Modelling and Comparing OECD Countries' Consumer Behaviour

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ABSTRACT

This thesis seeks to model and compare OECD countries' consumer behaviour. We build REPIH/RELCH and ECM formulations using long time-series based solely upon private sector measures of income for twenty OECD countries. No previous study features such a broad coverage of private sector data and models.

Using the Johansen procedure we build structural ECMs based upon consumption, income and inflation allowing for heterogeneous dynamics across countries and considering whether an intercept should be included in, or excluded from, the cointegrating vector. Models embodying asymmetric nonlinear adjustment towards equilibrium are also developed. We are not aware of any previous study which considers such flexibility of specification for twenty OECD economies. We build ECMs consistent with valid error-correction behaviour for eighteen countries and find evidence favouring asymmetric/nonlinear adjustment for twelve countries.

We derive a REPIH/RELCH model in logarithmic form to allow for current income consumers, durable expenditures and intertemporal substitution. We are aware of no previous study which simultaneously allows for all three of these features in a REPIH/RELCH model. This model is estimated with both GMM and IV methods. A proportion of current income consumers is found for all twenty countries and, in addition, accommodation for durability is evident for two economies. There is no evidence of intertemporal substitution.

Regressions are employed to explain the cross-country variations in the models' estimated parameters. We are not aware of any previous study attempting to explain variations in estimated elasticities from an ECM. We are able to explain the cross-country variations in the long run income and inflation elasticities but not the short run income and inflation elasticities or the adjustment coefficient. Only one previous study considers whether the cross-country variation in the estimated proportion of current income consumers can be explained by liquidity constraints. We use a broader range of proxies for liquidity constraints and additionally consider income uncertainty as a potential explanation. Application of iterative NL3SLS to the whole panel reveals that both factors explain the cross-country variation in this proportion.

CHAPTER 1

INTRODUCTION

The study of consumers' behaviour is of interest to orthodox economists because consumption is often suggested to be the ultimate end of all economic activity. Consumption is also the major constituent of aggregate demand, thus explaining its determination is highly important for understanding economic fluctuations and economic growth. Indeed, how it is modelled, in aggregate, may have major policy implications. For example, policy decisions which are heavily influenced by forecasts of consumption may be severely impaired if such predictions are highly inaccurate - Church *et al* (1994) note that UK consumption was severely underpricted in the late 1980s and overpredicted in the early 1990s. Berg (1994) highlights the importance of understanding the impact of economic policy on consumer behaviour by noting that the poor timing of financial deregulation and tax reform had severe adverse economic and social consequences in the Nordic countries. Empirical comparison of the determinants of different countries' consumer behaviour should enhance the understanding of expenditure decisions and policy effects upon them.

Four main theories have been forwarded as explanations of the essential aspects of consumer behaviour: Keynes's (1936) Absolute Income Hypothesis (AIH), Duesenberry's (1949) Relative Income Hypothesis (RIH), Modigliani and Brumberg's (1954) Life-Cycle Hypothesis (LCH) and Friedman's (1957) Permanent Income Hypothesis (PIH). Most contemporary analyses of consumption are based upon versions of the LCH or PIH frameworks. Both of these theories emphasise the dependence of consumption upon lifetime resources, introducing a prominent role for wealth and income expectations, and argue against Keynes's view that current income is the primary determinant of consumption.

In 1978 two seminal papers were published which established the framework for the majority of research into consumer behaviour to the present day. The first paper was by Hall (1978), who developed a means of testing an implication of the joint hypotheses of rational expectations and the PIH/LCH (REPIH/RELCH). Initial research focussed upon specifying the model to satisfy underlying statistical assumptions and to use the appropriate measure of consumption to enable the valid conduct of such a test. This literature established the joint hypotheses rejection. Subsequent analysis has sought to explain this rejection by relaxing the underlying assumptions of Hall's (1978) model. Three important studies of aggregate behaviour consider modifying this model to allow for variable interest rates (intertemporal substitution), Hall (1988), imperfect capital markets (current income consumers), Campbell and Mankiw (1991), and durable consumption expenditures, Caballero (1994). The evidence strongly suggests a role for a significant proportion of current income consumers. Although initially motivated by a belief in the potential prevalence of binding liquidity constraints, Carroll *et al* (1994) offer an alternative explanation for the presence of current income consumers, being uncertainty over (future) income (inducing precautionary saving). The majority of evidence appears to refute intertemporal substitution as an explanation for the rejection of the REPIH/RELCH. However, it has been argued that evident intertemporal substitution is found with the correct treatment of durability in the measure of consumer expenditure.

The second seminal paper was by Davidson, Hendry, Srba and Yeo (1978) - hereafter DHSY - which popularised the use of autoregressive distributed lag (ADL) error correction models (ECM). The DHSY model was shown to encompass typical empirical formulations of the AIH, RIH, PIH and LCH, to highlight the need to accommodate inflation effects in the wake of the 1973 OPEC oil shock and to emphasise the importance of lagged level terms to capture long run behaviour. Hendry and Ungern-Sternberg (1981) - hereafter HUS - encompassed the DHSY formulation, with the modelling of wealth effects providing the most notable extension. The direct modelling of (liquid) asset effects reduced the role of inflation to, at most, adjusting the conventional measure of disposable income for inflation induced losses on assets. The financial deregulation of the mid/late 1980s and reregulation of the late 1980s and early 1990s that occurred in many industrial countries led to substantial changes in wealth and the spendability of assets, especially illiquid assets such as housing. Broad wealth became the favoured asset measure to be incorporated into ECMs of consumer behaviour - see Patterson (1984) and Molana (1989). Indeed, when appropriately broad measures of wealth were incorporated into consumption functions, inflation appeared to have virtually no explanatory role, see Church, Mitchell, Smith and Wallis (1994) and Lattimore (1994). This suggests that, as Bean (1978) argued, inflation's role in consumption functions is to proxy for the erosion of nominally denominated assets when wealth effects are omitted from the model. The importance of wealth has been confirmed using the cointegration ECM methods of Engle and Granger (1987) and Johansen (1988), which facilitate the direct testing for long run equilibrium relationships, by, for examples, Drobny and Hall (1989) and Hall and Patterson (1991). Hence, ECMs based upon the fundamental variables of the LCH, income and wealth, provide the basis for the majority of structural modelling of consumer behaviour - the inclusion of current income can reflect naive income expectation formation, see Ando and Modigliani (1963), or the prevalence of liquidity constraints, see Miles (1992).

The majority of empirical analyses of the dominant REPIH/RELCH and ECM methodologies have focussed, primarily, on the UK and the USA, suggesting a need for similar analyses for a broader set of countries. A few recent analyses have applied these methods to a broader range of countries. These include Jin (1994), who applies a REPIH/RELCH formulation, modified for a proportion of current income consumers, to nineteen Organisation of Economic Cooperation and Development (OECD) economies over the period 1960-1988. Similarly, Carruth, Gibson and Tsakalotos (1996) examine the DHSY model for fifteen European Union (EU) countries over the period 1956-1990, while Pesaran, Shin and Smith (1997) consider the DHSY model using an ADL framework for twenty-four OECD economies over the period 1962-1992/3. However, none of these studies use the appropriate private sector measure of income, all use income measures which embody some component of public income. Thus, investigations of the REPIH/RELCH and ECM methodologies for a broad range of countries using measures of private income would provide improved inference.

The contribution to the literature provided by this thesis is the construction and comparison of REPIH/RELCH and ECM models using private disposable income for twenty OECD countries, which is available over the period 1955-1994. This analysis uses the broadest coverage of data based upon private sector measures of consumption and income relative to any previous study.

The models employed in our analysis utilise greater flexibility of specification compared to previous studies. For the ECM we utilise the Johansen (1988) cointegration method to test for long run relationships between consumption, income and inflation, where inflation is used to proxy asset effects (reliable wealth data of reasonable coverage is not available for the majority of OECD countries). We allow for heterogeneous dynamic specification across countries and for the intercept to be omitted or included in the long run consumption function. Structural error correction models, allowing for the endogeneity of contemporaneous variables, are then constructed using our favoured equilibrium consumption functions. We then consider ECM specifications allowing for asymmetric/nonlinear adjustment towards equilibrium. Such adjustment may be particularly appropriate for our measure of total consumption (which contains durables) because there may be a threshold which defines when disequilibrium changes from being tolerable to intolerable, suggesting that the speed of adjustment is related to the degree of disequilibrium.

The REPIH/RELCH model that we derive explicitly allows for current income consumers, durability and intertemporal substitution in a single, logarithmic, specification. Variants of this model is estimated for each country using Generaralised Methods of Moments (GMM), Instrumental Variables with Moving Average errors (IV-MA) and Autoregressive Integrated Moving Average (ARIMA) methodologies. No previous analysis has explicitly modelled all three of these features in a single formulation or used such a range of estimation methods. We note that the use of total expenditures to measure consumption is particularly appropriate for formulations, such as ours, which assume current income consumers spend all their income each period. This is because it means constrained consumers are not assumed to spend all their income on nondurables, as in many previous applications of modified REPIH/RELCH models, but can purchase durables as well.

The final contribution to the literature provided by this thesis is the systematic analysis of the cross-country variation in the estimated parameters of the ECM and REPIH/RELCH models, employing a wide range of country-specific policy, structural and other economic factors. Using cross-section regressions we investigate whether the countries' estimated short and long run income elasticities vary with factors hypothesised to explain the proportion of income that agents consume. Similarly, we assess whether the estimated short and long run inflation elasticities feature cross country variations consistent with factors postulated to explain the proportion of consumers' income held in assets - providing some insight into whether inflation acts as a proxy for wealth. We also examine whether the speed of adjustment to equilibrium varies with country specific institutional factors, that is, whether the estimated adjustment coefficient is correlated with variables that proxy the financial development of an economy. We are not aware of any previous analysis which investigates the cross country variation in the estimated parameters of ECMs. Finally, we investigate if the estimated proportions of current income consumers (obtained with both GMM and IV-MA methods) vary systematically with factors which proxy both liquidity constraints and future income uncertainty. We are not aware of any previous studies which consider the potential role for future income uncertainty in explaining such cross country variation. This cross-country variation is also examined with a REPIH/RELCH model applied to the complete time series/cross section panel which is estimated using iterative nonlinear three stage least squares (NL3SLS).

The thesis is organised as follows. Chapter two provides a review of the theoretical and empirical literature on consumer behaviour. The AIH, RIH, PIH and LCH theories of consumer behaviour are outlined and much of the empirical work on these early theories, especially the AIH, is argued to be invalid because the statistical models were often misspecified. It is also argued that the DHSY ECM encompasses the majority of empirical specifications of these theories. We find that two methods of analysing consumer behaviour have dominated the literature since 1978, being the ECM and REPIH/RELCH approaches. The ADL and cointegration ECM methodologies have highlighted the role of lagged level variables as providing important long run information. These investigations have also demonstrated the need to include broadly defined asset variables in consumption functions. Issues surrounding the role of (financial) deregulation, future income uncertainty, demographic factors, interest rates and income distribution are also discussed within this context. The REPIH/RELCH literature discusses the appropriate specification of the model to ensure one does not infer spurious excess sensitivity of current consumption to predictable income and the potential explanations for the rejection of the basic REPIH/RELCH. The literature on Campbell's (1987) implication of the REPIH/RELCH that saving anticipates declines in future labour income is also reviewed. Discussion on the evidence regarding REPIH/RELCH models modified to allow for current income consumers, variable interest rates, durability of expenditures and finite planning horizons (see Blanchard (1985)), is provided, and issues regarding the appropriate dating of instruments for expectational variables outlined. Finally, the emerging literature on solved out rational expectations consumption functions (see Muellbauer (1994)), is considered.

Chapter three conducts an analysis of the consumption, income and inflation data. The sources, construction and coverage of these series are given. Through visual inspection of data plots, basic descriptive statistics and augmented Dickey Fuller (ADF) tests we draw initial inferences regarding the characteristics of the data. The issues we address are whether consumption appears to be smooth relative to income, the degree to which the data are subject to outliers and structural breaks, whether consumption and income move/break together, why GDP is not a good proxy for income in studies of consumer behaviour and to gauge the series' orders of integration.

Chapter four tests for cointegration, using Johansen's (1988) multivariate cointegration framework between consumption, disposable income and inflation. As noted above, this specification may be regarded as an approximation of Ando and Modigliani's (1963) LCH with inflation capturing various wealth effects. Although there are other potential influences of long run consumer behaviour we concentrate on these three main variables for three reasons. First, because there is a lack of reliable long consistent time series on such factors, especially wealth, age structure and liquidity constraints. Second, whilst allowing for a certain degree of heterogeneity in model specification we also wish to consider similar models to facilitate meaningful cross-country comparison. This could be hindered if we considered a wide range of variables for each economy, especially if certain factors were retained in some countries' models and not others. Third, to minimise the size of the Vector Autoregression (VAR) given that inference becomes unreliable as degrees of freedom become scarce. Thus, we estimate equilibrium consumption functions using the Johansen VAR which capture the three main factors of consumer behaviour (with inflation acting as a wealth proxy). Preferred long run consumption functions are chosen for each country based upon tests of statistical significance on the cointegrating vectors' parameters and the adjustment coefficients as well as their theoretical plausibility. The postulate that consumption is homogeneous of degree one in income is tested and the issue of weak exogeneity is investigated. When the statistical evidence suggests two cointegrating vectors we consider the overidentifying restriction that consumption and income form one cointegrating relation and inflation a second, separate, stationary vector.

In Chapter five dynamic error correction models are developed for each country based upon the favoured long run consumption functions chosen in Chapter four. These are estimated using the method of Instrumental Variables (IV) when contemporaneous variables feature in the model. We consider the development of these models to allow for two broad forms of asymmetric adjustment to equilibrium - the Granger and Lee (1989) partitioned specification and the cubic nonlinear form employed by, for example, Hendry and Ericsson (1991).

A logarithmic REPIH/RELCH model modified to explicitly accommodate the presence of current income consumers, durable expenditures and intertemporal substitution is derived in Chapter six. This specification, along with formulations nested within it, are estimated by GMM, IV-MA and ARIMA methods for all twenty countries to determine the main causes of the rejection of the REPIH/RELCH. A review of previous literature on similar REPIH/RELCH models is provided and our results are compared with those of previous studies.

Chapter seven seeks to explain the cross-country differences in the parameters of the models estimated in Chapters four, five and six. That is, we use cross-section regressions to determine whether the long and short run income and inflation elasticities, the adjustment coefficient and the proportion of current income consumers systematically vary with postulated explanatory factors. We also employ a panel estimation technique to assess the factors explaining the cross-country variation in the proportion of current income consumers. A discussion of the relevant country-specific factors (based upon theoretical and empirical considerations) which potentially explain cross-country differences in consumer behaviour is also provided.

Chapter eight summarises the findings of the thesis and discusses their implications for policy and empirical modelling of consumer behaviour.

CHAPTER 2

REVIEW OF THE THEORETICAL AND EMPIRICAL LITERATURE ON THE CONSUMPTION FUNCTION

2.1 Introduction

This Chapter provides a review of the theoretical and empirical literature on the consumption function. The aim is to provide a background for the present thesis and motivate the research and methods employed herein.

Section 2.2 outlines Traditional theories of consumer behaviour being Keynes's absolute income hypothesis (AIH), the relative income hypothesis (RIH) of Duesenberry (1949) and Brown (1952), Friedman's (1957) permanent income hypothesis (PIH) and the life cycle hypothesis (LCH) associated with Ando, Brumberg and Modigliani (in various papers). Section 2.3 considers Hall's (1978) seminal work on the rational expectations permanent income/life cycle hypotheses (REPIH/RELCH) with the subsequent extensions for durable goods, Mankiw (1982) and Caballero (1994), variable interest rates, Hansen and Singleton (1983) and Hall (1988), current income consumers, Hall and Mishkin (1982) and Campbell and Mankiw (1991), and uncertainty over life length/planning horizon, Blanchard (1985) and Evans (1988). Section 2.4 reviews the autoregressive distributed lag (ADL) error correction models (ECM) pioneered by Davidson *et al* (1978) and extended by Hendry and Ungern-Sternberg (1981) and the extensions explored, primarily within this framework, to explain the breakdown of these consumption functions during the mid/late 1980s. The research on consumption in the cointegration ECM frameworks of Engle and Granger (1987) and Johansen (1988) are discussed in Section 2.5. Section 2.6 reviews the solved out rational expectations consumption functions associated with Muellbauer and recent evidence on the LCH. Section 2.7 provides conclusions.

2.2 Traditional Theories of Consumer Behaviour

In this section we outline the AIH, RIH, PIH and LCH.

2.2.1 The Absolute Income Hypothesis

Keynes (1936) argued that contemporaneous income (Y_t) was the primary determinant of consumption (C_t) and formed the "fairly stable relationship":

$$C_t = a_1 + b_1 Y_t. \qquad \textbf{(2.2.1)}$$

Keynes suggests that consumers will spend some proportion of their income motivating the "fundamental psychological law" that the marginal propensity to consume (MPC) will fall between zero and one ($0<b_1<1$). The average propensity to consume (APC) is suggested to exceed the MPC because autonomous consumption is positive ($a_1>0$): subsistence consumption.[1] Another proposition attributed to the Keynesian consumption function is that the MPC falls as the level of income rises. However, this last proposition "seemed to be an aside remark which was neither pressed hard nor included under the fundamental psychological law. Keynes does not state in any explicit way that there is a secular trend of the propensity to consume to decline with income." (Hadjimatheou, 1987, p. 2).

Davis (1952) reports the results of several early time series applications of **(2.2.1)** to aggregate US consumption. The evidence typically supported the proposition of positive autonomous consumption (APC>MPC) and that the MPC falls between zero and one. However, such models estimated using pre-World War II data systematically underpredicted post-War consumption, casting doubts over the proposition of a "fairly stable relationship". Kuznets (1946) and Goldsmith (1955) find that there is no autonomous consumption (constant APC) when using an averaged data set over a long period of time and conclude that *long run* data contradicts the AIH.

[1] Keynes recognised that windfall changes in non-human wealth, large fluctuations in the interest rate and alterations in income distribution could cause the MPC to vary, however, these factors were regarded as secondary to the "fundamental psychological law".

While Brady and Friedman (1947), using cross sectional budget data for various years, found that although autonomous consumption was positive, it increases through time. New theories were elaborated in an attempt to explain these apparent inconsistencies of the AIH with the data.[2]

2.2.2 The Relative Income Hypothesis

Duesenberry's (1949) RIH postulates that an individual's APC depends upon their percentile position in the overall income distribution. Those with relatively low living standards are portrayed as attempting to emulate the consumption patterns of the better off implying that low income earners exhibit larger APCs than high income earners. The RIH could explain why cross sectional studies indicated a Keynesian style consumption function with a positive intercept: the less well off maintain relatively high consumption levels despite their relatively low incomes. It is also consistent with Kuznets's (1946) finding of a constant *aggregate* APC: if all consumers' incomes increased by the same proportion then, because the income distribution remains unchanged, individuals', and therefore the *aggregate*, APC(s) would be constant. This theory could also explain the apparent upward shifting of the short-run Keynesian consumption function about a long term constant APC. Through habit and a desire to maintain their standards of living consumers would be reluctant to allow their consumption to fall, even with declining income. Therefore, consumption would depend on both the level of income and its size *relative* to the previous highest value of income and/or consumption.

Brown (1952) developed the RIH into the habit persistence model, arguing that last period's consumption, rather than its previous peak, determines current consumption, due to the persistence of habit, thus:

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[2] Spanos (1987) finds that application of statistical procedures available in the 1980s show that these early empirical findings are based upon severely misspecified consumption functions, so provide invalid inferences. This was especially true of the time-series studies where the statistical models employed did not account for autocorrelation and spurious/nonsense regression arising from the strong temporal correlations that typically prevail in such time-series data. Hence, little of the early evidence on the AIH was informative. Indeed, even now there is difficulty in turning **(2.2.1)** into a valid statistical model for testing Keynes's theory - certainly for time series analysis.

$$C_t = c_1Y_t + d_1C_{t-1}. \quad \textbf{(2.2.2)}$$

Lagged consumption replaces the intercept in a time-series regression of **(2.2.1)** to capture the temporally increasing constant uncovered by Brady and Friedman (1947). Assuming that consumption grows at a constant rate, g, such that $C_t = (1+g)C_{t-1}$ then, through substitution into **(2.2.2)**, and solving for C_t, we obtain the long-run consumption function:

$$C_t = \{c_1/(1-[d_1/(1+g)])\}Y_t. \quad \textbf{(2.2.3)}$$

(2.2.3) suggests the APC is constant in the long-run, which is consistent with the evidence presented by Kuznets (1946). The habit persistence form of the RIH appeared to be more data coherent than the AIH. Spanos (1987) suggests that this is due to the inclusion of lagged consumption. Indeed, Davis (1952) noted that the addition of lagged variables substantially improved the predictive performance of the Keynesian consumption function. The RIH, PIH and LCH may have been developed with this in mind.

2.2.3 The Permanent Income Hypothesis

Unlike the AIH and RIH, Friedman's (1957) PIH relates *actual* consumption, not expenditure, to *permanent*, rather than current, income. *Permanent* income is defined as the per-period level of household consumption that can be secured *for eternity* whilst maintaining the household's stock of *real* wealth from period to period. Consumers are, therefore, forward looking, planning their consumption over an *infinite* time horizon with an infinite planning horizon suggesting that consumers are as concerned about their heirs' utility as they are about their own. The simple PIH expresses *permanent* consumption, C_t^P, as some proportion, k, of *permanent* income, Y_t^P:

$$C_t^P = kY_t^P. \quad \textbf{(2.2.4)}$$

In **(2.2.4)**, k=1 because wealth only remains unchanged if all of a household's *permanent* income is consumed each period. Thomas (1994) notes that empirical estimates of k are typically below one, implying ever increasing wealth. To allow for this, Friedman modifies **(2.2.4)** to characterise households as accruing wealth as a precaution against unforseen circumstance. That is, k is specified to positively depend upon the ratio of non-human wealth to total wealth, w. k is also suggested to be related to the rate of return on wealth, r, and tastes, u, yielding the modified PIH:

$$C_t^P = k(r, w, u)Y_t^P. \qquad \textbf{(2.2.5)}$$

Further, the proportionality between C_t^P and Y_t^P, is an inessential part of the PIH. When proportionality is relaxed, by adding an intercept, permanent consumption remains a function of permanent, and not current, income - changes in current income only affect consumption to the extent that they cause a revision in Y_t^P.

If the PIH is *true* one can explain the early empirical findings of the AIH with reference to the measurement of income and consumption. Friedman postulates that both *measured* consumption and *measured* income are the sums of *permanent* (P) and *transitory* (TR) components, thus:

$$C_t = C_t^P + C_t^{TR} \qquad \text{and,} \qquad Y_t = Y_t^P + Y_t^{TR}. \qquad \textbf{(2.2.6)}$$

Transitory income is not expected to be permanent so does not influence *permanent* consumption.[3] Substituting **(2.2.6)** into **(2.2.4)** transforms the PIH into a relationship of observed variables, thus:

$$C_t = kY_t + (C_t^{TR} - kY_t^{TR}). \qquad \textbf{(2.2.7)}$$

If the PIH representation **(2.2.7)** is *true*, estimation of the AIH equation **(2.2.1)** by OLS would lead to a downward biased estimate of the slope coefficient, k, and an upward biased intercept.[4]

[3] To facilitate the testing of the PIH with observed data Friedman assumes that the transitory and permanent components of consumption and income are uncorrelated and that the transitory components of consumption and income are uncorrelated.

[4] This is because the error term, $(C_t^{TR} - kY_t^{TR})$, is negatively correlated with the explanatory variable, $Y_t = Y_t^P + Y_t^{TR}$, assuming Y_t^{TR} is non-negligible.

Thomas (1994) argues that with a long span of aggregate time series data, such as Kuznet's (1946), the transitory cycle will represent minor variations so, if the PIH is *true*, estimation of **(2.2.1)** will be subject to little bias and predicts the zero intercept observed by Kuznets. In contrast, estimating **(2.2.1)** with shorter time series, such as those reported by Davis (1952), will be subject to greater bias because the transitory (cyclical) components will represent a major portion the data's variation. This explains the positive intercepts reported by Davis and the smaller estimated slope coefficients compared to Kuznets. Thus, the PIH can rationalise the time-series results of the AIH. The PIH also predicts the cross section findings of Brady and Friedman (1947): the positive intercept and smaller slope relative to time series studies. This is because the transitory component of household income will likely represent a larger portion of the data's variation relative to *aggregate* income because individuals' transitory incomes will offset each other in aggregate series.

Friedman (1957) presents numerous cross section tests of the PIH, however, "many of his predictions can also be derived, if with a little more difficulty, from the RIH." (Thomas 1994, p. 262). One of Friedman's tests that does provide support for the PIH and not the RIH (or the AIH) is that the use of cross section data with relatively small variations in *transitory* income yield MPCs and income-elasticities which are larger than with samples containing greater variability in the transitory component.

Friedman's (1957) direct estimates of the PIH suggested little correlation between consumption and current income (0.33) but a high correlation between consumption and permanent income (0.88). However, Wright (1969), for example, re-ran Friedman's regressions with "minor" modifications for the war years and obtained estimates of the correlation between consumption and current income which were as high as 0.8, which is more consistent with the AIH than the PIH. Indirect time-series estimates of the PIH can be obtained using the following model:[5]

$$C_t = kqY_t + (1-q)C_{t-1} + u_t \qquad \textbf{(2.2.8)}$$

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[5] Using the adaptive expectations scheme, $Y^P_t - Y^P_{t-1} = q(Y_t - Y^P_{t-1})$, to eliminate Y^P_t from **(2.2.4)**, applying a Koyck transformation to the result, and using **(2.2.6)** to eliminate the unobserved transitory components one obtains the PIH model, **(2.2.8)**.

where, $u_t = C^{TR}_t - (1-q)C^{TR}_{t-1} + e_t - (1-q)e_{t-1}$, which is a first-order moving average process [MA(1)]. However, application of OLS to **(2.2.8)** will yield biased and inconsistent estimates, given this autocorrelation and the presence of a lagged dependent variable. Further, this statistical formulation of the PIH is the same as the habit persistence form of the RIH, excepting the error process so does not discriminate between the two hypotheses. Given the statistical problems of applying OLS to **(2.2.8)** and the controversial nature of Friedman's time-series results, the time-series evidence on the PIH seems uninformative. However, the cross-section evidence offers some support for the superiority of the PIH over both the AIH and the RIH.

2.2.4 The Life Cycle Hypothesis

The LCH associated with Ando, Brumberg and Modigliani (see, for example, Ando and Modigliani 1963) also attempts to explain the early empirical findings against the AIH. In the LCH the individual household maximises utility, U^T, subject to the budget constraint of total life-time resources:

$$\text{Max:} \quad U^T(C_t, E_tC_{t+1}, \ldots, E_tC_{t+L-T}) \qquad \textbf{(2.2.9)}$$

$$\text{s.t.} \sum_{i=t}^{t+L-T} E_tC^T_i\,(1+r)^{t-i} = A^T_t + \sum_{i=t}^{t+L-T} E_tY^T_i\,(1+r)^{t-i} \qquad \textbf{(2.2.10)}$$

where A^T_t is the household's net worth at the *start* of period t, $E_tC^T_i$ is actual consumption expected in period t for period i, and $E_tY^T_i$ is labour income expected in period t for period i. r is the (assumed) constant real interest rate, L is the lifespan of the household and T is the age of the household in period t. We assume the following homothetic utility function for the household:[6]

$$U^T = \sum_{i=t}^{t+L-T} U(E_tC^T_i)(1+\delta)^{t-i} \qquad \textbf{(2.2.11)}$$

[6] Homotheticity ensures that the *relative* expenditure shares *across* periods are independent of the size of *expected* life-time resources.

where δ is the rate of time preference (discount rate).[7] We further assume constant elasticity of substitution (CES) preferences, where the elasticity of substitution is $\sigma = 1/\beta$:[8]

$$U(E_tC^T_i) = [1/(1-\beta)](E_tC^T_i)^{1-\beta}. \qquad \textbf{(2.2.12)}$$

Solution of the model yields the *Euler* equation:

$$E_tC^T_{t+1} = [(1+r)/(1+\delta)]^{1/\beta}\, C^T_t. \qquad \textbf{(2.2.13)}$$

By analogy to **(2.2.13)** we can obtain an expression for every expected value of consumption in the life cycle as a function of current consumption, thus:

$$E_tC^T_{t+i} = [(1+r)/(1+\delta)]^{i/\beta}\, C^T_t, \qquad i = 1,\ldots, L\text{-}T. \qquad \textbf{(2.2.14)}$$

Summing **(2.2.14)** for all remaining periods of the household's life, from period t to t+L-T, and substitution into the budget constraint,[9] gives the optimal consumption function for a household of age T:[10]

$$C^T_t = \{A^T_t + \sum_{i=t}^{t+L-T} E_tY^T_i\,(1+r)^{t-i}\} \,/\, \{\sum_{i=t}^{t+L-T} [(1+r)/(1+\delta)]^{(i-t)/\beta}\,(1+r)^{t-i}\} \qquad \textbf{(2.2.15)}$$

The aggregate consumption function can be obtained by first summing **(2.2.15)** over the M households of age T and then summing the result for the N different aged households:

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[7] The larger is δ the greater is the preference for current rather than future consumption.

[8] Thus, marginal utility is: $\partial U/\partial C^T_t = (C^T_t)^{-\beta}$, and the elasticity of marginal utility with respect to consumption is: $\{\partial[\partial U/\partial C^T_t]/\partial(C^T_t)\}.\{C^T_t/(\partial U/\partial C^T_t)\} = \{-\beta(C^T_t)^{-\beta-1}\}\{C^T_t/[(C^T_t)^{-\beta}]\} = -\beta$. For $\beta>0$ marginal utility decreases as consumption increases (the utility function is strictly concave). β is sometimes referred to as the coefficient of relative risk aversion.

[9] This is equivalent to substituting into the derivative of the Lagrangian function with respect to the Lagrange multiplier (marginal utility of expected consumption) for period i.

[10] One substitutes the indifference curve, equating marginal utilities, into the budget constraint, to ensure they are tangential.

$$C_t = b_2Y_t + b_3E_tY + b_4A_{t-1} \tag{2.2.16}$$

where C_t is aggregate consumption and Y_t is *aggregate* (labour) income and A_{t-1} is the *aggregate* stock of *end of period* assets.[11]

This model is based upon the following assumptions: (1) No bequests are left or received; (2) The rate of return on assets is, and is *expected* to be, constant; (3) Capital markets are perfect: there are no liquidity constraints; (4) Each household exhibits the same utility function and uses the same discount rate; (5) The economy's age distribution, age distribution of income and age distribution of net worth, are constant (with respect to time); (6) Households make a detailed consumption plan for every period of their life; (7) Variations in the degree of uncertainty over future *expected* levels of income do not affect the allocation of consumption through time; (8) The household's planning horizon is its whole life-span (L), which is known with certainty; (9) The rate of time-preference is constant; (10) Households act according to the optimal consumption plan, as identified above.

Modigliani (1986) characterises the LCH as households attempting to maintain a constant level of consumption throughout their entire life in the face of an expected hump-shaped-income-profile by saving and borrowing. Typically, young consumers command below their average expected income so seek to borrow to attain their desired consumption level. As income rises during mid-life through to retirement the consumer is able to repay debts and accumulate savings sufficient to sustain their constant consumption level throughout retirement. Hadjimatheou (1987) identifies six major implications of this model as: (1) A country's saving rate is independent of its per-capita income; (2) Differences in nations' savings rates are not inconsistent with identical individual household's life-cycle behaviour; (3) For those countries with identical life-cycle behaviour, those featuring higher long-run growth rates will exhibit larger saving rates; (4) The wealth-income ratio is negatively related to an economy's growth rate; (5) An economy can accumulate substantial wealth relative to its income without bequests; (6) The prevailing

[11] The apparent difference in the asset variables' time subscripts is because A_{t-1} refers to end of period holdings while A^T_t to beginning of period stocks.

length of retirement is the major determinant of the wealth-income and saving ratios, for a *given* growth rate.

Muellbauer (1994) modifies this characterisation of the LCH to accommodate demographic effects. He argues that aggregate consumption will rise with an increase in the proportion of the population who are young and retired (lower savers). Further, older consumers will exhibit larger MPCs out of assets because they have a shorter remaining lifespan, compared to the young, in which to spend them. The difference between younger and older generation's MPCs out of assets will reduce in the presence of a bequest motive - the size of this difference will greatly affect the size of demographic, population and income growth effects on aggregate saving. Additionally, variation in the needs of children may affect aggregate consumption. Young couples with children expect increased expenses, and possibly reduced income, during the middle of their life compared to households without offspring. This may reduce the incentive of the young to borrow with the expected burden of mid-life debt repayments. Banks *et al* (1994) present simulation evidence based upon microeconomic UK data which suggests that households expecting more children reallocate expenditure *into* periods when those children are present while Deaton (1992) provides cross country evidence indicating highly volatile saving profiles throughout the life-cycle. Both sets of evidence are consistent with the needs of children influencing consumption.

Nevertheless, the innovation of the LCH is that it predicts that current consumption is determined by the wealth that a household *expects* to generate throughout its life: "in contrast to Keynes's approach, the life-cycle model in its pure form assumes that individual consumers are forward looking, planning over their life-span, with future changes in their economic circumstances having as much influence on their behaviour as their current situation." (Hadjimatheou 1987, p. 34). Similar to the PIH, consumption will only respond to current income to the extent that it affects expected life time resources: consumption will be smoother relative to the AIH. Indeed, both LCH and PIH suggest that a sustainable level of consumption will be determined by expectations about future resources. The only differences between the two hypotheses are that the horizon over which expectations are formed are the life cycle (LCH) and eternity (PIH); that the LCH explicitly specifies wealth as an explanatory factor whereas it is only implicit in the PIH; and that the response of consumption to changes in expected lifetime resources is lower in the PIH, due to a precautionary saving motive, compared to the LCH.

The length of the household's planning horizon has been subject to considerable debate in the literature. Barro's (1974) intergenerational altruism argument suggests that parents' utility is an increasing function of their children's utility such that the household may be characterised *as if* it were infinitely lived. Weil (1994) identifies two channels through which bequests can affect savings: by children influencing their parent's saving and parents influencing their children's saving. Weil (1994) finds evidence for the latter but not the former and argues that "This evidence supports the view that even if the old do not dissave themselves, they lower the saving of the young via bequests." (Weil 1994, p. 56). However, Carroll (1994) notes that the data used by Weil is comprised of 70% to 95% noise suggesting his results should be viewed with caution. Overall, Muellbauer (1994) argues that an infinite planning horizon may be unrealistically long because parents may apply a larger discount factor to their childrens' utility: children may already have assets and have better income prospects (in a growing economy). This is consistent with the general evidence against complete Ricardian equivalence (that current tax cuts will not raise current consumption because parents realise that their children will have to pay for them in the future).

To empirically implement the LCH one needs data on income expectations and wealth. Initial investigations employed simplifying assumptions regarding expectations, For example, Ando and Modigliani (1963) assume "naïve" income expectations:

$$E_tY = b_5Y_t. \qquad \textbf{(2.2.17)}$$

Substituting **(2.2.17)** into **(2.2.16)** yields:

$$C_t = b_6Y_t + b_4A_{t-1}, \qquad \text{where, } b_6 = b_2 + b_5. \qquad \textbf{(2.2.18)}$$

The appropriate concept of wealth is also an issue. Do total money balances or just outside money (net private balances) stimulate consumption (when prices fall in a recession)? See Pigou (1943) and Kalecki (1944). Should interest-bearing public debt be included in private wealth or is it discounted by *ultra-rational* consumers who realise that future taxes must rise to pay off this government debt? See Barro's (1974) debt neutrality hypothesis (Ricardian equivalence). Are liquid assets sufficient because they comprise the major variation in wealth or are financial markets sufficiently liberalised to facilitate significant illiquid asset expenditures making total wealth the appropriate measure? See Patterson (1984). The problem in defining wealth is compounded by the general lack of reliable data of sufficient coverage for a reasonable concept of wealth. Beyond the UK and the USA we are only aware of a broad measure of wealth being available for Australia (Lattimore, 1994) and Japan (Horioka, 1996). Consequently, indirect means of accounting for wealth have often been employed, particularly in early studies.

Two widely employed indirect ways of accommodating wealth are its elimination from the mathematical formulation and the use of proxy variables. Assuming no capital gains/losses (capital gains) on assets, we can use the following definition, $A_t = A_{t-1} + Y_t - C_t$, to eliminate wealth from **(2.2.18)**, to yield:

$$C_t = b_6 Y_t + (b_4 - b_6) Y_{t-1} + (1 - b_4) C_{t-1}. \tag{2.2.19}$$

Data constraints led to the use of **(2.2.19)** until its performance waned with the 1970s oil shocks. Another problem with this wealth elimination specification is its similarity to RIH and PIH empirical formulations, making it difficult to assess the adequacy of one theory over another. These factors motivated the use of proxies for wealth. The most widely used proxy is inflation which is negatively correlated with wealth because rising inflation reduces the real value of nominally fixed assets.[12] For example, inflation causes a capital loss on (outside) money and other financial and illiquid assets - see Grice (1981), Patterson (1984 and 1985) and Hendry (1994).[13] For the components of wealth where data is available one may combine the inflation effect with that asset. However, data on the full spectrum of assets is generally unavailable leaving a role for inflation as a proxy for omitted wealth effects.

[12] Inflation may also have substantial distributional wealth effects.

[13] Beyond monetary assets one should really adjust any capital gain/loss with the asset's inflation rate as well - see Pesaran and Evans (1984).

There have been non-wealth justifications for the inclusion of inflation in consumption functions, the most popular being Deaton's (1977) hypothesis of a negative relationship between consumption and *unanticipated* inflation.[14] Deaton postulated that consumers who do not need to know the prices of goods purchased at infrequent and irregular intervals cannot distinguish between absolute and relative price increases. Assuming inflationary expectations are determined by recent experience, following an adaptive expectations scheme, then, during periods of accelerating inflation, expectations will lag behind actual price increases. Therefore, consumers may mistake *unanticipated* price rises as relative rather than absolute increases, making all goods appear relatively more expensive, causing a substitution away from the consumption of all goods. Deaton (1977) finds support for his hypothesis in the sense that inflation is positively associated with the average propensity to save in the UK and US.

However, Hadjimatheou (1987) argues that assuming expectations are adaptive, rather than rational, characterises consumers as systematically failing to learn from their experience of accelerating inflation, which is probably unrealistic. MacDonald and Peel (1985) find, for Canada, France, Germany, Japan, the UK and the USA that Deaton's (1977) model is rejected if the rational expectations hypothesis is valid. While Blinder and Deaton (1985) conclude, using US data, that *anticipated* rather than *unanticipated* inflation affects consumption. Since Hadjimatheou (1987) suggests that inflation during the 1970s was probably unanticipated other interpretations for the statistical significance of inflation effects may be required. Bean (1978) argues that the negative association and statistical significance of inflation in consumption functions is better interpreted according to the classical wealth effect because of the potential severity of the erosion of the *real* value of nominally fixed assets. He presents empirical support for this hypothesis.

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[14] Examples of anticipated inflation influencing consumption include the following. Juster and Wathchel (1972a, b) argue that higher rates of anticipated inflation are associated with a greater variability of inflation which induces precautionary saving through increased real income uncertainty. In contrast, Springer (1977) argues that increased *anticipated* inflation causes an intertemporal shifting of future consumption to the present. Similarly, Bulkley (1981) argues that in years of rising anticipated inflation consumption will increase due to nominal wage contracts being agreed at discrete intervals of time. However, Hadjimatheou (1987) suggests that inflation during the 1970s was probably unanticipated rather than anticipated.

Ando and Modigliani (1963) use Goldsmith's (1956) wealth series to directly estimate various versions of the basic LCH equation **(2.2.18)** and find that it exerts a statistically significant impact upon consumption. In contrast, Evans (1967) found no significant role for net worth using post-war US data while Duesenberry (1971) points out that equally well specified consumption functions have been obtained without wealth as with its inclusion. However, a body of subsequent evidence appears to support the importance of wealth influencing consumer behaviour. Mayer (1972) empirically tests the major theories of consumers' behaviour and finds support for the inclusion of wealth. Mishkin (1977) enters several disaggregated components of wealth with statistical significance in a model of US durable expenditures finding that the different components have different impacts and arguing that the 1973-1975 collapse of US consumption is due to changes in US wealth. Modigliani (1981) establishes support for asset effects in a US consumption function. Pesaran and Evans (1984) find that a savings version of Ando and Modigliani's (1963) model, equation **(2.2.18)**, extended to allow for inflation induced capital gains and losses on ordinary shares outperforms the models of Deaton (1977), Davidson *et al* (1978) [hereafter DHSY], Hall (1978), Hendry and Ungern-Sternberg [hereafter HUS] and Muellbauer (1983) for the UK. A clear role for wealth effects are demonstrated for Australia by Lattimore (1994), for the UK by Church *et al* (1994), for the US by Muellbauer and Lattimore (1995) and for Japan by Horioka (1996). Indeed, many of these studies find that when well defined asset effects are incorporated in their model inflation effects have no role. Overall, this provides support for the role of wealth and, therefore, the LCH. It also suggests that inflation captures the role of wealth effects in the absence of good asset data.

Virtually all analyses of consumer behaviour are now based upon the PIH-LCH.[15] However, it relies on some questionable assumptions which have been examined in the recent literature. The assumptions which have been most widely relaxed are the formulation of expectations, the lack of liquidity constraints, income distribution, income uncertainty and demographic effects, along with the constancy of (real) interest rates, the length of planning horizon, the presence of transitory consumption and the role of durability. The removal of these assumptions have

[15] In the literature the LCH and PIH are treated as essentially the same theory (PIH-LCH), where the main difference, the length of planning horizon, is an issue to be resolved.

primarily been considered in the two dominant methodologies of consumer behaviour since 1978: the rational expectations permanent income life cycle hypotheses (REPIH/RELCH) originated by Hall (1978) and the error correction mechanism pioneered by DHSY. We consider the removal of these assumptions within these two methodologies.

2.3 The REPIH/RELCH Methodology

This section reviews the theory and evidence on Hall's (1978) and Campbell's (1987) implications of the REPIH/RELCH with subsequent extensions to accommodate durabilitity, intertemporal substitution, current income consumers and uncertainty over life length.

2.3.1 Testing Implications of the REPIH/RELCH

Hall's (1978) REPIH/RELCH model is motivated by two factors. First, the Lucas (1976) critique argued that econometric specifications would suffer parameter instability if changing policy regimes were not anticipated by the model. The *backward looking* expectations schemes (naive and adaptive) typically employed in the empirical implementation of the PIH and LCH cannot anticipate changes in policy regimes, so will portray consumers' behaviour as identical under old and new regimes. However, the *forward looking* rational expectations scheme characterise agents as utilising all information available at the time of expectation formation. To the extent that consumers can predict future policy changes, they will be able to revise their expectations and, therefore, their consumption. Hall (1978) introduces rational expectations into the PIH-LCH framework to make the model more robust to changing policy regimes. The second motivation is to ensure all right hand side variables are exogenous to avoid simultaneity bias. Hall's (1978) model assumes quadratic preferences:

$$U(E_tC_i) = [-(\tfrac{1}{2})(C-E_tC_i)^2] \qquad \textbf{(2.3.1)}$$

where C is some *bliss* level of consumption. Substitution of **(2.3.1)** into **(2.2.11)**, and maximising the resulting objective function subject to the budget constraint **(2.2.10)**, gives the Euler equation:

$$C_t - C = \{[(1+r)/(1+\delta)]\ (E_tC_{t+1} - C)\}. \qquad \textbf{(2.3.2)}$$

Rearranging **(2.3.2)** to normalise on E_tC_{t+1} gives:

$$E_tC_{t+1} = [(1+\delta)/(1+r)]\ C_t + [(r-\delta)/(1+r)]\ C. \qquad \textbf{(2.3.3)}$$

The expected level of consumption in period t+1 only differs from its actual level if there is a *surprise* in expected income arising after the expectation was made. Assuming this income innovation, ϵ_{t+1}, is stochastic means we can define actual consumption in period t+1 as:

$$C_{t+1} = E_tC_{t+1} + \epsilon_{t+1} \qquad \textbf{(2.3.4)}$$

$$\text{or, } E_tC_{t+1} = C_{t+1} - \epsilon_{t+1}. \qquad \textbf{(2.3.5)}$$

Substituting **(2.3.5)** into **(2.3.3)** yields Hall's (1978) REPIH/RELCH *levels* equation, thus:

$$C_{t+1} = b_7 + b_8C_t + \epsilon_{t+1}, \qquad \textbf{(2.3.6)}$$

where, $b_7 = [(r-\delta)/(1+r)]C$ and $b_8 = [(1+\delta)/(1+r)]$. Setting $r=\delta$ yields the following *difference* form of Hall's (1978) model:

$$\Delta C_{t+1} = \epsilon_{t+1}. \qquad \textbf{(2.3.7)}$$

(2.3.6) and **(2.3.7)** form the basis of Hall's (1978) joint test of the REPIH/RELCH. It suggests that the best prediction of next period's consumption is this period's consumption. The implication is that if *both* rational expectations and PIH/LCH are true then all available information at time t (t-1) should be incorporated in C_t (C_{t-1}) and so information dated in period t (t-1) and earlier should have no explanatory power for consumption in period t+1 (t). If such regressors can enter with statistical significance then one or both of these hypotheses is inconsistent with the observed data.[16]

Hall (1978), using US data, finds that neither lagged consumption nor income can be added to **(2.3.6)** with statistical significance. The insignificance of lagged income indicates that the information embodied in Friedman's (1957) adaptive expectations scheme has already been accounted for by consumers: it is an inferior expectations formulation. However, lagged stock prices (approximating wealth) were statistically significant. Although the *pure* REPIH/RELCH is rejected, Hall (1978) argues that the evidence is consistent with a *modified* version, where consumption takes time to adjust to wealth induced changes in permanent income. Davidson and Hendry (1981) and Daly and Hadjimatheou (1981) present evidence to refute Hall's (1978) model using UK data. Cuddington (1982) rejects the REPIH/RELCH for Canada and Johnson for Australia. These papers demonstrate a general rejection of the REPIH/RELCH.[17]

Flavin (1981) derives a *structural* REPIH/RELCH model which includes current and lagged income. Current income is expected to be statistically significant because it contains new information which will cause permanent income to be revised.[18] Lagged information is, as before, expected to be statistically insignificant if the REPIH/RELCH is true. This is tested using a reduced form from which the structural parameters of interest can be retrieved. This involves adding lagged income to **(2.3.7)**, with intercept, and testing for the statistical significance of

[16] This model rests on the following assumptions: (1) no credit restrictions or other non-linearities in the budget constraint; (2) no habits or adjustment costs; (3) the subjective discount rate, δ, is the same across consumers; (4) there are no measurement errors or transitory shocks to consumption; (5) the frequency of consumer decisions coincides with the data's periodicity; (6) the real interest rate is constant. (7) expectations are formulated rationally; (8) consumers plan between two adjacent periods which are not just the present and the future; (9) the utility function's form is quadratic and additive over time (yielding marginal utility which is linear in consumption); (10) the representative agent model can be applied to aggregate data.

[17] Berloffa (1997) finds, using microeconomic data, that the hypothesis of no excess sensitivity cannot be rejected for households with two earners once heterogeneity of individuals in the sample is accounted for.

[18] Flavin (1981) recognises that contemporaneous variables other than income may cause revisions in expected income and suggests that these influences will enter the error term. She proceeds *as if* current income is the only relevant news regarding income expectations.

these lags. Using US data she rejects the REPIH/RELCH, finding consumption to be *excessively sensitive* to lagged income. Interestingly, Flavin (1981) applies this test after detrending her data to recognise that "the model is intended to explain revisions in planned consumption which are caused by changes in expectations about future income," and goes on to suggest that her model "applies to the movement of consumption around a trend attributable to the trend in per capita income." (Flavin 1981, p. 989). She further comments on the problems of spurious/nonsense regression by stating that, "If the income process does include a trend, Hall's tests are misspecified under the alternative hypothesis." (Flavin 1981, p. 1004).

Mankiw and Shapiro (1985) find that incorporating trends in the *levels* random walk test equation, or equivalently detrending variables as Flavin (1981) does, is inappropriate if they feature stochastic rather than deterministic trends because the test statistics will have non-standard limiting distributions. Specifically, the tests will be biased toward accepting excess sensitivity to predictable changes in income because fatter tailed distributions than standard (higher critical values) are appropriate. Aggregate consumption and income series are generally found to be difference rather than trend stationary, indeed Harvey (1997) points out that the assumption of (linear) trend stationarity is unduly simplistic. The majority of subsequent research into the REPIH/RELCH utilise the difference transformation to induce stationarity.

Muellbauer (1983) derives a more powerful test of the REPIH/RELCH utilising CES preferences and using the growth of consumption as the dependent variable. Substitution of **(2.3.5)** into the Euler equation, **(2.2.13)**, gives:

$$\Delta \ln C_{t+1} = b_9 + \epsilon_{t+1}. \qquad \textbf{(2.3.8)}$$

The *surprise* in expected income is defined as, $\epsilon_{t+1} = E_tY_{t+1} - E_{t-1}Y_t$, which when substituted into **(2.3.8)** yields Muellbauer's (1983) *surprise* consumption function, $\Delta \ln C_{t+1} = b_9 + (E_tY_{t+1} - E_{t-1}Y_t)$. Muellbauer (1983) operationalises this test equation by approximating the *surprise* in permanent income using the following distributed lag in income, $\ln Y_{t+1} = b_{10} + b_{11}\ln Y_t + b_{12}\ln Y_{t-1} + b_{13}\ln C_t + u_{t+1}$, and assuming proportionality between the income innovation and the *surprise* in expected income, $u_{t+1} = b_{14}(E_tY_{t+1} - E_{t-1}Y_t)$. The basic form of Muellbauer's (1983) REPIH/RELCH equation is:

$$\Delta \ln C_{t+1} = b_9 + b_{14}\hat{u}_{t+1} + v_{t+1}. \qquad \textbf{(2.3.9)}$$

Muellbauer (1983) finds that b_{14} is statistically insignificant in various forms of **(2.3.9)** applied to UK data, appearing to support the REPIH/RELCH. However, he notes that the sample period is not homogenous, and argues that different exchange rate regimes may change the transmission mechanism to macroeconomic shocks. Modifying **(2.3.9)** to allow for variable real interest rates and testing the REPIH/RELCH over two separate periods corresponding to fixed and flexible exchange rate regimes, Muellbauer (1983) finds evidence of excess sensitivity to the income innovation in the former and intertemporal substitution in the latter. Overall, Muellbauer (1983) suggests rejection of the REPIH/RELCH.

In recent tests, Gausden and Myers (1997) find excess sensitivity using a panel of regional UK data as do Fan and Wong (1998) using aggregate time series for Hong Kong. Caballero's (1994) assessment is that "Researchers now seem to agree that *Hall's (1978)* implication of the PIH does not hold in the data, regardless of the country and sample used." (p. 107; my comments in italics).

Campbell (1987) develops an alternative testable implication of the (RE)PIH. If true, consumption is proportional to *permanent* income, therefore, when *current* income is below (above) *permanent* income consumers are dissaving (saving) in the [rational] expectation that *current* income will rise (fall). "Put another way, dissaving anticipates rising income and saving anticipates falling income. People save for a 'rainy day.'" (Campbell 1987, p. 1250). Saving should be at least as good a predictor of declines in future *labour*, not capital, income as any other forecast based upon publicly available information.[19]

Campbell (1987) finds that US consumption and *labour* income are first difference stationary and constructs a stationary linear combination of consumption and *disposable* income, which is

[19] Saving increases capital income which would offset declining labour income.

presented as a quasi-saving variable. The weak implication of the PIH is supported because (quasi) saving Granger causes changes in *labour* income. However, the stronger implication is rejected because the forecasts of the present value of future declines in labour income exhibit a larger standard deviation than *actual* saving. The lower variability in *actual* saving implies that consumption features a higher correlation with current income than predicted by the PIH (*excess sensitivity*).

MacDonald and Speight (1989) find that saving *positively* Granger-causes changes in UK labour income, contradicting the REPIH/RELCH. However, Attfield *et al.* (1990), after amending an error in MacDonald and Speight's calculation of real labour income, find that saving *negatively* Granger-causes changes in labour income, consistent with the *weak* implication of the REPIH/RELCH. Campbell and Clarida (1987) confirm support for this weak implication for the UK and Canada. MacDonald and Speight (1989) found support for the stronger implication of the REPIH/RELCH when allowance is made for transitory consumption. However, Attfield *et al.* (1990) argue that they fail to allow for an MA(1) error process implied by the presence of transitory consumption. Allowing for this MA(1) process Attfield *et al.* (1990) reject the strong implication for the UK.[20]

Given the persistence of positive aggregate saving in industrial economies Campbell's (1987) formulation implies, according to Muellbauer (1994), continual declines in future aggregate income, which is counterfactual. This suggests rejection of (linear) REPIH/RELCH models, based upon quadratic utility functions and employing the simplifying assumption that the interest rate is constant and equal to the subjective discount rate, such as Campbell's.

Muellbauer (1994) argues that another implication of the REPIH/RELCH Euler equation is that *surprises* in consumption should equal the shocks to permanent income. *If* actual income exhibits a unit root (Muellbauer (1994) suggest that standard tests indicate a unit root in per capita

[20] However, MacDonald and Speight (1990) maintain their support for the REPIH/RELCH when allowing for both MA *and* non-constant autocovariance (ARCH) processes.

income[21]), all income shocks are translated into changes in permanent income and the best estimate of permanent income is current income. Thus, the REPIH/RELCH implies that the variance of the consumption innovation should be at least as large as the variance of the current income innovation. Muellbauer (1994) cites evidence that the variance of consumption growth is only half that of current income growth for various industrial countries. Thus, consumption is *excessively smooth* compared to the prediction of the REPIH/RELCH *if* income is persistent (features a unit root).

Overall, the evidence suggests rejection of various implications of the basic linear REPIH/RELCH. Subsequent research has focused upon the modification Hall's (1978) random walk model by removing one or more of the assumptions upon which it is based.

2.3.2 Adjustment Costs and Durability

Caballero (1994) suggests that the REPIH/RELCH applied to durable expenditures is more emphatically rejected than when applied to non-durables. This will have implications for models applied to total expenditures or with any significant element of durability.

The common practice of specifying convex (usually quadratic) adjustment costs in representative agent models is argued to be counterfactual at the microeconomic level because durable purchases are typically "sporadic and lumpy rather than continuous and smooth" (Caballero 1994, p. 108). The intermittent adjustment of the stock of durables at the microeconomic level can arise from fixed adjustment costs.[22] Such costs may cause microeconomic agents to tolerate *small* departures from an ever changing *optimal* level of the durable stock. Once departures are no longer considered *small* the consumer abruptly buys or sells to make the disequilibrium tolerable once more.

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[21] Tests allowing for fractional integration, for example, Sowell (1992), and broken trends, for example, Perron (1989), may suggest that income is not best characterised as exhibiting a unit root.

[22] Transactions costs, time spent searching amongst heterogeneous products and imperfections in secondary markets (like *lemons*) are examples of fixed adjustment costs.

Hall's (1978) principal insight suggests that individuals smooth their consumption over time and so abrupt changes in marginal utility must be brought about by *surprises* in permanent income. Assuming separability across goods and time this model can be applied to the services from durables. Assume that apart from trends (due to economic growth) additions to the aggregate stocks of durables (K_t) are random:

$$\Delta K_t = \epsilon_t^D, \qquad E_{t-1}(\epsilon_t^D) = 0. \tag{2.3.10}$$

Let the stock of durables, K_t, depreciate geometrically at the rate, γ, as follows:

$$K_t = (1-\gamma)K_{t-1} + CD_t \tag{2.3.11}$$

where CD_t is the *flow* of expenditures on consumer durables. Taking first differences yields:

$$\Delta K_t = (1-\gamma)\Delta K_{t-1} + \Delta CD_t. \tag{2.3.12}$$

Substituting **(2.3.10)** into **(2.3.12)** and re-arranging gives Mankiw's (1982) MA(1) specification for durable expenditures under the REPIH\RELCH:

$$\Delta CD_t = \epsilon_t^D - (1-\gamma)\epsilon_{t-1}^D. \tag{2.3.13}$$

With durables lasting for more than one period the stock does not require replacement expenditures each period except for depreciation. This depreciation implies a coefficient on the MA(1) term just slightly less negative than minus one. Using post-war quarterly US data, Mankiw (1982) finds that the MA coefficient is not significantly different from zero, suggesting rejection of the REPIH/RELCH (which implies an MA process for durables). Caballero (1994) suggests that an amended version of **(2.3.10)**, where adjustment costs causes replacement expenditures to be spread over several periods, may explain the data better. That is:

$$\Delta K_t = \alpha\epsilon_t^D + (1-\alpha)\epsilon_{t-1}^D, \tag{2.3.14}$$

which when substituted into **(2.3.12)** gives,

$$\Delta CD_t = \alpha\epsilon_t^D + \{(1+\alpha\gamma-2\alpha)/\alpha\}\alpha\epsilon_{t-1}^D - \{[(1-\alpha)(1-\gamma)]/\alpha\}\alpha\epsilon_{t-2}^D \quad \textbf{(2.3.15)}$$

Letting $v_t = \alpha\epsilon_t^D$ yields:

$$\Delta CD_t = v_t + \{(1+\alpha\gamma-2\alpha)/\alpha\}v_{t-1} - [\{(1-\alpha)(1-\gamma)\}/\alpha]v_{t-2}. \quad \textbf{(2.3.16)}$$

Notice that when α is sufficiently below unity the coefficient on the first moving average term can be close to zero. Further, the sum of the moving average terms is $-\{1-(\gamma/\alpha)\}$. Caballero (1994) argues that provided adjustment costs are not excessively large, $\alpha > \gamma$, the large negative MA(1) coefficient indicated by the model without adjustment costs, **(2.3.13)**, is spread out over several MA terms and reflected in the sum of their coefficients.[23] Caballero (1994) estimates a model like **(2.3.16)** with fifteen moving average error terms using quarterly data for three durable expenditure categories. The sums of the coefficients, in all three cases, are statistically significant and negative, which is consistent with the slow adjustment interpretation ($\alpha > \gamma$), supporting the REPIH/RELCH modified for durability.[24]

2.3.3 Intertemporal Substitution

Summers (1982), Hansen and Singleton (1983), Muellbauer (1983), Wickens and Molana (1984) and Mankiw, Rotemberg and Summers (1985), for examples, have all estimated models which relax the assumption of fixed real interest rates. They all find a statistically significant and positive coefficient on real interest rates, suggesting this as a potential explanation for the rejection of the *pure* REPIH/RELCH. However, Hall (1988) develops a model which explains

[23] If $\alpha = \gamma$, the moving average error terms will sum to zero and if $\alpha < \gamma$, their sum will be positive.

[24] According to the Box and Jenkins method of ARIMA model identification the number of statistically significant autocorrelation coefficients indicates the order of MA process. Further, if more than the first four consecutive autocorrelation coefficients are statistically significant this is typically considered indicative of a nonstationary process. *If*, therefore, 15 MA error terms implies that the first 15 autocorrelation coefficients are statistically significant, this *might* suggest the process is nonstationary. Thus, one *may* wish to view Caballero's (1994) results with caution.

and rejects (encompasses) the findings of these researchers. To develop an intertemporal substitution model one takes the natural logarithms of both sides of **(2.2.13)**, allowing interest rates to vary and substituting **(2.3.5)** into the result yields:

$$\ln C_{t+1} = \{\ln[1/(1+\delta)]\}/\beta + (1/\beta)\ln(1+E_t r_{t+1}) + \ln C_t + \epsilon_{t+1}. \qquad \textbf{(2.3.17)}$$

Hall's (1988) encompassing model is obtained by using the approximation $E_t r_{t+1} \approx \ln(1+E_t r_{t+1})$, defining $b_{15} = \{\ln[1/(1+\delta)]\}/\beta$ and $\sigma = (1/\beta)$, and subtracting $\ln C_t$ from both sides of **(2.3.17)**, thus:[25]

$$\Delta \ln C_{t+1} = b_{15} + \sigma E_t r_{t+1} + \epsilon_{t+1}, \qquad \textbf{(2.3.18)}$$

where σ is the intertemporal elasticity of substitution which has also been interpreted as the reciprocal of the coefficient of relative risk aversion, $1/\beta$, see Hansen and Singleton (1983).[26]

Hall (1988) argues that previous researchers' instrumental variable (IV) estimates of the intertemporal elasticity of substitution are subject to bias arising from instrumenting interest rates with variables dated in period t-1 which, due to time aggregation, will be correlated with the error term. Using instruments dated no earlier than period t-2, which will not be subject to this bias, Hall (1988) finds the intertemporal elasticity of substitution to be negative and therefore concludes that the inference of a large and positive intertemporal elasticity of

[25] Strictly, Hall (1988) uses interest rates lagged one period behind consumption growth. We specify interest rates and consumption as contemporaneous to be consistent with Campbell and Mankiw (1991).

[26] Hall (1988) cites work within the ordinal certainty equivalence literature and representations of intertemporal preferences under uncertainty which depart from the expected utility framework and suggest the inverse relationship between σ and β does not always hold. "It is an unambiguous conclusion that the intertemporal elasticity of substitution alone controls the relation between consumption growth and the expected real interest rate.... the bivariate relation between consumption and real interest rates does not necessarily reveal anything about risk aversion." (Hall 1988, pp. 344-345). Obstfeld (1994) suggests that risk aversion and intertemporal substitution parameters can be separately identified when allowance for uncertainty is made using "non-expected-utility preferences".

substitution was due to time aggregation bias and that the adoption of appropriate estimation techniques led to implausible negative values of σ. This intertemporal substitution model has been subsequently rejected by, for example, Campbell and Mankiw (1989) and Jin (1994).

Hahm (1998) argues that the evidence against intertemporal substitution may be due to using the incorrect measure of consumption and excluding instruments of sufficient lag length. Regarding the former, Hahm (1998) highlights the aggregation of housing services with other expenditures. This may be particularly problematic because, for example, both homeowners and renters will find it prohibitively costly to continuously alter their housing consumption in the face of frequently changing interest rates - housing consumption will either have a lower degree of intertemporal substitution relative to non-durables, or adjustment will not be as smooth as for non-durables. It is therefore argued that use of non-durables is a more appropriate measure of consumption and that the addition of services could cause misleading results. Hahm (1998) finds evidence favouring the presence of both current income consumers and intertemporal substitution for non-durable US consumption. When these models were estimated using non-durables and services as the measure of consumption, no statistically significant relationship between consumption growth and interest rate is revealed. This suggests the need to accommodate consumer expenditure series including durable components, especially housing services.

Attanasio and Weber (1989) separate the effects of intertemporal substitution and relative risk aversion by considering the correlations of rates of return on two UK assets of different risk with consumption. Using a certainty equivalence model applied to cohort data (which excludes those of ages likely to be liquidity constrained) they obtain Three Stage Least Squares (3SLS) estimates, allowing for time aggregation, which indicate a positive and statistically significant intertemporal elasticity of substitution.

Overall the evidence is ambigous, if generally unfavourable, regarding the role of variable real interest rates in explaining the rejection of the REPIH/RELCH.

2.3.4 Liquidity Constraints, Current Income Consumers and Precautionary Saving

Hadjimatheou (1987) suggests that consumption may be more sensitive to *current* income than predicted by the PIH-LCH for the following reasons. Firstly, a large proportion of the population may have few assets upon which to draw when their current income is insufficient to support their optimal level of consumption. Secondly, in periods of high unemployment a substantial percentage of consumers may find that their labour has become an "*illiquid*" good: it cannot be exchanged for money with which to maintain their desired level of consumption. Thirdly, capital market imperfections may make it impossible, or extremely expensive, for consumers to borrow against their potential future earnings. Fourthly, those who do own durables may be discouraged from selling them to obtain funds to support their present consumption due to substantial transactions costs. Thus, removal of capital market imperfections can be considered using variables like unemployment to proxy the degree of liquidity constraints or allowing consumption to be more sensitive to current income.[27]

Hadjimatheou (1987) argues that there are several justifications for entering unemployment in the consumption function: as a proxy for uncertainty about future income, income distribution, the cyclical movement of consumption and liquidity constraints. He further notes that unless unemployment is introduced in a way which discriminates between these alternative hypotheses there is no way of assigning the expected sign to unemployment a priori. This is borne out by the empirical evidence which is neither clear on the significance or sign of unemployment in a consumption function. For example, Arestis and Hadjimatheou (1982) and Muellbauer (1983) find a positive relationship between unemployment and consumption while Ouliaris (1981), Flavin (1985) and Malley and Moutos (1996) find a negative effect. Townend (1976) finds both positive and negative effects (depending upon whether unemployment is entered contemporaneously or with a lead) while Koskela and Viren (1986) find that unemployment does not significantly affect consumption for nine countries, though it has an implicit positive relationship for three. This suggests that alternative means of accommodating liquidity

[27] One could also use direct measures of credit, although one needs to ensure they reflect supply rather than demand side factors.

constraints is desirable.

Hall (1978) suggests the following modification to the REPIH/RELCH, "The simplest alternative hypothesis supposes that a fraction of the population simply consumes all of its disposable income, instead of obeying the life cycle-permanent income consumption function." (Hall 1978, p. 977). The rest of the population consume according to their optimal life-time plan. Thus, one may define aggregate consumption as the sum of unconstrained (C_t^U) and constrained consumption (C_t^C):

$$C_t = C_t^C + C_t^U. \tag{2.3.19}$$

A popular specification utilising this idea was introduced by Hall and Mishkin (1982) who, in essence, utilise the above equation in difference form:

$$\Delta C_t = \Delta C_t^C + \Delta C_t^U. \tag{2.3.20}$$

Unconstrained consumers are specified as following Hall's (1978) model, equation **(2.3.7)**, multiplied by the proportion of the population who are unconstrained, (1-π). Constrained consumers consume all of their income each period which is their share, π, out of current income. In first differences constrained consumption is:

$$\Delta C_t^C = \pi \Delta Y_t. \tag{2.3.21}$$

Substituting (1-π) multiplied by **(2.3.7)** and **(2.3.21)** into **(2.3.20)** gives the essential form of Hall and Mishkin's (1982) model:[28]

$$\Delta C_t = \pi \Delta Y_t + (1-\pi)\epsilon_t. \tag{2.3.22}$$

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[28] Hall and Mishkin (1982) derive the model in a slightly different manner by assuming aggregate consumption is given by, $C_t = (1-\pi)C_t^U + \pi C_t^C$, taking $(1-\pi)C_t^U$ as unconstrained consumption and πC_t^C as constrained. Our representation yields essentially the same model except we call C_t^U and C_t^C unconstrained and constrained consumption, respectively.

Deaton (1992) suggests that *excess smoothness* can be resolved by **(2.3.22)** because it implies that the innovation in consumption is a weighted average of the permanent income innovation for unconstrained and constrained consumers, ϵ_t. That is, if the innovations in income of the constrained and unconstrained are not highly correlated and the variance of ϵ_t is less than the variance of the income surprise then the variance of total consumption, the weighted average of income innovations, will be less than that of income: *excess smoothness*.[29]

Jappelli and Pagano (1989) seek to determine whether liquidity constraints are the cause of excess sensitivity by considering whether this excess sensitivity is larger in countries whose capital markets are less well developed. They estimate the following model for Sweden, the US, the UK, Japan, Italy, Spain and Greece using annual data with consumption measured as per capita non-durable expenditures.

$$C_t = b_{16} + b_{17}C_{t-1} + \pi(Y_t - b_{17}Y_{t-1}) + \epsilon_t. \qquad \textbf{(2.3.23)}$$

The excess sensitivity parameter, π, is significant for all countries except Sweden. Its value varies widely across countries being highest for Italy, Spain and Greece and lowest for the US and Sweden. They conclude that "the fact that consumer debt is low in countries where the excess sensitivity of consumption is high can be interpreted as evidence that liquidity constraints in the form of quantity rationing are at the source of the empirical failures of the LC-PIH in time-series tests." (Jappelli and Pagano 1989, p. 1089). However, the reliability of inference may be undermined by the inclusion of variables lagged one period to instrument Y_t in equation **(2.3.23)** and the assumption that the regressors are trend stationary.

Zeldes (1989) estimates an Euler equation for constrained and unconstrained consumers, where those with (near) zero wealth are assumed constrained. Using US household panel data Zeldes (1989) finds that lagged income can only be added to the constrained consumers' sub-sample and concludes that liquidity constraints are the cause of the REPIH/RELCH's failure for the

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[29] However, if income has a unit root then the income innovations may be highly correlated and so the consumption variance may not be less than the variance of income.

US.[30]

Fissel and Jappelli (1990) argue that if current income consumers are liquidity constrained their behaviour is affected by supply factors which may vary. Variations in the supply of credit suggests that the proportion of liquidity constrained consumers is *endogenously* determined and cannot be presumed constant. In contrast to much previous literature, Fissel and Jappelli (1990) assume that a consumer may be constrained at different points in time, depending upon explanatory factors such as current resources and proxies for future resources including age, schooling, sex and race. They find that both the probability of being liquidity constrained and the proportion of total income consumed by constrained consumers in the US varies significantly over the period 1969-1982: both are endogenously determined.

Campbell and Mankiw (1991) - hereafter CM - derive a stationary log-linear form of the current income consumer model for non-durable expenditures. They assume that the growth in total consumption is approximately given by:[31]

$$\Delta \ln C_t \approx \pi \Delta \ln C^C_t + (1-\pi)\Delta \ln C^U_t. \qquad \textbf{(2.3.24)}$$

Substituting **(2.3.8)**, and $\Delta \ln C^C_t = \Delta \ln Y^C_t = \Delta \ln(\pi Y_t) = \Delta \ln Y_t$, into **(2.3.24)** gives CM's equation (19), reproduced below:

$$\Delta \ln C_{t+1} = b_9 + \pi \Delta \ln Y_{t+1} + \epsilon_{t+1}. \qquad \textbf{(2.3.25)}$$

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[30] Jappelli (1989) suggests caution in interpreting evidence from studies such as Zeldes (1989), which assume that high wealth households are unconstrained and low wealth households constrained. For example, it is suggested that high wealth consumers may be liquidity constrained due to the transactions costs involved in realising illiquid assets.

[31] Strictly this approximation implies, $C_t \approx (C^C_t)^{\pi}(C^U_t)^{(1-\pi)}$, which is not the definition of total consumption given by **(2.3.20)**. However, CM note that, using US data, log and level formulations yield similar estimates of π, suggesting the log-linear approximation does not significantly affect the estimate of the parameter of interest. Jin (1994) notes that the interpretation of π in the CM model is the proportion of expenditures made by current income consumers rather than the proportion of income accruing to them.

Relaxing the assumption of constant interest rates, analogous to **(2.3.18)**, gives rise to the following extended current income consumer model:

$$\Delta \ln C_{t+1} = b_{18} + \pi \Delta \ln Y_{t+1} + \sigma E_t r_{t+1} + \epsilon_{t+1}. \qquad \textbf{(2.3.26)}$$

CM estimate these models using IV with instruments dated in period t-2 or earlier to allow for any first order autocorrelation arising from time averaged data, white noise measurement error in the levels of consumption and income or taste shocks and expenditure measures incorporating a durable component. They find no support for intertemporal substitution so their favoured models exclude interest rates. In these preferred specifications the proportion of current income consumers ranges from 0.2 in Canada to 0.35 in Sweden and the US, to nearly 1.0 for France and 0.35 or 0.65 (depending upon seasonal adjustment) in the UK. There is no reliable estimate for Japan. They note that countries with larger values feature less well developed consumer credit markets, consistent with these being estimates of the proportion of liquidity constrained consumers. Except for the UK, they find no evidence that π varies through time suggesting that factors such as unemployment may have offset the effects of deregulation (which were expected to cause time variation).[32]

Church *et al* (1994) estimate an analogue of the Weale model, a REPIH/RELCH specification augmented for current income consumers and, following Blanchard (1985), finite planning horizons. Recursive estimation indicates that the percentage of liquidity constrained consumers has remained relatively constant at 16%. This implies that financial deregulation has had no effect or that there are two offsetting influences. For example, if the prospect of liberalisation raised income growth expectations during the 1980s households would wish to increase borrowing, compared to the situation without deregulation, raising the proportion of constrained consumers. Conversely, the actual relaxation of constraints lowers the proportion of frustrated

[32] For the UK CM estimate that the proportion of current income consumers is increasing over the period 1957-88 suggesting, counterfactually, that credit rationing has intensified! Muellbauer (1994) argues that this raises doubts over the rational expectations assumption of the Euler equation for the UK because relating consumption growth solely to expected income growth means that the only way the Euler equation can explain the mid-1980s consumer boom is by allowing the share of credit-constrained households to grow.

households, possibly by a similar amount to the offsetting influence. They are unable to establish the superiority of one hypothesis over the other.

Bacchetta and Gerlach (1997) find that the proportion of UK credit constrained consumers rises through time for the UK, falls for the US and does not vary for Canada, France and Japan. In contrast, Bayoumi (1993) and Blundell-Wignall *et al* (1995) present evidence consistent with deregulation UK causing a reduction in UK credit constraints through time, possibly with or without an offsetting influence. Darby and Ireland (1994) derive a model where consumers are forward looking, have finite planning horizons and there is a proportion of current income consumers which is allowed to vary with the degree of financial liberalisation. Their model indicates that π was 33.4% prior to deregulation, falling to a low of 16.1% in 1988 and growing back to 25% by 1990. Their simulations, based on the estimated model, indicate an unambiguous rise in UK consumption due to deregulation of 2.87% per annum, which is similar to the 2.25% reported by Bayoumi (1993). McKiernan (1996) presents evidence indicating that π varies from 0.1 to 0.6, with a mean value of 0.33, in the USA and that this time variation is related to liquidity constraints. This evidence suggests that the existence of a proportion of current income consumers explains the rejection of the REPIH/RELCH and that this proportion systematically varies with the degree of liquidity constraints for some countries.

Jin (1994) estimates the proportion of current income consumers and also assesses whether they vary with credit conditions across countries. Jin (1994) derives analogues of CM's formulations which ensures the coefficient on current income is appropriately interpreted as the proportion of current income consumers.[33] The derived models are:

[33] Unconstrained consumers' behaviour is described by equation **(2.3.8)** lagged one period, $\Delta \ln C_t^U = b_9 + \epsilon_t$, and constrained consumers spend all their income, which is some proportion, π, of total income thus, $C_t^C = Y_t^C = \pi Y_t$, or in growth form, $\Delta \ln C_t^C = \Delta \ln(\pi Y)_t = \Delta(\ln\pi + \ln Y_t) = \Delta \ln Y_t$. It is then noted that $\Delta \ln C_t \approx \Delta C_t/C_{t-1}$ which, after substitution of the *exact* definition of non-durable consumption, $C_t = C_t^U + C_t^C$, yields, $\Delta \ln C_t \approx \Delta(C_t^U + C_t^C)/C_{t-1} = (\Delta C_t^U + \Delta C_t^C)/C_{t-1}$. The approximation, $\Delta \ln C_t \approx \Delta C_t/C_{t-1}$, also implies, $\Delta C_t \approx C_{t-1}\Delta \ln C_t$. Substituting this into $(\Delta C_t^U + \Delta C_t^C)/C_{t-1}$, *approximately* yields, $(C_{t-1}^U \Delta \ln C_t^U + C_{t-1}^C \Delta \ln C_t^C)/C_{t-1}$. Rearranging, $C_t = C_t^U + C_t^C$, gives $C_{t-1}^U = C_{t-1} - C_{t-1}^C$, which when substituted into $(C_{t-1}^U \Delta \ln C_t^U + C_{t-1}^C \Delta \ln C_t^C)/C_{t-1}$, gives, $([C_{t-1} - C_{t-1}^C]\, \Delta \ln C_t^U + C_{t-1}^C \Delta \ln C_t^C)/C_{t-1}$. Substituting our expressions for constrained and unconstrained consumption and consumption growth into this last expression yields, $\Delta \ln C_t \approx$

$$\Delta \ln C_t = b_9 + \pi(Y_{t-1}/C_{t-1})\Delta \ln Y_t - b_9\pi(Y_{t-1}/C_{t-1}) + z_t \qquad \textbf{(2.3.27)}$$

where, $z_t = [1-\pi(Y_{t-1}/C_{t-1})]\epsilon_t$.

Allowing rates of return to vary yields:

$$\Delta \ln C_t = b_{19} + \pi(Y_{t-1}/C_{t-1})\Delta \ln Y_t + \sigma r_t - b_9\pi(Y_{t-1}/C_{t-1}) - \pi\sigma(Y_{t-1}/C_{t-1})r_t + z_t. \qquad \textbf{(2.3.28)}$$

Jin (1994) estimates these models using generalised methods of moments (GMM) for nineteen OECD countries implicitly accounting for any moving average error, due to the use of total consumption for the dependent variable, by employing Newey and West (1987) adjusted standard errors. Using an approximate measure of private disposable income Jin secures estimates of the proportion of current income consumers which are statistically significant and between zero and one for all but Luxemborg and Switzerland of the nineteen OECD economies he considers.[34] The REPIH/RELCH modified by a proportion of constrained consumers is therefore interpreted to have been supported. Similar estimates are obtained from a pooled regression.[35] These pooled estimates are found to vary with the degree of liquidity constraints, supporting their constrained consumer interpretation. There is little evidence supporting intertemporal substitution.

Bacchetta and Gerlach (1997) argue that CM's model, **(2.3.25)**, may be overly restrictive. "First, most consumers are able to borrow, at least to some extent. This suggests that if the liquidity

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$\{[(C_{t-1}-\pi Y_{t-1})(b_9+\epsilon_t)] + [(\pi Y_{t-1})\Delta \ln Y_t]\}/C_{t-1}$. Which after simplification gives **(2.3.27)**.

[34] Private disposable income is conventionally measured as real national disposable income less real general government disposable income. The approximate measure used for all nineteen countries is national disposable income minus government consumption expenditure, which does not account for the general government saving component of government income.

[35] Jin's (1994) *pooled* estimates of the proportions of credit constrained consumers are Australia 0.257; Austria 0.350; Belgium 0.489; Canada 0.473; Denmark 0.528; Finland 0.539; France 0.326; Germany 0.428; Greece 0.337; Ireland 0.639; Italy 0.499; Japan 0.544; Luxembourg 0.043; the Netherlands 0.493; Norway 0.368; Sweden 0.496; Switzerland 0.316; the UK 0.414) and the USA 0.369.

constraint interpretation is to be taken seriously, credit should be incorporated in the analysis. Secondly, the proportion of constrained consumers is unlikely to be constant. It seems plausible that fewer people are constrained in good times or when access to credit is easier." (Bacchetta and Gerlach 1997, p. 210). If credit constraints vary, it is argued that variables other than income need to be explicitly included in a model. Thus, they essentially suggest the following extension of **(2.3.25)**:

$$\Delta \ln C_t = b_{20,t} + \pi_t \Delta \ln \mathbf{x}_t + \epsilon_t, \qquad \textbf{(2.3.29)}$$

where π_t is a vector of time-varying coefficients which correspond to the growth rates of the vector of variables, including income and those characterising credit conditions, $\mathbf{x}_t$. When credit and income growth variables are included simultaneously the former is significant while the latter sometimes becomes insignificant. This highlights the importance of credit variables in such regressions, perhaps overshadowing the income variable.

Gausden and Myers (1997) find evidence suggesting that the inclusion of unemployment growth causes income growth to become insignificant in a CM-type model for the UK. This seems to add support to Bacchetta and Gerlach's (1997) extension of the CM model. While Acemoglu and Scott (1994) find, for the UK, that a consumer confidence indicator enters with significance in a CM-type model and causes income growth to become insignificant.[36] They argue that this indicates that the rejection of the REPIH/RELCH is due to precautionary saving rather than the presence of liquidity constraints.[37] This interpretation may also rationalise the dominance of unemployment over income growth in modified REPIH/RELCH specifications.

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[36] Carroll *et al* (1994) find evidence which supports the predictive power of consumer confidence for consumption growth in the US while Fan and Wong (1998) offer evidence against this for Hong Kong. The latter suggest this result may be due to the confidence indices including information beyond expectations about future income, especially about their future well being (with the transfer of sovereignty from Britain to China).

[37] Muellbauer (1994) points out that income uncertainty cannot be accommodated in a (RE)PIH model utilising a quadratic utility function. Pemberton (1993) argues that this is because it produces marginal utility which is linear in consumption so does not reflect the motive underlying precautionary behaviour, being that low levels of consumption yield disproportionately low levels of (marginal) utility. The CES/CRRA utility function does.

Unlike the studies reviewed above, Carroll (1994) considers liquidity constraints and precautionary saving variables simultaneously to gauge if one or the other is dominant. Using US household data he finds consumption to be excessively sensitive to current income and excessively smooth to changes in future expected income. Carroll (1994) finds that liquidity constraints cannot explain these results. However, he does find that precautionary saving can explain these findings using three measures of income uncertainty.[38] That is, all three measures of income uncertainty depress consumption, however, only the favoured measure does so with statistical significance. Carroll's (1994) results are argued to be consistent with precautionary saving, rather than liquidity constraints, explaining both excess smoothness and excess sensitivity.

A similar examination of both liquidity constraints and income uncertainty is provided by Hahm and Steigerwald (1999) who find, using US time-series data, that "the excess sensitivity of consumption to current income may be partially explained by the role of time-varying income uncertainty operating through precautionary saving." (Hahm and Steigerwald 1999, p. 39). They also find that, after conditioning income on income uncertainty, income growth rates have less explanatory power for consumption growth. Thus, excess sensitivity is due, in some part at least, to income uncertainty, possibly in conjunction with liquidity constraints.[39]

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[38] Standard measures of income uncertainty are the standard deviation/variance of income *after the predictable component of income changes have been removed* and normalised by mean income/the square of mean income, to make them dimensionless. However, neither of these measures are considered good proxies of uncertainty "in the sense of a measure which theory is a sufficient statistic for the amount of precautionary saving that will be induced by a given income distribution." (Carroll 1994, p. 136). Carroll suggests that if the theory of precautionary saving is correct a more appropriate measure *might* be the *equivalent precautionary premium* derived by Kimball (1990) which is "a direct measure of the intensity of the precautionary saving motive at the point of zero precautionary saving." (Carroll 1994, p. 136). All three measures of income uncertainty are shown to yield similar patterns of uncertainty across occupational and educational types with all suggesting the highest degree of income uncertainty for the self employed and the lowest uncertainty for professionals and highly educated workers.

[39] Muellbauer (1994) argues that the large amount consumers spend on other types of insurance indicates that it would seem "incontrovertible" that they build up precautionary saving as insurance against future income uncertainty.

Overall, there is strong evidence supporting the presence of a proportion of current income consumers and that this proportion, in general, varies across countries. Further work on whether liquidity constraints and/or income uncertainty explain this variation seems warranted. Whether the proportion of current income consumers varies through time or if credit variables, unemployment and indicators of consumer confidence should be added to the model or replace income growth altogether remain contentious.

2.3.5 Uncertainty Over Life Length and the Finiteness of the Planning Horizon

Muellbauer (1994) offers a warning to the research on the linear REPIH, which requires the use of a quadratic utility function, being that it is assumed that age is irrelevant to decisions, agents act as if they are infinitely lived and so their probability of survival is independent of age.[40] Uncertainty over the length of life, even in the absence of a bequest motive, may lead to the leaving of substantial assets, particularly when life is ended prematurely implying that the subjective discount rate incorporates an individual's survival probability. This helps explain why, according to survey evidence, the retired do not dissave as much as implied by the simple LCH. Indeed, consumption needs of the elderly become more uncertain as the likelihood of health failure increases requiring more protection to cover such expenses. Thus, the MPC out of assets is likely to lower and far less variable than predicted by the basic LCH. Further, aggregation bias becomes a less important explanatory factor in differences in the saving rates of economies with different rates of population and income growth. Uncertainty over life length can also explain large scale bequests.

Blanchard (1985) develops a model which allows the discount rate on non-interest income to be greater than the interest rate, so characterising consumers with finite planning horizons. This facilitates aggregation over households of varying age through the introduction of uncertainty

[40] Muellbauer and Murphy (1994) argue that consumers uncertainty is greater the further are projections of income into the future, suggesting that *far* future incomes should be more heavily discounted than *near* future incomes. While Pemberton (1993) convincingly criticises the "unreality about the notion that consumers make detailed allocation plans for the far-distant future" (p.10), which is implicit in all work based upon the PIH-LCH.

over households time of death. To obtain a tractable model Balnchard (1985) makes two assumptions. Firstly, all consumers are assumed to face a constant instantaneous probability of death, p, throughout their life so that their expected life length is 1/p. Although uncertainty surrounds any particular individual's time of death a *large* cohort of agents' uncertainties is assumed to decline non-stochastically through time. Secondly, it is assumed that life assurance companies exist that will pay consumers the proportion 1/p of their wealth each period until they die, when the company receives the agent's wealth (there are no bequests or unpaid debts left upon death). The large size of cohorts allow life assurance companies to be free of risk and it is assumed that they reap no profit.

A household's behaviour depends upon p. When p=0 the planning horizon, 1/p, is infinite while p>0 indicates a finite planning horizon. Blanchard (1985) derives a continuous time aggregate consumption function, using a logarithmic utility function, which depends upon the parameter p. In general, the model predicts that both household and aggregate consumption are functions of human and non-human wealth. When the planning horizon is finite (p>0) the discount rate on labour income exceeds the interest rate and so aggregate consumption (growth) remains a function of human and non-human wealth. However, in the infinite horizon case (p=0) non-human wealth is eliminated leaving the standard Hall (1978) style consumption function.

The main policy implication is that a present reduction in taxes financed by a future increase in taxes (either directly or to finance increased government debt) has a decreasing impact upon current consumption as p tends to zero (the planning horizon tends to infinity). However, the more finite the planning horizon, the greater the impact of such a policy on current consumption because some agents, particularly those nearing the end of their life, will receive the increase in wealth from the tax cut but not expect to be alive when taxes are increased in the future. Thus, when planning horizons are finite governments can, in principle, utilise fiscal policy to smooth consumption over the business cycle.

Evans (1988) derives a discrete time analogue of Blanchard's (1985) model to empirically test whether planning horizons are infinite or not or, equivalently, whether consumers are Ricardian or not. Evans (1988) formulation provides a well specified model which nests both Ricardian equivalence and an alternative in which households treat government debt as net wealth. In this model, aggregate consumption depends upon (expected) resources and the planning horizon may be finite because a fraction of households, p, die each period.

$$C_t = b_{21}\{(1+r_t)A_{t-1} + \sum_{i=0}^{\infty}(1-p)^i b_{22,it}E_tY_{t+i}\}, \qquad \textbf{(2.3.30)}$$

where $0 < b_{21} < 1$, $0 \leq p < 1$ and $b_{22,it}=1$ for i=0 and,

$$b_{22,it} = 1 / \prod_{j=1}^{i}(1+F_{jt}), \quad i > 0, \qquad \textbf{(2.3.31)}$$

where F_{jt} is the forward real interest rate in period t on bonds that will be issued in period t+j-1 and will mature in period t+j. Using the budget constraint to eliminate income from **(2.3.30)** and invoking the common assumption that the forward real rate of interest is constant and equal at every horizon, $b_{22,it} = b_{22}^i$, gives, after some manipulation:

$$C_t = \{(1-b_{21})/[b_{22}(1-p)]\}C_{t-1} - [b_{21}p/b_{22}(1-p)]A_{t-1} + u_t, \qquad \textbf{(2.3.33)}$$

where,

$$u_t = b_{21}\sum_{i=0}^{\infty} (1-p)^i b_{22}^i (E_t-E_{t-1})(C_{t+i} + pA_{t+i}). \qquad \textbf{(2.3.34)}$$

Ricardian equivalence holds if the coefficient on A_{t-1} is not significantly different from zero (p=0). However, if this coefficient is negative and significant (p>0), Blanchard's (1985) alternative cannot be rejected. Evan's extends **(2.3.34)** to allow interest rates to vary, giving:

$$\Delta \ln C_t - \ln(1+r_t) = b_{23} - [(1-b_{24})/b_{24}]p(A_{t-i}/C_{t-i}) + u_t \qquad \textbf{(2.3.35)}$$

If the OLS estimate of the coefficient on (A_{t-i}/C_{t-i}) is statistically insignificant, Ricardian equivalence holds while a significant and negative estimate suggests that Blanchard's alternative is favoured. Estimating **(2.3.33)** and **(2.3.35)** for the US, Evans (1988) finds evidence rejecting

Blanchard's (1985) alternative to Ricardian equivalence.[41] Intervention analysis suggests that tax cuts did not have a significant impact upon US consumption further supporting Ricardian equivalence. Ricardian equivalence is argued to yield a *reasonable approximation* of quarterly postwar US data.

However, Hayashi (1982), Weale (1990) and Darby and Ireland (1994), using specifications based upon Blanchard's (1985) work, find that the discount rate exceeds the real rate of interest, which refutes Ricardian equivalence.[42] Muellbauer (1994) alternatively interprets this as evidence favouring the existence of a risk premium consistent with income uncertainty rather than uncertainty over life length.

Uncertainty over life length can explain current income consumers without appeal to income uncertainty or liquidity constraints. Leung (1994), employing the standard simulation models of Yaari (1965), finds that many individual consumers' savings will, with uncertainty over survival, be depleted before death due to a finite planning horizon. Thus, those retired consumers with depleted saving may become current income consumers for the remainder of their lives.[43]

Weil (1991) suggests that there exist many hypotheses concerning the determinants of saving and little empirical evidence on its *true* causes. Indeed debate remains over the most simple theories' assumptions: are consumers forward-looking?[44] are there significant altruistic linkages between generations of families? are liquidity constraints important in determining current

[41] Despite the error terms of the models **(2.3.33)** and **(2.3.35)** theoretically featuring uncorrelated error terms, the actual data used induces autocorrelation due to time aggregation problems. GMM with autocorrelation consistent standard errors is employed with C_{t-2} and A_{t-2} used as instruments for C_{t-1} and A_{t-1}.

[42] Assuming constant rates of return one may view the effective rate at which consumers discount future income as: $[1/(1+\delta)] = [(1-p)/(1+r)]$, see Church *et al* (1994). With p=0, planning horizons are infinite and $[1/(1+\delta)] = [1/(1+r)]$: the discount and real interest rates are equal.

[43] Leung (1994) argues that this may explain the widespread under-saving reported in many national surveys.

[44] Weil (1991) argues that it is unlikely that most agents are forward-looking optimisers although they may make up most of the saving population.

consumption? The length of planning horizon is regarded as an unresolved issue.

2.4 Analysing Consumption in the ADL Error Correction Methodology

This section reviews the literature on the autoregressive distributed lag (ADL) ECMs of DHSY and HUS along with an examination of the break down of these consumption functions in the mid/late 1980s.

2.4.1 The DHSY ECM

The DHSY ECM is one of the most influential econometric specifications of modern times. It encompassed analogues of representative contemporary specifications including, for example, Wall *et al*'s (1975) transfer function and wealth elimination forms of the LCH, equation **(2.2.19)** above. The DHSY model utilises the microeconomic postulates that household consumption is homogenous of degree one in income (unit income elasticity), implying a constant APC, and is homogenous of degree zero in prices.[45] *Imposition* of these homogeneity postulates on an unrestricted ADL in levels yields:[46]

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[45] According to microeconomic homogeneity postulates a rise in prices will not increase *real* consumption, given that *real* income and *real* wealth remain intact, because their *nominal* income and wealth would have risen in the same proportion as the price rise. If increased prices raise consumption this is through *money illusion*: household's mistake equiproportionate increases in *nominal* income and wealth as *real* increases. Branson and Klevorick (1969) find evidence of a statistically significant positive price effect upon USA consumption. However, Burch and Werneke (1975) note that the findings of Branson and Klevorick (1969) are biased towards this finding and question the reliability of Branson and Klevorick's (1969) inferences. The price level is not a typical variable employed in consumption functions.

[46] Assuming a first order lag polynomial in log-linear form the general model is:

$$\ln C_t = K^* + \alpha_1 \ln C_{t-1} + \beta_1 \ln Y_t + \beta_2 \ln Y_{t-1} + \delta_1 \ln P_t + \delta_2 \ln P_{t-1}.$$

Imposing the microeconomic homogeneity postulates that consumption is homogeneous of degree one in income ($\alpha_1+\beta_1+\beta_2=1$; or $\alpha_1=1-\gamma$ and $\beta_2=\beta_1-\gamma$) and that consumption is homogeneous of degree zero in prices ($-\delta_1=\delta_2$) gives the basic form of the DHSY error correction model **(2.4.1)**.

$$\Delta \ln C_t = K^* + \beta_1 \Delta \ln Y_t - \delta_1 \Delta \ln P_t - \gamma_1 \ln(C/Y)_{t-1}. \qquad \textbf{(2.4.1)}$$

The short-run decision, determined by income growth and inflation, is modified by some proportion of last period's (log of the) APC (the error-correction term) to ensure coherence with consumers' long run target consumption-income ratio. The well known static and dynamic equilibrium relations are given as:[47]

$$C/Y = \exp\{K^*/\gamma_1\}, \qquad \textbf{(2.4.2)}$$

$$C/Y = \exp\{[K^*-g(1-\beta_1)-\delta_1\mu_1]/\gamma_1\}, \qquad \textbf{(2.4.3)}$$

where g and μ_1 are the constant long run growth paths of consumption/income and prices, respectively.[48]

DHSY make two important points regarding these equilibriums. First, although the observed downward trend in the UK's APC seems inconsistent with the assumed long-run unit income elasticity, DHSY argue that the target APC need only be constant along *given* growth paths. If there was an upward shift in the growth rate of prices, μ_1, as occurred in the UK in the 1970s, the target APC would fall according to the dynamic equilibrium **(2.4.3)** whilst maintaining the

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[47] The static equilibrium is obtained by assuming that the variables do not change from period to period: $X=X_t=X_{t-1}=\ldots=X_{t-i}$. The dynamic equilibrium is secured by assuming constant long run growth paths.

[48] Currie (1981) argues that static long run solutions to ADL models are sensible and well determined but the dynamic equilibriums often are not. This is because the latter's long run solutions depend upon variables' growth rates as well as their levels. In many instances there may be no theoretical justification for inclusion of such growth rates in the equilibrium solution and so suggests a set of restrictions which may be used to eliminate them. However, in the case of the DHSY consumption function, the growth in prices can be justified due to the mismeasurement of income, the erosion of the real value of wealth and consumers mistaking absolute for relative price increases while income growth's inclusion can be rationalised following Modigliani's (1986) LCH and Brown's habit persistence hypothesis. Thus, the DHSY model may be theoretically sound although Currie (1981) further suggests that it may be difficult to estimate such long run dynamic effects from the relatively short data period employed by DHSY.

unit income elasticity postulate. Second, if g and μ_1 are relatively constant the target APC will be reasonably constant suggesting that the intercept, K^*, and error-correction term, $\ln(C/Y)_{t-1}$, in **(2.4.1)**, will be "almost perfectly collinear".[49] Under such conditions the estimated version of **(2.4.1)** may exhibit excessively large standard errors of the intercept and error-correction terms. Since the static and dynamic long run solutions are not well defined when the error-correction term is excluded but are when the intercept is omitted they argue that the latter can be removed because, in this sense, it has no theoretical role.

DHSY employ both of these arguments in their empirical application of variants of quarterly analogues of **(2.4.1)** to UK data. An estimated version of the general equation in log-level form, which *implies* an error-correction model like **(2.4.1)**, featured a substantially larger intercept standard error than the estimated version of equation **(2.4.1)** (without inflation or an error-correction term). DHSY argue that this large intercept standard error is indicative of collinearity between the implied error-correction term and intercept. They therefore estimate **(2.4.1)** (without price effects) excluding the intercept. This model is found to be data coherent but suffers from predictive failure. This predictive failure motivated the inclusion of price effects to account for the impact of the observed acceleration of inflation in the 1970s upon consumers' behaviour. Two versions of **(2.4.1)** (with price effects) are estimated: one excluding the intercept and one excluding the error-correction term. In both formulations the price effects are found to be significant, however, only the forecasting performance of the former is found to be *acceptable*.[50] This is suggested to imply that it is not the inflation effects by themselves which secure parameter constancy, rather it is their combination with the error-correction term. This is argued to support the view that the upward shift in inflation is lowering the target APC (μ_1 is not constant). Therefore, DHSY have "no hesitation in dropping the constant term instead of the *error correction term*." (DHSY p. 688, my italics), because of their belief in the collinearity of the intercept and error-correction term (μ_1 is constant).

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[49] DHSY implicitly assume that a relatively constant target suggests an equally constant *actual* APC.

[50] Hendry and Mizon (1998) confirm that the addition of inflation to the DHSY model is necessary to remove predictive failure.

However, the favoured DHSY model relies on contradicting assumptions concerning inflation: the former rests on the upward shifting of μ_1 while the latter requires that μ_1 is constant. This implies that either the unit elasticity along given growth paths or the "near perfect collinearity" of the intercept and *actual* APC must be unacceptable. Stewart (1998) argues that the latter assumption is invalid because the variable included in **(2.4.1)**, $\ln(C/Y)_{t-1}$, is clearly variable and cannot be perfectly collinear with the intercept.[51]

There are objections to excluding the intercept. Patterson (1985) suggests that the intercept, K^*, has a theoretical place in **(2.4.1)** because it allows the static equilibrium to deviate from unity. When the intercept is excluded, $K^*=0$, the APC in static equilibrium is unity which "implies not only that there is no saving, but also that there is no expenditure on replacing the depreciating part of the stock of consumers' durables;" (Patterson 1985, p. 471). A negative intercept is required for the APC to be below unity in static equilibrium.[52] Pesaran and Evans (1984) argue that a statistically significant intercept should be included in a regression to avoid biasing its error term. Stewart (1998) notes that excluding a (positive) intercept from DHSY's model will bias the error correction term's coefficient (downwards, making it more negative). This offers an alternative explanation of DHSY's finding that the error correction term is only statistically significant and negative when the constant is excluded. Thus, recent articles which have demonstrated that the favoured DHSY formulation, excluding intercept and modified using time-varying parameters (see, Harvey and Scot, 1994; Song, 1995; and Gausden and Brice, 1995), can successfully model and predict UK consumption over the turbulent period of the late 1980s and early 1990s, may need to be viewed with caution.[53]

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[51] The apparent large standard error of the intercept found in the log-level version of **(2.4.1)** is due to the dependent variable being the log of consumption which features larger units of measurement than models with the growth rate of consumption as the dependent variable.

[52] Nickell (1985) argues that a non-zero intercept may be required in an ECM to track a growing *target* variable. In the DHSY model the target is the APC and one might not expect this to grow (fall) at a *constant* rate indefinitely, although it may over any particular finite sample.

[53] Harvey and Scott (1994) re-estimate the favoured DHSY specification, augmented to allow for time-varying seasonality, using quarterly UK data up to 1992 and find that this modified form does not suffer from predictive failure when the sample is split in several places.

Bollerslev and Hylleberg (1985) empirically examine potential explanations for the decline in the UK's APC through appropriate modification of the DHSY specification. They find that the UK's non-durable APC is not adequately explained by a shift in inflation, or a fall in the relative price of non-durables to durables or being the result of using an inappropriate measure of income. The favoured explanation of the falling APC, at least over their sample, is provided by a declining linear trend. This suggests a below unit income elasticity which needs explanation - it could not follow a linear trend indefinitely because this implies the APC will eventually become negative.

Harnett (1988) reestimates the DHSY model using quarterly current price data for the USA. The estimated model features similar coefficients to those obtained using UK data which is considered surprising because of the dissimilarities in the saving ratios of the UK and USA. Harnett also compares plots of the USA's APC and inflation. Although he accepts that the gradual increase in the USA's inflation since the 1960s may explain the fall in its APC to 1974 he does not accept that the increased volatility in inflation since 1973 explains the rise in the APC from 1974. It is argued that there are some omitted explanatory factors.

Sarantis and Stewart (1998b) demonstrate, using post second world war annual data for Greece, Portugal and Spain, that an unobserved component, in the form of a time-varying trend (see, Harvey, 1989), can be added with statistical significance to a DHSY ECM and captures

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Song (1995) re-estimates the DHSY model allowing the parameters to be a function of their value last period, inflation and wealth, using quarterly UK data up to 1991. The forecasting performance of this time-varying parameter specification of DHSY's model is demonstrated to represent a significant improvement over its fixed parameter counterpart as well as a fixed-parameter version of the HUS model. This is attributed to the ability of the time-varying parameter model to allow consumers to revise their decisions in the face of regime changes - they cite the OPEC oil shocks (inflation) and the financial deregulation (wealth) as the relevant events. Gausden and Brice (1995) re-estimate DHSY's model using quarterly UK data up to 1988. They allow the parameters to be time-varying by specifying additional explanatory variables as the products of the original variables in the model and a time trend. The parsimonious form of this time-varying parameter model is shown to feature greater in-sample fit compared to the original DHSY specification. Additionally, the systematic under-prediction of the 1985-1988 consumer boom exhibited by the fixed parameter model is eliminated with the time-varying coefficient formulation.

nonstationary behaviour. It is argued that use of unobserved components can control for explanatory factors, such as wealth, and help ameliorate misspecification problems associated with omitted variables.

Carruth *et al* (1996) estimate the DHSY model, imposing the unit income elasticity and including intercept, for a panel of fifteen European Union (EU) countries. Although generally well specified, for only eight (six) of the fifteen countries is the error correction term (inflation) negative and statistically significant. They note that the addition of lagged income, relaxing the unit income elasticity, improves the results, yielding statistically significant and correctly signed error correction terms for fourteen of the fifteen countries. They suggest that there is a "weakness of imposing the unit elasticity hypothesis of the DHSY model." (Carruth *et al*, 1996, p. 9).[54]

Pesaran, Shin and Smith (1997) estimate an analogue of the DHSY model, relaxing the unit income elasticity for twenty-four OECD countries and find evidence of valid error correction behaviour for twenty countries. However, seven of the twenty-four countries' regressions are subject to evident misspecification possibly reflecting omitted explanatory factors

2.4.2 HUS, the Integral Control Mechanism and Perceived Income

HUS utilise the following simplified version of the DHSY model to demonstrate why it was unable to provide a *complete* account of the dynamic adjustment of consumption:

$$\Delta \ln C_t = \beta_1 \Delta \ln Y_t - \gamma_1 \ln(C/Y)_{t-1}, \qquad \textbf{(2.4.5)}$$

where the dynamic equilibrium of **(2.4.5)** is:

$$C = \exp\{[-g(1-\beta_1)]/\gamma_1\}Y. \qquad \textbf{(2.4.6)}$$

[54] Carruth *et al* (1996) also find evidence against the hypothesis of a common aggregate consumption function for EU countries suggesting different responses of consumption to shocks across the EU.

They argue that, in response to any shock, the adjustment of consumption to its steady state growth path is monotonic. In the typical case of $\beta_1<1$, consumption grows more slowly than income, so consumption will converge to its equilibrium, **(2.4.6)**, from below (above) when income increases (decreases). The monotonic adjustment process implies that the error correction term is increasing (decreasing) consumption each period, however, is always below (above) its target, giving rise to a cumulative underadjustment. Since consumption is an expenditure and income an accrual, the stock of wealth must be changing to accommodate the cumulative underadjustment. For example, when income is rising, the stock of assets must be increasing because consumption is continually below its target. Expressed in the language of Phillips (1954 and 1957), **(2.4.5)**, incorporates derivative $[\Delta \ln Y_t]$ and proportional $[\ln(C/Y)\ ]_{-1}$control mechanisms but excludes an integral control mechanism (ICM) $[\Sigma \ln(Y/C)_t]$, which is accumulated savings or the integral of past discrepancies between income and consumption. HUS propose extending the DHSY model by including wealth effects.

To incorporate wealth they assume "a prior steady-state utility maximising exercise leads agents to seek to maintain constant ratios between *C/Y* ... and between *A/Y* (*ceteris paribus*), namely: $C^E=K^aY$ and $A^E=B^aY$ where *E* denotes 'dynamic equilibrium'." (HUS, p. 240, my italics). The link between stock and flow variables is given by the following definition of wealth:[55]

$$A_t \equiv A_{t-1} + Y_t - C_t. \qquad \textbf{(2.4.7)}$$

To be consistent with **(2.4.7)** in steady state the following relation must hold (where g denotes income growth):

$$K^a = 1 - \{g/(1+g)\}B^a. \qquad \textbf{(2.4.8)}$$

The logarithmic steady-state approximation of **(2.4.7)** is:

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[55] Comparing **(2.4.7)** with the budget constraint **(2.3.32)** shows that the term rA_{t-1} is included in the latter. We note this without using notation to distinguish disposable income and labour income.

$$\Delta \ln A_t^E \equiv H^a(\ln Y_t - \ln C_t^E), \quad \text{where,} \qquad H^a = (1+g)/B^a, \tag{2.4.9}$$

and the logarithmic long run targets are:

$$\ln C_t^E = \ln K^a + \ln Y_t, \quad \text{and,} \qquad \ln A_t^E = \ln B^a + \ln Y_t. \tag{2.4.10}$$

Recognising that these targets are not always achieved they set up the following one period quadratic loss function to assign priorities to removing disequilibria:

$$q_t = \lambda_1(\ln A_t^P - \ln Y_t - \ln B^a)^2 + \lambda_2(\ln C_t^P - \ln Y_t - \ln K^a)^2 + \lambda_3(\ln C_t^P - \ln C_{t-1})^2 - 2\lambda_4(\ln C_t^P - \ln C_{t-1})(\ln Y_t - \ln Y_{t-1}). \tag{2.4.11}$$

The first two terms penalise deviations of *planned* values (denoted with a P superscript) from their respective steady state outcomes. The third term seeks to stabilise behaviour in a non-growing world by attaching a cost to the change in planned consumption this period and its actual value last period. These three terms are squared to penalise large deviations more than small ones. The fourth term is introduced to temper the third term by subtracting losses when there is growth in the economy, allowing planned consumption to grow. The λ_is are constant parameters.

Minimising the expected value of **(2.4.11)** with respect to $\ln C_t^P$ taking account of **(2.4.9)** holding for *planned* quantities yields the basic HUS model:

$$\Delta \ln C_t = \theta_0 + \theta_1 \Delta \ln Y_t - \theta_2 \ln(C/Y)_{t-1} + \theta_3 \ln(A/Y)_{t-1} + u_t. \tag{2.4.12}$$

This is the DHSY model with the ICM, $\ln(A/Y)_{t-1}$, replacing inflation. The dynamic equilibrium is obtained by assuming, $g = \Delta \ln C_t = \Delta \ln Y_t = \Delta \ln A_t$, which when substituted into **(2.4.12)** yields:[56]

$$C/Y = [\exp\{[\theta_0 - (1-\theta_1)g]/\theta_2\}](A/Y)^{(\theta_3/\theta_2)}. \tag{2.4.13}$$

[56] The logarithmic form of the solution is: $\ln(C/Y) = \{[\theta_0 - (1-\theta_1)g]/\theta_2\} + \theta_3 \ln(A/Y)^{(\theta_3/\theta_2)}$.

HUS demonstrates that in non-stochastic steady state **(2.4.12)**, solved for planned values, essentially yield the desired equilibrium consumption-income and wealth-income relationships given by **(2.4.10)**. Their connection is given by **(2.4.13)**.[57]

The second substantial innovation of HUS was to adjust the conventional measure of disposable income for inflation induced losses on (liquid) assets. It is argued that during inflationary periods the real value of assets declines, causing a rise in its rate of return to compensate for this loss. Since the conventional measure of personal disposable income includes interest earnings but excludes changes in the real value of assets, an upward shift in inflation will cause this conventional measure to artificially rise. That is, accelerating inflation causes conventionally measured personal disposable income to rise due to increased interest receipts without any corresponding reduction due to inflation induced losses on assets. Therefore, HUS propose adjusting conventionally measured disposable income by subtracting inflation induced losses on assets. This yields a measure of a person's income which is more consonant with Hick's widely accepted definition being: "what he can consume during a week and still be as well off at the end of the week as he was at the beginning" (quoted in Ungern-Sternberg, 1986, pp.741-742). The adjusted measure (*perceived* income), Y_t^*, is calculated using the formula:

$$Y_t^* = Y_t - \rho E(\Delta \ln P_t) A_{t-1}, \qquad \textbf{(2.4.14)}$$

where the expected rate of inflation, $E(\Delta \ln P_t)$, is used because consumers' expenditure decisions are suggested to be based upon their *perceived* real income rather than its actual value. The proportionality coefficient, ρ, should equal unity in the absence of scaling errors due to the mismeasurement of $E(\Delta \ln P_t)$ or A_{t-1}.

HUS's estimate of ρ (=0.5) is found to be significantly different from zero, supporting the need to adjust income, however, it is also significantly less than one, suggesting some possible mismeasurement of $E(\Delta \ln P_t)$ or A_{t-1}. The appropriate measurement of the latter has received

[57] Salmon (1982) argues that the inclusion of an ICM ensures a zero steady state error whereas the DHSY model, with its constant target, was subject to a constant equilibrium error.

much attention in the literature. HUS use the stock of net liquid assets, L_t, to proxy total wealth, probably due to data constraints, and argue that L_t will exhibit similar variations to total wealth, A_t, because they are the most spendable (variable) component of wealth. Ungern-Sternberg (1981) employs the narrower wealth concept of monetary assets (liquid assets less building society and mortgage loans) because it appropriately represents an erosion of income. Ungern-Sternberg (1981) find that ρ is significantly different from zero and insignificantly different from unity for both West Germany (ρ=1.16) and the UK (ρ=0.85), suggesting that the narrower concept of monetary assets is appropriate.

Patterson (1984 and 1985) suggest that the definition of perceived income, **(7.2.19)**, clearly indicates that components of wealth beyond liquid assets should be considered in the adjustment. Steel (1987) finds evidence favouring the adjustment of Belgian income with accumulated saving where ρ=0.96. This estimate is close to unity, possibly indicating the appropriateness of broader wealth measures for adjusting income. In contrast, Patterson (1991b) considers inflation losses on five components of wealth and favours restricting the adjustment to liquid assets. However, Carruth and Henley (1992) find a significant role for housing equity withdrawal in the adjustment of income when building a model of UK consumer durable expenditures. It appears that disposable income needs to be adjusted for inflationary losses on assets although the appropriate definition of wealth remains an unresolved issue.

HUS suggest that using perceived income may render the separate inclusion of inflation terms unnecessary. "It should be stressed that the use of Y_t^* is in principle complementary to the theory in Deaton (1977), although in practice the explanations are likely to be more nearly substitutes." (HUS, p. 248). They find lagged price variables to be insignificant in an unrestricted log-level HUS formulation when perceived income is employed, confirming the redundancy of additional inflation effects. Rossi and Schiantarrelli (1982), using a HUS formulation for Italy, are also unable to find evidence supporting an additional role for inflation and, indeed, interest rates or relative price variables. Patterson (1985) finds no role for inflation beyond the adjustment of income in a DHSY consumption function of the UK. Muellbauer and Murphy (1989) find that adding variables, including liquid and illiquid assets, to an annual analogue of the DHSY model for the UK yields similar coefficients to DHSY's favoured specification on all except the inflation terms. They argue that "this result also provides a valuable insight into the role of inflation effects in consumption functions: they appear to be primarily an imperfect proxy for real asset and debt effects." (Muellbauer and Murphy, 1989, p. 62). Molana (1989), commenting on the evidence of HUS, argues "that inflation is relevant only so far as it devalues the potential purchasing power of households, there is no direct relation between consumption and fluctuations in the price level. Therefore this finding undermines the models presented by Deaton and Davidson *et al.*" (Molana, 1989, p. 213). He goes on to develop a model which adjusts income for capital gains on various assets suggesting that one should not restrict adjustment to the income variable but incorporate these capital gains terms as separate regressors. "In our study we shall use the latter approach. This is because we believe that a correct adjustment of income cannot be obtained and even if one could obtain such a measure there would be doubts whether capital gains and/or losses should be restricted to have the same effect as income." (Molana, 1989, p. 216).

HUS's model applied to UK data is free from evident misspecification and variance dominates an analogue of DHSY's model so is regarded as the preferred specification. Similarly, Davidson and Hendry (1981) and Patterson (1985) find that the HUS model variance dominates the DHSY specification using quarterly UK data. Further, Davidson and Hendry (1981) find that both the DHSY and HUS specifications variance dominate an analogue of Hall's (1978) model[58] applied to UK data.[59]

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[58] Davidson and Hendry (1981) argue that although ECMs do not employ expectational hypotheses they can frequently mimic rational behaviour. Indeed, Nickell (1985) shows that ECM's provide optimal responses of agents in a dynamic environment for a variety of different circumstances. Since "feedback and forward-looking behaviour can 'look-alike' in many states of nature... the problem is not one of reconciling error correction or expectational interpretations, but of distinguishing their separate influences." (Davidson and Hendry, 1981, p. 191).

[59] Bean (1986) finds that Muellbauer's (1983) extended version of Hall's (1978) REPIH/RELCH, modified for surprises in real interest rates, the change in income, hours worked and government expenditure, "just" encompasses an analogue of the DHSY model applied to quarterly USA data. In contrast, Harnett (1988) finds that the DHSY model variance dominates a REPIH/RELCH specification for the USA. Whether ECMs encompass REPIH/RELCH models for the USA is unresolved.

Hendry (1983a) subjects the DHSY and HUS models to six criteria for rigorously assessing models using both annual and quarterly UK data.[60] The estimated analogue of the DHSY model is found to satisfy *data coherency*, because it exhibits Gaussian residuals with a 0.5% standard error, *valid conditioning* (weak exogeneity) because OLS and IV estimates are almost identical, *parameter constancy* (no structural instability),[61] *data admissibility*, since a savings analogue of the DHSY model provides similar inference to its consumption counterpart and *theory consistency*, because the DHSY model exhibits theoretically expected signs and satisfy the unit income elasticity postulate.[62] However, although the DHSY model *encompasses* many previous researchers' findings it is, itself, both nested[63] and variance dominated by the HUS formulation.[64]

Ungern-Sternberg (1981) considers the performance of the HUS formulation, using monetary assets to proxy A_t, for West Germany, the UK and the USA. A well specified model is found for both West Germany and the UK but not the USA. For both West Germany and the UK there is evidence that the role of assets is to adjust for the mismeasurement of the conventional measure of disposable income (negative income effect). However, only in the UK is there evidence of a statistically significant ICM, which is argued to reflect the personal sector's attempts to rebuild their asset positions (real balance effect). The combination of these effects is argued to explain

[60] An interesting discussion of the Hendry methodology, focusing on the pursuit of *true* models, the need for enduring models (parameter stability) and the role of judgement versus data evidence is provided by Lawson (1981 and 1983) and Hendry (1983b).

[61] It is argued that the DHSY model appears not to be subject to the Lucas critique because of its ability to "mimic" forward looking behaviour.

[62] Hendry (1983a) does, however, warn against imposing theoretical models upon data. For example, it may be interpretation rather than the sign of a coefficient which is incorrect.

[63] It should be noted that "a nesting model which formed the union of all other hypotheses would automatically, but rather vacuously, encompass so a parsimony criterion remains pertinent." (Hendry, 1983a, p. 215).

[64] Pesaran and Evans (1984) find that a quasi-differenced LCH model modified for capital gains on ordinary shares encompasses analogues of the Deaton (1977), DHSY, HUS and Muellbauer (1983) specifications, featuring the largest maximised value of the log-likelihood function. Hendry (1983) notes that this is for total UK expenditures and, given that he finds the capital gains term cannot be added to the HUS model for non-durables, suggests that capital gains primarily influence durable consumption.

the fall in the UK's actual APC, however, the former provides the sole explanation for the trend in West Germany's APC. Ungern-Sternberg (1986) confirm these results using substantially revised German data (the model is robust through time).

Rossi and Schiantarelli (1982) find that the HUS model adequately characterises the fall in Italy's non-durable APC since the 1960s. The favoured results employ financial assets as the appropriate definition of wealth.[65] Although the ICM remains significant when total net wealth was used, the performance of the model deteriorates. It is argued that, measurement problems aside, this suggests that the narrower definition of wealth is the relevant concept, possibly due to its greater degree of liquidity.

In contrast, Patterson (1984) argues that the ICM theoretically refers to total net wealth - the integral of past discrepancies between consumption and income. Currie, Holly and Scott (1990) argue that it is unrealistic to assume capital markets are so imperfect to completely prevent consumption out of illiquid assets. Steel (1987) finds that an ICM based upon accumulated savings enters a well specified Belgian non-durable consumption function significantly, if with a negative sign.[66] It is argued that this supports the wider definition of wealth. Similarly, Patterson (1984) secures a well specified HUS formulation for the UK incorporating two ICMs, the liquid asset to income ratio and an illiquid asset to income ratio, which variance dominates a model employing the total asset-income ratio. The liquid asset to income ratio exhibits a significantly larger elasticity than the illiquid asset variable, reflecting the easier spendability of the former.[67]

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[65] Financial assets are defined as monetary assets plus privately held bonds and equities.

[66] This apparently counterintuitive sign is rationalised as follows. High accumulated savings will facilitate greater access to loans when credit markets are imperfect. If such loans are for durables they may require some additional funds which may be obtained by substituting expenditures away from non-durables.

[67] Harnett (1988) finds that the liquid asset to income ratio is statistically significant while the tangible asset to income ratio is insignificant in an ECM for the UK, confirming the different spendability of assets and suggesting that tangible assets cannot be (easily) converted into consumption.

Molana (1989) argues that the increased volatility of UK financial asset prices from the mid-1960s to mid-1970s led households to substitute them with physical assets, which are suggested to maintain their value, reflecting risk aversion of UK consumers. With the subsequent reduction in volatility this portfolio reallocation has abated. This indicates that households adjust their portfolios towards some desired target share of assets. If this is the case, Molana argues that the long run aggregate consumption function will need to include end of period wealth and the asset's price, Q_t, relative to consumer prices. The following long run consumption function is suggested:

$$C = Y^{\theta_4} A^{\theta_5} P^{\theta_6} Q^{\theta_7} \exp\{\theta_8\}. \qquad \textbf{(2.4.15)}$$

Two prior beliefs regarding the long run elasticities, θ_i, are suggested. First, the "extreme" assumption of the unit income elasticity is relaxed by postulating that consumption is homogenous of degree one in *lifetime resources*: $\theta_4+\theta_5=1$, where, $\theta_4<\theta_5<1$. Second, consumption is homogenous of degree zero in prices, $\theta_6+\theta_7=0$: absolute prices have no long run impact upon consumption although relative prices do. Imposing these restrictions, dividing both sides by A, then by (A/Y) and rearranging gives the HUS formulation extended for relative prices:

$$(C/Y) = (A/Y)^{(1-\theta_4)}(Q/P)^{\theta_6}\exp\{\theta_8\}. \qquad \textbf{(2.4.16)}$$

Using quarterly UK data Molana estimates a general dynamic log-levels formulation of **(2.4.16)** and finds that assets and asset prices feature joint statistical significance suggesting an improvement over the DHSY and HUS formulations. Regarding the latter, the role of bond prices and use of total rather than liquid assets is emphasised. However, both postulated homogeneity restrictions are rejected, possibly suggesting that consumption and wealth income ratios should not be imposed and that there is long term money illusion.[68]

[68] Patterson (1991b) rejects the restriction that consumption is homogenous of degree one in income and wealth at the 5%, but not 1%, level.

Harnett (1988) estimates a model of the USA's non-durable current price consumption which includes the lagged values of financial (negatively) and tangible (positively) assets as well as the lagged ratio of household liabilities to financial (positively) assets (the sign of correlation is given in brackets).[69] Although the asset variables do not enter as ratios to income, as is typical of ICMs, they are lagged which is consistent with them depicting long run behaviour.[70]

Davis (1984) notes that all the major UK macroeconomic modelling agencies of the early 1980s used variants of ECMs, suggesting their superior performance for the purpose of macroeconomic modelling. It is also argued that allowance for inflation effects through the accommodation of losses on liquid assets and the use of ICMs appears to be superior to the simple incorporation of the (difference) of the log of prices.

2.4.3 Explaining the 1980s Breakdown of ADL ECM Consumption Functions

These previously well specified ADL ECMs broke down when samples were extended beyond the mid-1980s, especially in the UK, as policy in developing countries switched towards the liberalisation of markets. For example, Carruth and Henley (1990) observe that all consumption functions adopted by the UK's major macroeconomic modelling agencies systematically underpredicted the rapid growth in UK consumption after 1985.

Muellbauer and Murphy (1989) - MM hereafter - identify eight potential explanations for this breakdown. First, *increased equity withdrawal*, facilitated by the easing of credit restrictions, allowed previously frustrated consumption plans to be realised: the MPC out of illiquid assets

[69] Inflation is significant in the model, however, because income is not adjusted for inflation induced losses on assets, it may be picking up this effect.

[70] Harnett highlights the positive influence of the liabilities to financial asset ratio as of particular interest. This ratio rose dramatically between 1973 and 1982 suggesting that Americans' response to high inflation was to adjust their portfolios to take advantage of the cheaper real cost of borrowing (with sometimes negative real rates of interest), to realise assets whose real values may be eroding and to increase consumption. This is argued to contrast with UK consumers' response to inflation: to maintain the real values of asset ratios which are held as a precaution against future uncertainty.

increased. "It is an empirical question whether this increased fungibility applied only to owner occupied housing wealth or as well as other assets such as pension rights and life insurance savings." (MM, p. 40).[71] Second, the *increased value of wealth* raised expenditures. The end of credit rationing led to an increased effective demand for housing which would persist each period with the continual entry of first time buyers. With an inelastic supply of housing in the short term, due to lags in the response of the construction industry, increased loans for house purchase would lead to an insatiable excess demand for housing each period, raising property prices, homeowners wealth and, therefore, consumption - UK house prices soared in the mid/late. Third, *reduced economic uncertainty* raised consumption. After the turbulent 1970s with oil shocks and union activity the 1980s were argued to be a period of steadier and longer expansion of real income. Reduced income uncertainty lowers precautionary saving and raises consumption. Fourth, *demographic change*, especially the increase in the relatively high consuming young following from the 1960s baby boom, raised expenditure.[72] Fifth, the manifesto *promising tax cuts* may, if credible, raise consumption by increasing the expected future disposable income stream. Sixth, is *cuts in publicly supplied substitutes for private consumption*. For example, cuts in public health provision and higher education maintenance grants may increase private expenditures on health (insurance) and subsistence during college. Seventh, the *declines in 'non-discretionary' or 'institutional' saving* "as life-insurance companies have reduced their contribution rates in response to the increases in financial asset values relative to the incomes being guaranteed." (MM, p. 30). Eighth, is *errors in the national accounts data* on consumption and income. Only the first four explanations are empirically considered by MM.

MM estimate a HUS style ECM using annual UK data. Their specification includes separate

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[71] Miles (1992) argues that the most significant effects of equity withdrawal were related to the housing market for three reasons. First, loans for house purchase represent the largest flow of gross credit to the personal sector. Second, building societies were allowed to deviate from previously government dictated interest rates facilitating the removal of credit rationing and mortgage queues. Third, the type of loans institutions could offer changed with second mortgages and loans unrelated to home improvements proliferating on an enormous scale. Owner occupiers no longer needed to move house to release equity for consumption purposes.

[72] Lee and Robinson (1989) outline an alternate mechanism through which the population's age structure influences the consumption of the elderly by encouraging the trade down of housing to release equity for consumption.

financial and physical asset ICMs and features a proportion of constrained consumers which varies with financial deregulation.[73] The coefficient on illiquid assets is also allowed to change after 1981 to test for equity withdrawal effects. MM find that credit liberalisation has increased UK consumption through increased equity withdrawal, an increase in the proportion of the population who are young and increased income uncertainty. However, it is unclear whether the proportion of constrained consumers falls.[74]

Miles (1992) reports regression results which show that when an equity withrawal term is added to the Pesaran and Evans (1984) UK savings function, using data extended to 1988, it is negative and statistically significant, substantially improves fit and almost completely removes structural instability which is evident without this term. He argues that the majority of equity withdrawal (80%) was spent on consumption rather than financial assets or paying off debt. Bayoumi (1993) estimates a saving function using a panel of UK data for the eleven standard regions and finds that the main impact of deregulation was through increased real wealth which reduced saving by over 5%.

Carruth and Henley (1990) estimate two HUS-style ECMs using UK data. In the first the ICM is defined as the ratio of financial assets to income while the second redefines the wealth variable as financial assets plus the housing stock. The first model provides unsatisfactory forecasting performance while the second features good predictive capabilities. Since the inclusion of housing wealth in the second specification is the main reason for the improved forecasting performance this supports the hypothesis that previously illiquid assets became more liquid.

Patterson (1991b) extends the HUS formulation using five ICMs to allow for the different spendabilities of net liquid assets (NLA), net other financial assets (OFA), equity in life

[73] This time varying proportion is: $\pi_t = \pi_0 - \delta_3 FLIB_t$, where π_0 and δ_3 are constant parameters and $FLIB_t$ is a financial deregulation dummy variable.

[74] It has been argued that "In Muellbauer and Murphy (1989), we attempted to estimate both shifts in π and changes in the spendability of illiquid assets. But this proved very difficult and, in any case, our 1989 model was lacking in theoretical foundations and suffered from considerable overfitting." (Muellbauer and Murphy, 1994, p. 16).

assurance and pension funds (LAPFE), net housing wealth (NHW), and the stock of consumer durables (SCD). This model is estimated using quarterly UK data. The favoured formulation includes the OFA, LAPFE and SCD asset variables. The statistical insignificance of liquid assets is argued to be due to financial deregulation because such assets are particularly important for constrained consumers and, with the reduction in liquidity constraints, their importance relative to other assets declines. However, that there is no role for housing wealth appears to indicate this asset has not become more liquid.

Carruth and Henley (1992) empirically consider the extent to which durables are affected by financial deregulation via the house price boom of the late 1980s and early 1990s collapse. They argue that changes in consumption will likely be concentrated upon durables, being relative luxury items and through complementarity between expenditure for durables and housing - a new house needs furnishing. Applying a HUS-style model to quarterly UK data they find an important role for housing equity withdrawal, working through an income effect, and a direct effect of house prices leading to complementary durable spending.

Using UK micro data Attanasio and Weber (1994) find that increased house prices can only explain part of the 1980s consumer boom. In particular, it cannot explain the rise in the young's consumption. Indeed, increased house prices may reduce the current consumption of the young who are saving for future house purchase. They find that increased income expectations due to the perceived productivity miracle of the 1980s are likely responsible for the majority of the consumer boom.

Lehmussaari (1990) argues that financial deregulation in the Nordic countries (Denmark, Finland, Norway and Sweden) allowed previously unfulfilled demand for household credit to be increasingly satisfied.[75] He utilises a HUS-style ECM modified to include short-run wealth effects and inflation.[76] All countries' models use annual data and are estimated from 1971 (or

[75] Berg (1994) outlines the process of deregulation in the Nordic countries which began in the early-mid 1980s.

[76] For Denmark total wealth (financial and housing) was employed, the housing price index was used for Finland and Norway with net financial wealth being used for Sweden.

1972) to 1987. The following favoured models were chosen. The DHSY ECM including short-run wealth effects and excluding ICM and inflation effects, for Denmark and Norway; the DHSY ECM (without ICM) with short-run inflation, but not wealth, effects for Sweden; while for Finland, a HUS style model (including both proportional and integral control mechanisms) with short run wealth, but not inflation, effects. Structural change is evident for Denmark, Finland and Norway. The elasticity of consumption to wealth showed a dramatic rise for Norway, and notable, but more modest increases for both Denmark and Finland. There is no evidence of structural change following deregulation in Sweden, possibly due to the "grey" market that had already developed in the second half of the 1970s. Deregulation is suggested to have facilitated increased access to credit markets raising consumption both directly and indirectly through increased wealth for all the Nordic countries except Sweden.[77]

Koskela *et al* (1992) confirm that the 1972 and 1987 house price booms reduce the Finnish saving ratio using a modified Deaton (1977) model with quarterly data over the period 1970-1989.

Muellbauer (1994) suggests that the phasing out of ceilings on paid interest loans after the mid-1970s introduced deregulation into the USA. Bovenberg and Evans (1990) note the substantial decline of the US *national* saving rate during the 1980s which is due to declines in both *public* and *private* saving; the latter primarily attributable to reduced *personal* saving. Using an ECM and simulations they find that the decline in the USA's aggregate *personal* saving rate since 1980 is due to increased wealth, falling inflation and demographic changes. The sharp increase in share and house prices, raising wealth, are suggested to be related to financial deregulation. Increased real after tax interest rates are found to have a significant impact in moderating this

[77] Berg (1994) provides a sobering summary of the impact of policy on the Nordic countries. Financial deregulation increased credit availability during the 1980s, increasing demand and raising asset prices. In the wake of this liberalisation (the late 1980s and early 1990s), came tax reforms which encouraged saving in financial assets and debt repayment so lowering demand for tangible assets and thus asset prices. The bad timing of these policies left households with high debts facing deteriorating asset prices (wealth). The subsequent debt-deflation probably deepened and lengthened the recession. "The economic and social costs of these policy mistakes have, of course, been very high, as they were in the UK where similar policy mistakes occurred." (Berg, 1994, p. 52).

fall in the saving rate, suggesting that after liberalisation households were more able to substitute consumption through time. Although demographic effects are found to be very important it is recognised that they may be picking up unmodelled trend-like factors.

2.5 Analysing Consumption in the Cointegration-Error Correction Methodology

Engle and Granger (1987) introduced a means of estimating and testing for cointegrating (equilibrium) relationships and, if they exist, representing them as an ECM. The Johansen (1988) method was developed to overcome the shortcomings of the Engle and Granger (1987) procedure. (Details of these methods and their relative merits is provided in Chapters 5 and 6).

Drobny and Hall (1989), using quarterly UK data, test ten logarithmic variants of DHSY's dynamic equilibrium, **(2.4.3)**, which are nested in **(2.5.1)**, for cointegration using the Engle and Granger (1987) procedure:

$$\ln C_t = b_{25} + b_{26}\ln Y_t - b_{27}\Delta\ln P_t - b_{28}\Delta\ln Y_t. \qquad \textbf{(2.5.1)}$$

Regardless of specification cointegration is always rejected indicating omitted factors in the long run consumption function, **(2.5.1)**.[78] Following Borooah and Sharpe (1986), Drobny and Hall (1989) consider whether the increase in income inequality, expected after the Conservative government took office in 1979, reduced consumption, as the lower MPC rich became more important in aggregate behaviour. Income distribution effects are modelled through the changing structure of income tax. Assuming higher rate tax payers have relatively low MPCs, a reduction in the tax rate differential, TAX_t, as occurred in the UK during the 1980s, would reduce aggregate consumption. Possible liquidity constraints from the housing market are captured by the quarterly mortgage rate, RM_t, and, following HUS, they include real financial wealth in the

[78] They find that this inference is similar to long run relations implied by estimated variants of DHSY's ADL model **(2.4.1)**, which are found to systematically overpredict consumption in the early/mid 1980s. This is interpreted to indicate a relationship between the ADL and cointegration ECM methodologies. In contrast, Larsson *et al* (1998) find evidence favouring cointegration for **(2.5.1)** for twenty-three OECD countries using a panel cointegration test.

following test equation:[79]

$$\ln C_t = b_{29} + b_{30}\ln Y_t + b_{31}\ln(A/Y)_t + b_{32}TAX_t + b_{33}RM_t. \quad \textbf{(2.5.2)}$$

The TAX_t variable is found to be crucial to securing cointegration and the favoured model includes all variables in **(2.5.2)**. A valid ECM using this favoured cointegrating vector is developed thus Drobny and Hall (1989) find support for a HUS-style relationship extended to account for income inequality.[80]

Molana (1991) develops a modified LCH theory for the "typical macroeconomic agent", rather than representative household, to be more appropriate for aggregate data. The utility function of the "typical macroeconomic agent" is argued to depend upon consumption *and* wealth, the latter is justified when there is uncertainty over future income or when liquidity constraints prevail. The resulting first order condition is shown to include wealth rather than income. "The empirical implications of the above analysis are straightforward. First, the dependence of the rate of growth of consumption on wealth may explain the empirical failure of the simple *REPIH/RELCH* model. Second, if an *error correction* model of consumption is to be constructed on the basis of life-cycle theory, the cointegration analysis will have to be centred on the relationship between consumption and wealth rather than income." (Molana, 1991, p. 388). Using quarterly UK data Molana (1991) finds a stationary linear combination between the logs of consumption and wealth but not the logs of consumption and income. Correspondingly, a valid ECM representation is only obtained when the cointegrating relation between consumption and assets is used suggesting that consumption and wealth, rather than consumption and income, constitute an equilibrium. This is argued to explain the parameter instability found in ECMs which concentrate on a long run relation between consumption and income. It is recognised that although income may not, on its own, explain long run consumption it may form part of an

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[79] A dummy variable is incorporated to capture pre-announced VAT changes.

[80] Hall (1991) repeats the work of Drobny and Hall (1989) using the Johansen procedure and confirms support for cointegration between UK consumption, income, the wealth-income ratio and the tax differential.

equilibrium relationship *with* wealth.[81]

Currie, Holly and Scott (1990) use quarterly UK data to examine the consumer boom of the 1980s. When only consumption, income and liquid assets are considered, cointegration is rejected. This contrasts with their favoured cointegrating relation which includes consumption, income, liquid assets, net housing wealth and illiquid financial assets (their coefficients decline in this order reflecting these variables different degrees of spendability). This suggests the inadequacy of liquid assets on their own. Their evidence fails to support the coefficients on housing wealth and illiquid assets varing with MM's $FLIB_t$ variable so they reject a change in the fungibility of these assets. A well specified ECM is constructed which features demographic and real interest rate effects determining short run behaviour. They do not find that consumption has become more sensitive to interest rates in the deregulation era. Their model suggests that the rise in consumption is primarily explained by increases in the following factors: income (44%), liquid assets (20%), illiquid financial assets (17%) and housing assets (8%). Financial liberalisation is argued to have worked through increases in illiquid assets where equity withdrawal facilitated this increase. The rise in wealth, due to deregulation, is suggested to be the primary cause of the 1980s consumer boom.

Hall and Patterson (1992) apply Johansen's procedure to Patterson's (1991a) model using the quarterly UK data employed by Patterson (1991b). Patterson (1991a) extends the HUS formulations of Patterson (1984 and 1985) to incorporate all of the disaggregated components of total wealth, in a complete integrated simultaneous equations system of consumption and portfolio decisions where the target asset-income ratios are allowed to vary with their rates of return. This yields a system where the (log of the) consumption target depends upon (the log of) income, income growth and the changes in the rates of returns on the various assets and the (log of) asset targets depend upon (the log of) income and the rates of return on each asset. They find evidence for eight cointegrating relations and that all the wealth income ratios feature the correct positive sign, except for other financial assets, in the equation of interest (normalised on the

[81] This evidence also supports Molana's view that consumers derive utility from wealth so explaining the empirical failure of Hall's (1978) REPIH/RELCH formulation.

APC). They proceed to develop a six equation dynamic system, based upon the consumption and asset income ratios' growth rates, and find that each feature negative and significant error correction terms. Thus, all the asset-income ratios influence the long run evolution of the change in the APC. It is argued that "These results can hopefully be improved upon - they represent what is considered as the much needed start on research on a systems view of consumption and wealth allocation. It seems unlikely that given the developments, particularly in financial markets in the 1980s, that it will be sufficient to consider households decisions on expenditure on non-durables and services separately from decisions on the accumulation of net assets." (Hall and Patterson, 1992, p. 1169).

Holmes (1993) considers whether incorporating housing equity withdrawal into an ECM of the APC can resurrect UK consumption functions. In contrast to Carruth and Henley (1992) who adjust income with equity withdrawal, Holmes includes it directly as a separate variable. He extends Arestis and Hadjimatheou's (1982b) model by incorporating real equity withdrawal (REQW) into the PIH as follows:

$$C_t = h[r_t, (A/Y)_t, REQW_t, P_t/P_{t-1}].Y_t^P. \qquad \textbf{(2.5.3)}$$

Dividing **(2.5.3)** by Y_t and taking the log of both sides gives:

$$\ln(C/Y)_t = b_{34} - b_{35}r_t + b_{36}\ln(A/Y)_t + b_{37} \ln REQW_t \pm b_{38}\Delta \ln P_t + b_{39} \ln(Y^T/Y)_t, \qquad \textbf{(2.5.4)}$$

where $Y_t^P = mY_t^T$, and Y_t^T is a weighted average of past and present observations on the trend value of income. Holmes (1993) applies this model to both durable and non-durable UK expenditures using the Johansen procedure and finds evidence of five cointegrating relations for both consumption measures. Holmes identifies the favoured cointegrating vector for both durable and non-durable expenditure APCs as those which are consistent with the theoretical priors indicated by **(2.5.4)**. For both models all the coefficients are statistically significant, including REQW, and feature expected signs. The coefficients for non-durable expenditures are smaller in magnitude than those for durables reflecting the greater sensitivity of the latter to explanatory variables. In particular, Holmes notes that the smaller coefficients in the non-durable equation on $\ln(A/Y)_t$ and $\ln REQW_t$ reflects a greater propensity for households to use equity withdrawal to finance durable rather than non-durable expenditure, which is consistent with Carruth and Henley's (1992) findings. The larger coefficient (in magnitude) on interest rates in the equation for durables is argued to reflect the greater importance of borrowing for durable expenditures. Well specified ECMs are developed for both expenditure types. The coefficients on the disequilibrium terms for both models are not statistically different, suggesting similar speeds of adjustment.

Church *et al* (1994) note that the consumption functions which failed to forecast the 1980s UK boom when respecified to accommodate financial deregulation (especially illiquid wealth and possibly equity withdrawal) and shifting expectations, were able to account for this boom. Using eight ECMs, five from large-scale macroeconomic models and three from City institutions, they consider whether the factors that explain the 1980s boom can also characterise the subsequent slump of the 1990s. Using cointegration tests, including both the Engle and Granger (1987) and Johansen (1988) procedures, they find some support for the existence of a stable single long-run relationship between *total* consumption expenditure, disposable income and total housing and financial wealth. Their evidence rejects the disaggregation of consumption or wealth in the long run relationship. Dynamic ECMs for total consumption are found to be less prone to predictive failure towards the end of the 1980s than formulations explaining non-durable expenditures. However, all eight models' forecasts systematically overpredict expenditure in 1989-92 with the City models explaining the 1990s downturn better than the large-scale macroeconomic models. They suggest that this *may* be due to their use of more recent information, being more parsimonious and including unemployment (in the cointegrating relation), possibly capturing income uncertainty effects. The level or difference of the unemployment rate is found to yield better results than transformations based upon its log. Church *et al* (1994) conclude the following. First, the use of total expenditure, if a less pure definition of consumption, is preferable to non-durables plus the imputed services from durables as the latter is very difficult to measure and, it is suggested, the costs of attempting to do so appear to outweigh the potential benefits. Second, most modern consumption functions are based upon the LCH and increasingly focus upon wealth with greater support for cointegration being secured between consumption, income and broader definitions of wealth. Third, none of the models successfully predicted the

1990s downturn. Adding variables in the long run relation appears not to have been fruitful, with the possible exception of unemployment. Greater success has been achieved by augmenting the short-run dynamics, although a *magic* variable has yet to be found.

Church and Curram (1996) compare four of the ECM consumption functions of the UK considered in Church *et al* (1994) with neural networks based upon the same variables and they find that they feature similar modelling and forecasting performances. It is concluded that there are no major non-linearities in consumption and that "the role of judgement in choosing the appropriate explanatory variables is the most important factor." (Church and Curram, 1996, p. 266). This highlights the need to find explanatory factors to characterise the 1990s downturn.

Horioka (1996) finds, using the Johansen procedure, a unique cointegrating vector between total private consumption, disposable income (or labour income) and total wealth using annual Japanese data.[82] His model indicates that the massive capital gains occurring between 1986-1989 are responsible for 24%-68.7% of the increase in consumption during that period. Similarly, the capital losses endured between 1990 and 1992 are estimated to have depressed consumption by 1.3%-3.5% (percentage points) during 1990-1993. This supports the role of broadly defined wealth. Consideration of additional explanatory factors of Japanese consumption is recommended. Horioka (1997) examines one such variable by considering whether the Japanese household saving rate cointegrates with the age structure of the population, specifically the dependency ratio (the proportion of those aged nineteen and under in the working population) and the retired proportion of the (working) population. Using annual per-capita data he finds, according to the Johansen procedure, evidence for cointegration. OLS, Johansen and Stock and Watson (1993) estimates of the cointegrating vectors reveal that both age structure variables exhibit the expected negative signs and are generally statistically significant. This evidence is argued to support the LCH and suggests that the high Japanese saving rate is, at least in part, due to the relatively young age structure of the present population.

[82] Horioka (1996) argues that the use of current income is especially appropriate for Japan citing evidence indicating that a particularly substantial proportion of Japanese households are subject to liquidity constraints.

Malley and Moutos (1996) consider the role of unemployment, proxying income uncertainty, on quarterly USA motor vehicle expenditure. It is argued that transactions costs, indivisibilities and the problem of *lemons* are associated with durable expenditures to a greater extent than non-durables suggesting larger expected costs of committing to a durable purchase. Hence the influence of income uncertainty will be more pronounced on durables and motivates their analysis of motor vehicle expenditures. The level and change of unemployment is entered in the long and short run components of the Johansen vector error correction model (VECM), respectively. They find evidence of a single unique cointegrating relation between (the log of) motor vehicle expenditure, (the log of) income, the rate of unemployment and the interest rate. All variables are statistically significant and have the expected sign, income having a positive influence and the rates of unemployment and interest negative associations. The unemployment rate is found to be strongly exogenous with rising unemployment causing (preceding) increased saving rather than the other way around, which is consistent with the precautionary saving interpretation of unemployment.[83] The estimates of the cointegrating relation, by various methods, suggest that a one percentage point increase in unemployment causes a reduction in motor vehicle consumption of between 1.55 to 2.58 percentage points. Two policy implications of such a large precautionary saving motive are suggested. First, the response of consumers to changes in income may be close to that suggested by the AIH. Second, the greater is a country's social security provision the lower will be the precautionary saving motive which "provides support for those who are sceptical about the desirability of a single currency for Europe without first eliminating differences between countries social security systems." (Malley and Moutos, 1996, p. 598).

Evidence based upon the Engle and Granger (1987) and Johansen (1988) procedures are subject criticism. For example, Swamy and Tavlas (1992) highlight the difficulty in their ability to

[83] Malley and Moutos (1996) cannot discount unemployment acting as a proxy for liquidity constraints. However, it is argued that if many consumers are unable to borrow when unemployed, there is an incentive to accumulate savings as an insurance against unemployment suggesting a complementary connection between liquidity constraints and precautionary savings. Nevertheless, they present evidence which indicates that unemployment remains a prime measure of income uncertainty. Merrigan and Normandin (1996) suggest that instrumented squared consumption growth is an appropriate alternative proxy for income uncertainty.

uniquely identify underlying equilibrium relationships. If the nonstationary terms (typically entered as lagged level variables) do not represent long run information then one needs to explain the importance of these terms, which is evident in the ADL and cointegration ECMs success in modelling consumer behaviour. Possible alternative rationalisations are provided by Muellbauer's (1994) solved out rational expectations consumption function which suggests that lagged levels terms may represent habits, adjustment costs and/or expectations. Whatever the interpretation of these lagged levels terms they are argued to "convey substantial statistical advantages, providing that measurement errors in such data are stationary." (Muellbauer, 1994, p. 35).

2.6 Recent Evidence on LCH and Solved Out REPIH/RELCH Models

The first subsection considers Muellbauer's solved out REPIH/RELCH models with habits/ECM behaviour while the subsecond section reviews some recent evidence on LCH models.

2.6.1 Solved Out REPIH/RELCH Models with Habits/ECM Behaviour

Muellbauer and Murphy (1994) extend Ando and Modigliani's (1963) LCH formulation to explicitly model income expectations, account for uncertainty, credit constrained consumers, adjustment costs/habit formation, differences in assets' liquidity and aggregating over individuals in the face of demographic and income distribution changes. Using a CES/CRRA utility function they derive an approximation of unconstrained consumption in logarithmic form as a function of unconstrained income and the asset to income ratio (augmented by interest). Lagged consumption is added to accommodate habits/adjustment costs and/or expectations. Constrained consumption is approximated by current income plus a stochastic error. Assuming that unconstrained and constrained consumers' incomes move in parallel they derive the following consumption function:

$$\ln C_t = (1-\pi)b_{40} + b_{41}\ln Y_t + (1-b_{41})\ln C_{t-1} + (1-\pi)\delta[A_{t-1}/Y_t](1+r_{t-1}) + (1-b_{41})\pi\Delta\ln Y_t + u_t - \pi u_{t-1} \qquad \textbf{(2.6.1)}$$

This equation approximates the behaviour of consumers with simple, perhaps myopic, income expectations. Letting myopic consumers constitute the proportion, s, of the population, and the proportion, (1-s), having forward looking expectations one can replace the term $b_{41}\ln Y_t$ in **(2.6.1)** with:

$$b_{41}\ln Y_t = b_{41}[s\ln Y_t + (1-s)E_t\ln Y_t] = b_{41}[\ln Y_t + (1-s)(E_t\ln Y_t - \ln Y_t)], \qquad \textbf{(2.6.2)}$$

yielding:

$$\ln C_t = (1-\pi)b_{40} + b_{41}\ln Y_t + (1-b_{41})\ln C_{t-1} + b_{41}(1-\pi)(1-s)[E_t\ln Y_t - \ln Y_t] \qquad \textbf{(2.6.3)}$$
$$+ (1-\pi)\delta[A_{t-1}/Y_t](1+r_{t-1}) + (1-b_{41})\pi\Delta\ln Y_t + e_t - \pi e_{t-1}.$$

This basic specification is augmented by variables designed to capture demographic, income distribution, income uncertainty and interest rate effects and wealth is disaggregated into liquid and illiquid assets. Expected income is, employing rational expectations, proxied by a lead moving average in the growth rate of income forecasted using explanatory variables known in the current period.

The extended version of **(2.6.3)**, with consumption growth as the dependent variable, is estimated with IV using annual US data. A well defined equation which is robust to structural shifts in 1980/1 is obtained. They estimate the habits parameter (which can also be interpreted as an adjustment coefficient), $-b_{41} = -0.52$; the proportion of credit constrained consumers, $\pi = 0.33$; the rate at which future income is discounted, $\delta = 0.175$, which is high; and $s = 0.57$, suggesting that 57% of unconstrained households have myopic (random walk) income expectations. They also find statistically significant intertemporal substitution, demographic and uncertainty effects. The evident income distribution effect was suggested to be economically small.

For the UK they define a financial liberalisation variable, FLIB, as the *unexplained* rise in the loan to value ratio. Other variables featured in the UK, and not USA, equations, due to institutional differences, are measures of the intensity of credit controls and the uncertainty effects of strikes. The estimated UK equation yields similar coefficients to that of the USA except there is a weaker interest rate effect, possibly reflecting the use of both credit controls and interest rates for demand management policy, and differences in demographic effects. Notably, *after* the UK's liberalisation, the long run wealth effects are almost identical for both countries. The similarity of these two countries' specifications is argued to indicate that these models are structural, however, the explanations of the movements in their saving ratios are different. The rise in the liquid to illiquid asset ratio and the increase in spendability of illiquid assets is suggested to account for 70% of the fall in the UK's saving ratio from the end of the 1970s to the late 1980s. In contrast, the rise in the USA's asset ratio only accounted for approximately 30% of the decline in its saving ratio. Income uncertainty is argued to be greater in the UK than in the US. The UK model also featured some degree of parameter instability. This is argued to be due to the government's "misleading" suggestion that the 1988 tax cuts were irreversible making income expectations inaccurate and that the model assumes a constant proportion of credit constrained consumers.

Muellbauer (1994) suggests that a similar specification extended to cover the early 1990s explains the 1988-1991 decline in UK consumption by the sharp rise in real interest rates and the sustained slump to 1993 by declines in the real prices of shares and houses and the rapid rise in debt. Similar effects are suggested to plausibly explain the boom and bust in the Scandinavian countries.

Lattimore (1994) suggests that the fall in Australia's household saving ratio during the 1980s is caused by extremely slow income growth. The model employed to explain this relatively rapid consumption growth utilises a basic LCH formulation, similar to that used by Muellbauer and Murphy (1994), taking a form similar to a HUS-style ECM. The proportion of liquidity constrained consumers is estimated to range from 30%-48%. The only expectation variable which is significant embodies one year ahead projections suggesting Australians plan over very short horizons. The house-price to income ratio (multiplied by the proportion of non-home owners) is deemed essential to obtain sensible wealth effects and is consistent with prospective homeowners increasing saving. The impact of increased wealth raising home-owners consumption is captured by the ICM. "The fact that inflation is not at all significant when variable addition tests are conducted suggests that its relevance in models in the tradition of Davidson *et al.* (1978) can be traced to omitted wealth variables." (Lattimore 1994, p. 65). No clear role was found for either nominal or real interest rates. Private sector credit growth is positive and statistically significant and evidence suggests that it captures exogenous supply constraints rather than being demand determined. Unemployment is negative and statistically significant characterising uncertainty and liquidity constraint effects. There are significant demographic effects which particularly indicate saving in middle age. The parameter stability of the model suggests that deregulation has not significantly altered the fungibility of Australia's assets. Overall the model implies that consumers are contingent planners rather than forward looking which could be due to liquidity constraints, uncertainty and/or myopia.

Regarding expectations formation, Muellbauer (1994) suggests relaxing the *Muthian* assumption of costless information employed in the REPIH/RELCH. Two alternative approaches are outlined. First, rules of thumb which have no real theoretical basis may be used. Second, microeconomic theories of consumption may be employed to infer expectations processes. For example, the representative consumer's optimization problem may be defined to trade off information acquisition costs against the benefits of improved forecasting accuracy. "Once information acquisition and processing costs are introduced, the benefits of full information intertemporal optimization may be too small to warrant the costs - simple rules of thumb are likely to be optimal" (Muellbauer 1994, p. 23). Backward looking error correction mechanisms are suggested to be one rule of thumb that can be used in the face of prohibitive information collection costs. In models where both backward and forward looking expectations are accommodated, Muellbauer (1994), in contrast to Lattimore (1994), cites evidence supporting forward looking expectations of income growth for UK, USA and especially Japanese aggregate consumption.[84]

[84] Parameter stability tests are argued to be an appropriate means of discerning forward versus backward looking behaviour.

2.6.2 Recent Evidence on LCH Models

Pain and Westaway (1994) seek to explain the systematic underprediction (overprediction) of UK consumption in the 1980s (1990s) using an LCH based overlapping generations model. They utilise variables which directly embody the effects of the prevailing financial conditions, equity withdrawal and real consumer credit, because such measures characterise variations in the degree of financial conditions.[85] These deregulation proxies are found to reflect supply constraints, rather than being demand determined. Their estimated model suggests that long run total consumption is positively influenced by total wealth and augmented income and negatively related to the rate of interest.[86] Unemployment (precautionary saving) and lead consumption (changes in expected future income or tax) are also controlled for. Overprediction of the 1990s slump is removed when consumer credit and equity withdrawal are held at their actual historical levels in simulation. They argue that difficulties in predicting the 1990s slump are due to problems in accurately predicting the sharp and unexpected decline in credit.

Jappelli and Pagano (1994) extend Modigliani's (1990) model of the national saving rate to allow for liquidity constraints. Their empirical analysis uses a panel of nineteen OECD countries over three periods, 1960-70, 1971-80 and 1981-87. It is found that GDP growth is positively and generally significantly correlated with national saving, as is government saving, offering evidence against Ricardian equivalence, while the dependency ratio is always highly insignificant. The maximum loan-to-value ratio (LTV) is found to be negative and statistically significant in the savings function and the estimates suggest that a 10% increase in LTV reduces national saving by 2%. The product of LTV and GDP growth is added to the model to test whether the degree to which GDP affects savings is related to the severity of liquidity constraints. This variable is negative and statistically significant suggesting that the impact of

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[85] Pain and Westaway (1994) do not portray constrained consumers as following a simple rule, such as consuming all their income, because this assumes that agents have no access to credit, rather than limited access. Nor do they construct dummy variables to proxy the evolution of formal controls, such as mortgage rationing, because their typical implementation imply that credit restrictions were completely removed by deregulation rather than varying in intensity.

[86] Consumer credit and equity withdrawal are added to disposable income to yield augmented income.

GDP growth upon savings is less in countries with better access to credit. This provides evidence that variations in liquidity constraints explain differences in *national* saving behaviour.

Pesaran, Haque and Sharma (1999) employing panel estimation methods find that only fiscal variables have a statistically significant impact upon OECD private savings to GDP ratios.[87] Indeed, they find that the government surplus almost completely offsets private saving (the coefficient is 0.9) indicating virtually complete Ricardian equivalence. Their results contrast with many previous studies which are suggested to be subject to bias.

2.7 Conclusions

The RIH, PIH and LCH were all developed to explain the perceived empirical failings of the AIH. However, the statistical procedures employed in these early studies provided invalid inference on the AIH, especially those using time-series data. The observational equivalence of tractable forms of the RIH, PIH and LCH made it difficult to empirically assess the superiority of one theory over the other. Indeed, the PIH and LCH are similar in many respects, particularly their characterisation of consumers as forward looking and that lifetime (or future) resources rather than current income determines consumption. Both suggest a role for wealth, either explicitly or implicitly, although data constraints has restricted research using explicit asset variables. Nevertheless, the PIH-LCH provides the framework for the majority of modern work on consumer behaviour, if much recent analysis has focused upon relaxing its stronger assumptions.

The REPIH/RELCH framework introduced rational expectations and has been further modified to allow for durability, variable interest rates, current income consumers and a finite planning horizon. Excess sensitivity to current income growth appears to be the most popular explanation for the rejection of the REPIH/RELCH. Whether the degree of excess sensitivity is variable through time and whether it is due to liquidity constraints or precautionary savings appears to

[87] They find no evidence that output growth, the inflation rate, the real interest rate the ratio of wealth to GDP (wealth is measured as cumulated savings) or the dependency ratio have long run statistical effects on OECD countries' private savings rate.

be the focus of current research. The majority of evidence on the REPIH/RELCH has been confined to the UK and the USA.

The other dominant strand of research into consumer behaviour since 1978 has been the ECM framework pioneered by DHSY. Their model encompassed typical empirical formulations of the AIH, RIH, PIH and LCH, highlighted a need to account for inflation and emphasised the importance of lagged level terms to account for long run behaviour. HUS, and others, subsequently encompassed the DHSY model with the incorporation of wealth. When asset effects are directly modelled there is no role for inflation except, possibly, to adjust the conventional measure of disposable income for inflation induced losses on assets. The move towards deregulation, especially financial liberalisation, in industrial economies during the 1980s led to sharp increases in wealth (and possibly its spendability) which in turn caused consumer booms. This led to the need to account for broadly measured asset effects. The cointegration methodology provides ambiguous evidence on whether consumption, income and inflation, on their own, constitute a long run relation, however, it does suggest that wealth is an important determinant of long run consumer behaviour. The subsequent slump in industrial economies during the early 1990s requires other explanatory factors, candidates include; changes in income uncertainty, *carefully* measured credit variables, age structure (and other demographic effects), interest rates and income distribution. Once again the majority of the evidence is for the UK and USA.

Recently developed solved out rational expectations consumption functions allows one to determine whether consumers are forward or backward looking. The recent evidence is mixed, with some countries' consumers appearing to be forward looking (though not into the distant future) and others are better characterised as contingent planners. This suggests that both rational expectations and ECM formulations are worthy of pursuit.

CHAPTER 3

ANALYSING CONSUMPTION, INCOME AND INFLATION DATA

3.1 Introduction

This Chapter conducts an analysis of the main time series data under study. The next section outlines definitions and coverage of the series, the third draws some basic initial inferences while section 3.4 draws conclusions. The information obtained is from a combination of visual inspection of the data, basic descriptive statistics and inference from augmented Dickey Fuller tests. The information sought is whether agents smooth consumption relative to income, the degree to which series are subject to outliers and structural breaks, whether consumption and income move/break together through time, whether GDP is a good proxy for income and, most importantly, to gauge the series' orders of integration. This information will influence the modelling strategy conducted in future chapters.

3.2 Data Definitions And Coverage

The data to be analysed are annual observations available over the period 1955-1994 for twenty of the member countries of the Organisation for Economic Cooperation and Development (OECD). These are: Australia (AUL), Austria (AUT), Belgium (BEL), Canada (CAN), Denmark (DEN), Finland (FIN), France (FRA), Germany (GER), Greece (GRE), Iceland (ICE), Ireland (IRE), Italy (ITA), Japan (JAP), the Netherlands (NET), Norway (NOR), Spain (SPA), Sweden (SWE), Switzerland (SWZ), the UK (UK), and the USA (USA).

The series to be considered are the natural logarithms of real (1990) per-capita total private consumers' expenditure (LC*), real (1990) private disposable income (LY*) and real (1990) GDP (LG*). Where * indicates the country. The log of the consumers implied price deflator (LP*) is the ratio of current to real consumption multiplied by 100, so 1990~100. Two measures of the logs of the real (1990) per-capita average propensities to consume (APC) are constructed, being the logs of the ratios of real consumption to real income (LCY*) and real consumption to real GDP (LCG).[1] Differences of these series are prefixed by a D for a first difference and DD for a second difference.

Natural logarithms are taken of all the series to compress variation so reducing potential heteroscedasticity and to help linearise typically exponential series potentially reducing the degree of differencing required to induce stationarity.[2] Logarithms also facilitate testing hypotheses regarding ratios of variables, for example, whether consumption is homogenous of degree one in income. Further, researchers typically use models in logarithmic form, for example, Campbell and Mankiw (1991) - CM hereafter - and Davidson *et al* (1978) - DHSY hereafter. Since we estimate both DHSY and CM-style models it is appropriate to investigate the properties of the logs of variables.

A more detailed discussion of the data, including definitions, construction, sources, coverage and transformations is given in appendix 3.1A. A particularly noteworthy point is that the Federal Republic of Germany and the German Democratic Republic were united monetarily, economically and socially on July 1, 1990 and unified on October 3, 1990. Data refers to west Germany over the period 1955-1990 (inclusive) and to *unified* Germany from 1991-1994. The fall in per-capita GDP, disposable income and consumption in 1991 likely reflects the impact of eastern Germany's lower living standards on the overall aggregates. Dummy variables will be considered to accommodate this effect in both univariate and multivariate analysis.

3.3 Data Analysis

In this section we consider whether consumption is smoothed relative to current income (as predicted by theory), identify breaks and outliers in series (which may affect inference and modelling), and casually assess co-movements in consumption and income (suggesting

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[1] We also note the orders of integration of the unemployment (U*) and real interest (RI*) rates, which have a peripheral role (as instruments, for example) in the thesis.

[2] Following the literature we do not take the log of unemployment.

cointegration and cobreaking). We also consider whether GDP and income feature similar temporal evolutions and identify all series' orders of integration.

3.3.1 Is Consumption Smooth Relative to Current Income?

The permanent income and life cycle hypotheses (PIH-LCH) suggest that household consumption is smoothed relative to current income. Optimal consumption, based upon future expected income, will be a relatively constant *sustainable* flow (Pemberton, 1993). *Sustainable consumption* is unaffected by transitory changes in income, only responding to alterations in expected income: consumption will be smooth relative to current income.

Such smoothing of consumption rests on potentially questionable assumptions such as perfect capital markets. The presence of liquidity constraints suggests a greater sensitivity of consumption to current income than predicted by the pure PIH-LCH. For example, Campbell and Mankiw (1991) and Jin (1994) find evidence of a significant proportion of current income consumers in many countries. However, if there is also a significant proportion of consumers who follow the PIH-LCH one may expect some degree of consumption smoothing.

Deaton (1992) argues that uncertainty over future income may induce precautionary saving and act as a disincentive to borrow (for consumption purposes) upon the basis of uncertain expected income. Uncertainty provides a direct disincentive to use capital markets to smooth consumption, making expenditure more reliant upon current income.

Banks *et al* (1994) suggest that the needs of families may make it optimal *not* to smooth consumption over the life cycle. For example, young couples planning children may expect reduced income and increased expenses in mid-life, compared to those not raising children, and so may be less inclined to borrow when young because the burden of debt-repayments in mid-life may be considered too high combined with the extra costs of child rearing.

The discounting of future income, according to Muellbauer (1994), reduces its present value so making current income more important in consumption allocation decisions. High discounting would therefore reduce the propensity to smooth consumption.

Muellbauer (1994) also argues that consumers' information may be insufficient to yield an accurate view of expected future income. Rules of thumb, like backward looking error correction behaviour, may provide a more accurate description of how consumption decisions are made. If true this questions the degree of consumption smoothing.

Pemberton (1993) argues that theories suggesting households make detailed plans of consumption in each future period of their lives is fanciful. A more realistic plan is to allocate consumption between two periods - the present and the future. There is no detailed plan to make consumption in each future period equal (or yield the same utility), rather consumption is continually allocated between the present and future upon the basis of (potentially continuously revised) income expectations. Due to the less certain notion of the future this may imply less consumption smoothing relative to the predictions of the PIH-LCH.

Further, our use of total consumer's expenditure may be more volatile than the *actual* consumption of goods and services because it contains durable expenditures. This is because durable *expenditures* are intermittent with the consumption services they provide being a smoother flow. However, as Jin (1993) argues, this problem may be partially ameliorated for two reasons. First, the low annual frequency of the data means that many (semi) durables will have been consumed within the year reducing the non-synchronisation between expenditure and *actual* consumption. Second, aggregation of households' consumption offsets the unusually high expenditures of some households in a particular year, due to the purchase of durables, against the low spending of other households who made their durable purchases in earlier years (or intend to make them in later years). Overall, durability may cause a small bias against the consumption smoothing hypothesis for total expenditures.

This smoothing hypothesis is assessed with two formal measures of dispersion: the standard deviation and coefficient of variation. Table 3.1 reports both measures of dispersion for the logs and growth rates of aggregate annual consumption (LC and DLC) and disposable income (LY and DLY). For thirteen out of twenty countries the standard deviation of LC *exceeds* that of LY.

TABLE 3.1: Descriptive Statistics

	Standard Deviations		Coefficients of Variation		Correlations	
	LC	LY	LC	LY	LY/LG	LCY/LCG
AUL	0.201	0.184	0.044	0.042	0.988	0.101
AUT	0.297	0.327	0.123	0.153	0.998	0.391
BEL	0.276	0.292	0.225	0.273	0.996	0.387
CAN	0.267	0.293	0.059	0.066	0.995	0.070
DEN	0.170	0.163	0.063	0.064	0.964	-0.007
FIN	0.297	0.279	0.089	0.085	0.997	0.064
FRA	0.265	0.261	0.088	0.091	0.996	0.655
GER	0.277	0.263	0.066	0.064	0.994	0.512
GRE	0.355	0.401	0.530	0.832	0.999	0.440
ICE	0.341	0.298	0.688	1.120	0.973	0.055
IRE	0.226	0.226	0.040	0.041	0.985	-0.027
ITA	0.359	0.307	0.076	0.069	0.995	0.864
JAP	0.427	0.420	0.064	0.064	0.996	-0.234
NET	0.256	0.256	0.062	0.064	0.995	0.166
NOR	0.248	0.238	0.089	0.086	0.992	-0.186
SPA	0.301	0.293	0.040	0.040	0.995	-0.517
SWE	0.155	0.131	0.057	0.050	0.962	-0.634
SWZ	0.170	0.191	0.045	0.052	0.997	0.211
UK	0.227	0.220	0.041	0.041	0.991	0.422
USA	0.213	0.199	0.048	0.045	0.994	0.661
AVE20					**0.990**	**0.170**

TABLE 3.1 (continued)

	Standard Deviations		Coefficients of Variation		Correlations		
	DLC	DLY	DLC	DLY	DLY/DLG	DLC/DLY	DLC/DLG
AUL	0.0136	0.0229	0.6913	1.3078	0.6716	0.7249	0.6824
AUT	0.0179	0.0209	0.6060	0.6413	0.6601	0.7001	0.7218
BEL	0.0187	0.0247	0.7198	0.8545	0.7604	0.7669	0.7157
CAN	0.0253	0.0279	1.1515	1.1702	0.7893	0.8220	0.9246
DEN	0.0286	0.0392	1.4228	1.9138	0.5407	0.5746	0.7998
FIN	0.0330	0.0331	1.1966	1.2603	0.7109	0.8000	0.9059
FRA	0.0153	0.0205	0.5804	0.7666	0.7220	0.8132	0.8312
GER	0.0249	0.0301	0.9406	1.1806	0.7983	0.9036	0.8694
GRE	0.0255	0.0423	0.7585	1.1018	0.8760	0.8652	0.8484
ICE	0.0617	0.0751	2.4519	3.4844	0.7471	0.8754	0.8208
IRE	0.0306	0.0350	1.2314	1.3946	0.6212	0.6614	0.6217
ITA	0.0233	0.0270	0.6501	0.8702	0.6132	0.7392	0.8733
JAP	0.0289	0.0324	0.6446	0.7313	0.8613	0.8902	0.8886
NET	0.0217	0.0257	0.8167	0.9643	0.8414	0.8164	0.7583
NOR	0.0266	0.0195	1.1152	0.7685	0.3615	0.4795	0.6192
SPA	0.0334	0.0333	1.1126	1.0401	0.8658	0.8236	0.8795
SWE	0.0218	0.0218	1.503	0.7059	0.5161	0.5655	0.7286
SWZ	0.0175	0.0238	1.0417	1.2377	0.8022	0.8368	0.8345
UK	0.0226	0.0235	1.0447	1.0622	0.6484	0.8040	0.8154
USA	0.0167	0.0166	0.8184	0.8808	0.7644	0.8089	0.8745
AVE20					**0.7086**	**0.7636**	**0.7907**

TABLE 3.1 notes. The standard deviation and coefficient of variation is reported for LC, LY, DLC and DLY for all twenty countries over the period 1960-1994. The correlation coefficient for LY and LG, LCY and LCG, DLY and DLG, DLC and DLY and DLC and DLG are also given. "AVE20" denotes the arithmetic mean of the specified statistic for the twenty countries.

Using the coefficient of variation this figure falls to nine out of twenty countries, suggesting that the LCH-PIH prediction of consumption smoothing does not hold for about half of the countries under study. However, for sixteen (seventeen) of the twenty countries DLY exhibits a larger standard deviation (coefficient of variation) than DLC. For only three countries (Norway, Spain and Sweden) do we find unambiguous evidence that the smoothing postulate does not hold when using growth rates.[3]

The evidence is mixed regarding consumption smoothing. Given the potential for liquidity constraints, backward looking behaviour, income uncertainty the needs of families and the component of durability in our consumption measure, one might be inclined to interpret this as evidence against the general smoothing of consumption relative to income. This provides some tentative evidence for the need to consider the relaxation of some of the stronger assumptions employed in the pure PIH-LCH. This will be considered using formal modelling in later chapters.

3.3.2 Identification of Breaks and Outliers

It is desirable to identify breaks and outliers in data as they may need accommodation when modelling, using either explanatory or dummy variables. For example, data constraints may cause a broadly measured wealth variable to be omitted from a consumption function. Abrupt changes in consumption (income) which are not matched by similar changes in income (consumption) imply sharp changes in savings and, therefore, wealth. In such circumstances, the consumption function will omit an abruptly changing explanatory factor which will need accommodating.

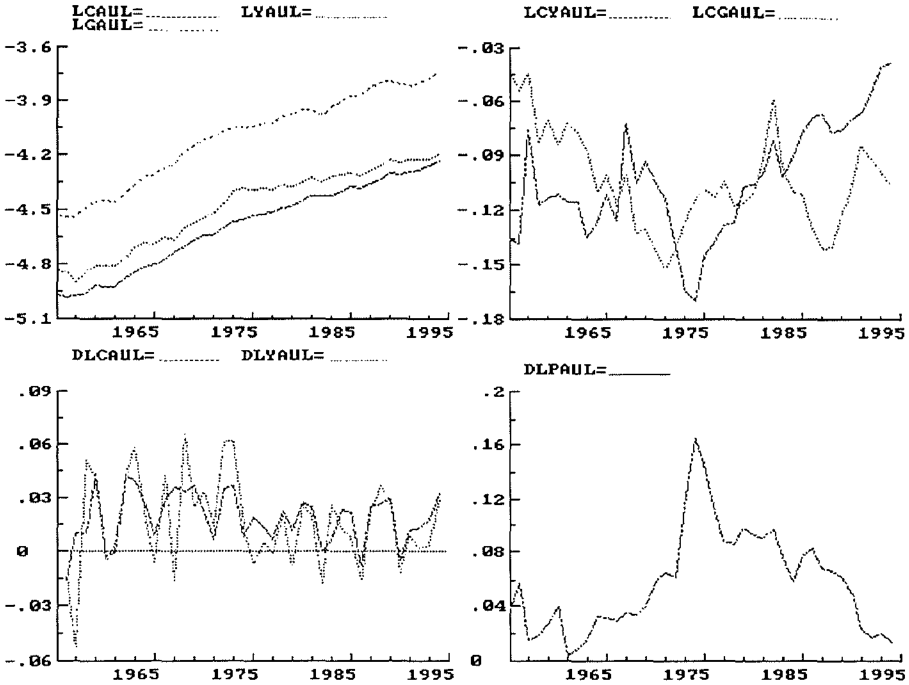
To identify features of interest we inspect our data which is plotted in Figure 3.1.[4] A plot, which

[3] The USA exhibits a smaller standard deviation for income relative to consumption but a larger coefficient of variation. This suggests that it is unclear whether consumers in the USA smooth their consumption. This may be surprising because the USA has relatively perfect capital markets suggesting that one might expect clear consumption smoothing in this country.

[4] A more detailed analysis by country is provided in appendix 3.2A.

FIGURE 3.1: Consumption Data Plots

Australia



Austria

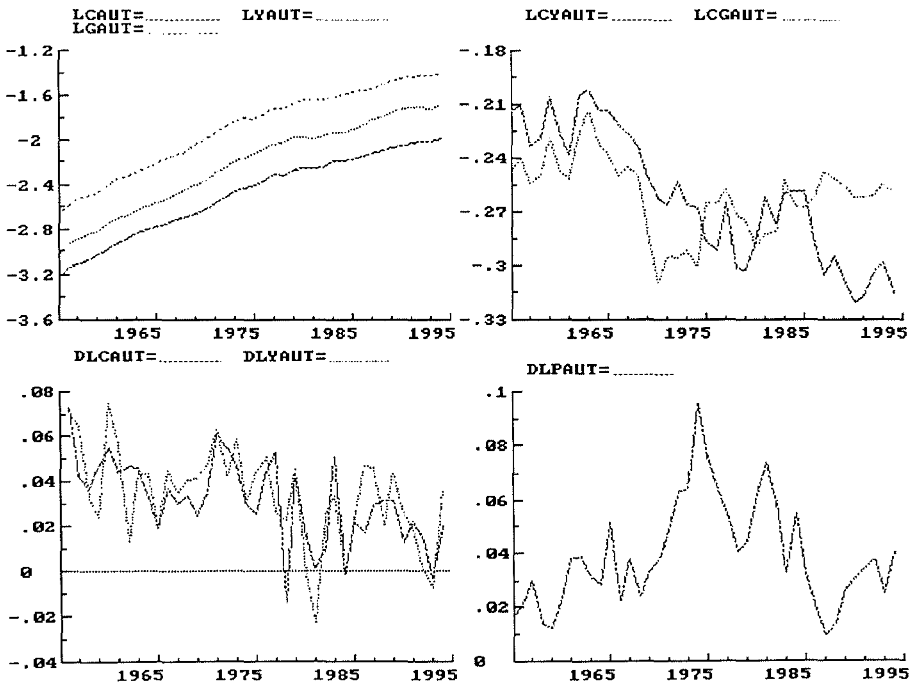
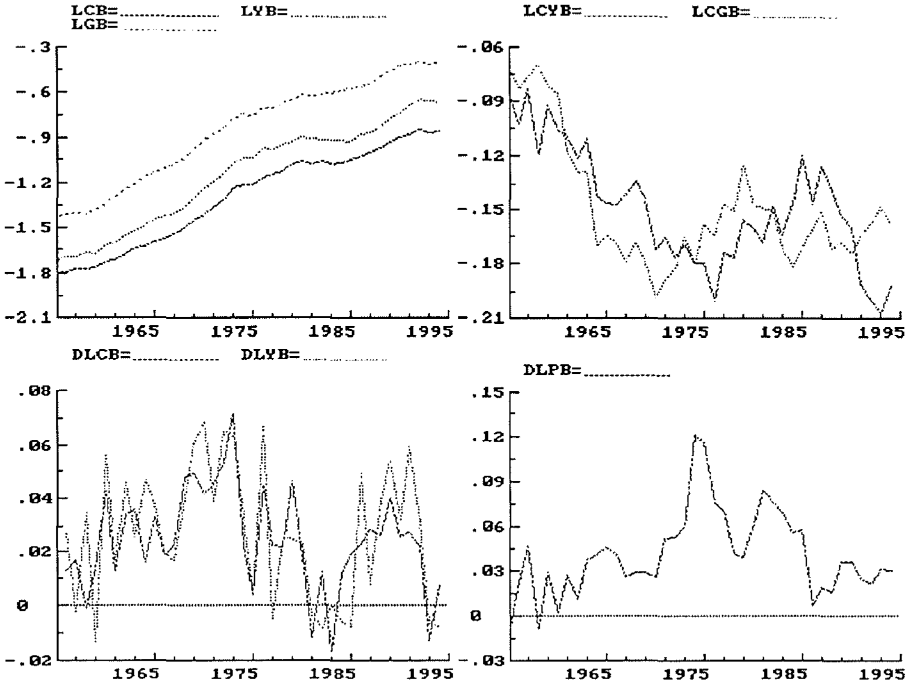


FIGURE 3.1 (Continued): Belgium



Canada

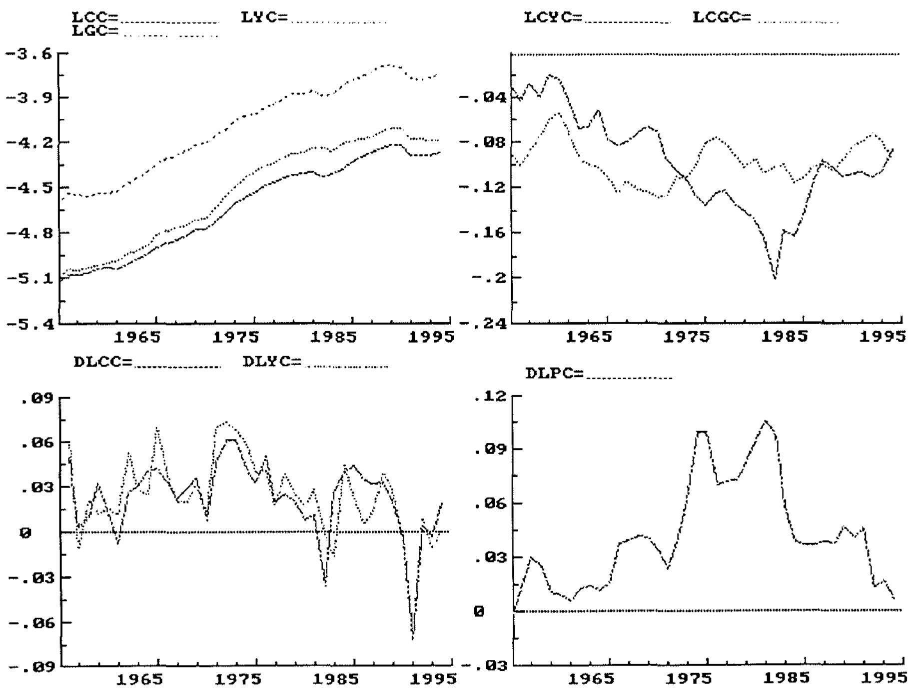
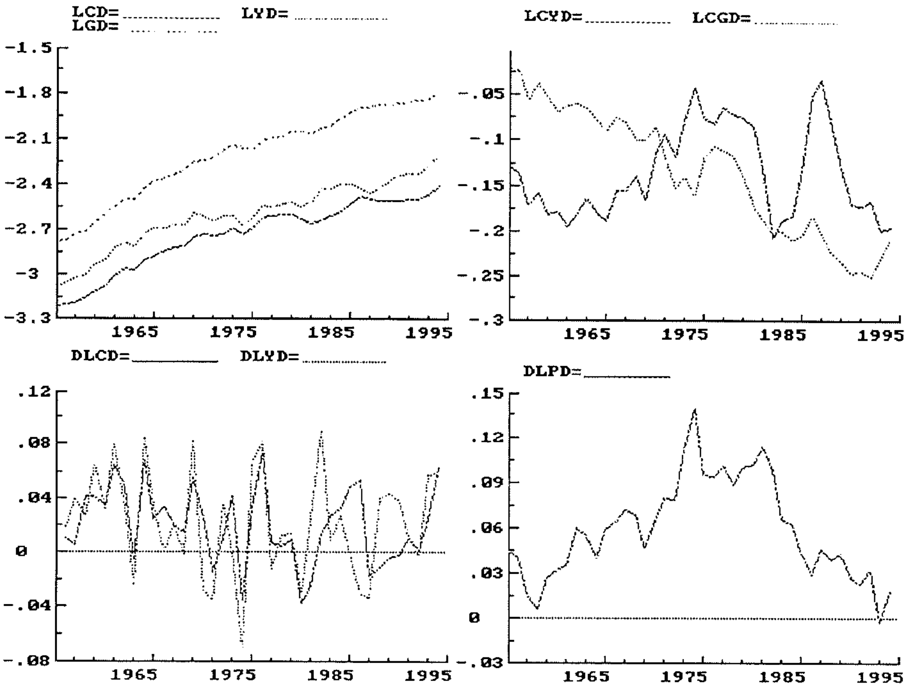


FIGURE 3.1 (Continued): Denmark



Finland

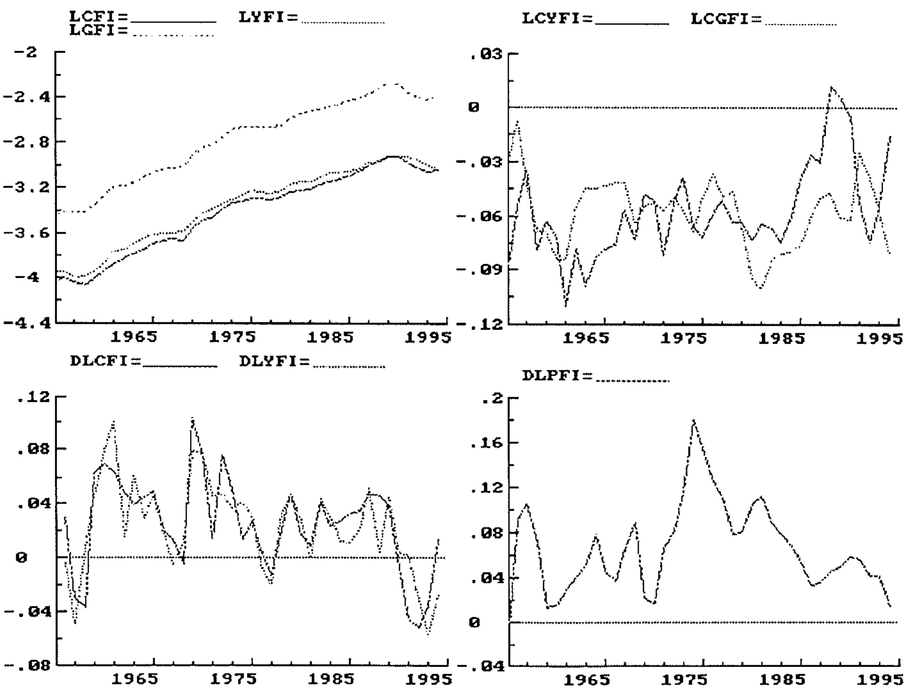
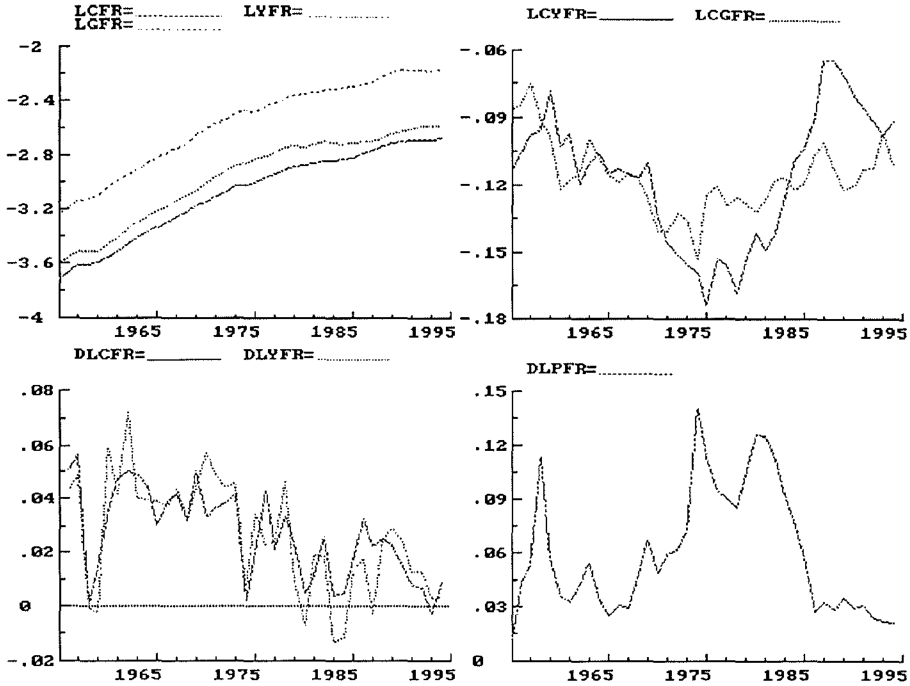


FIGURE 3.1 (Continued): France



Germany

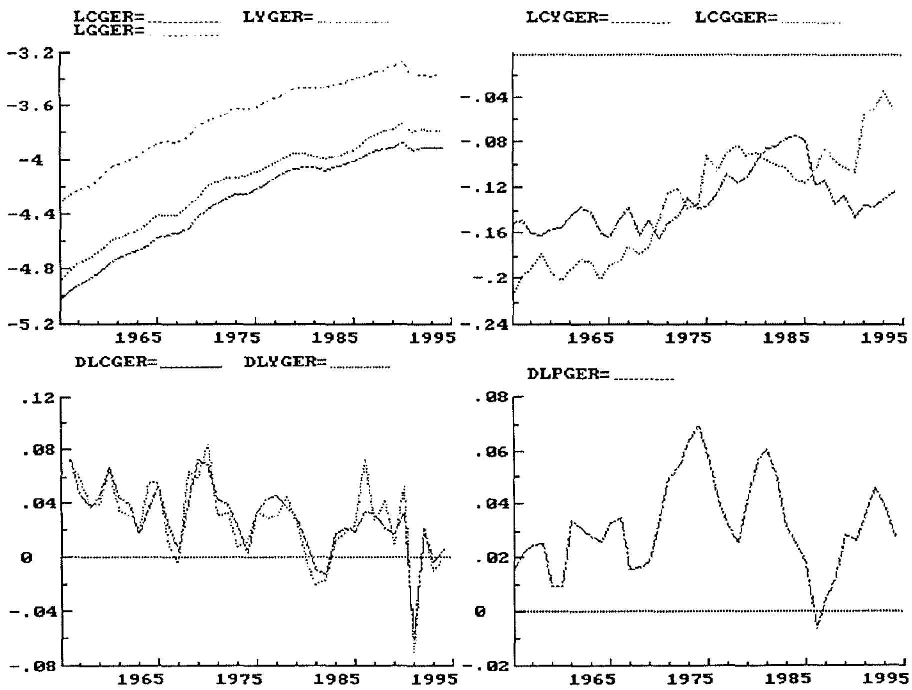
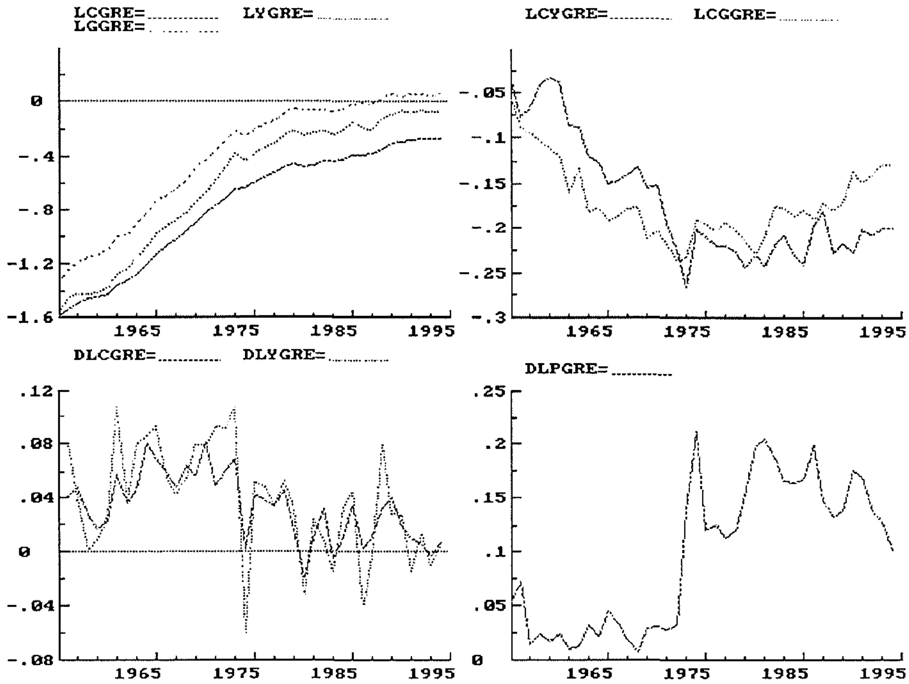


FIGURE 3.1 (Continued): Greece



Iceland

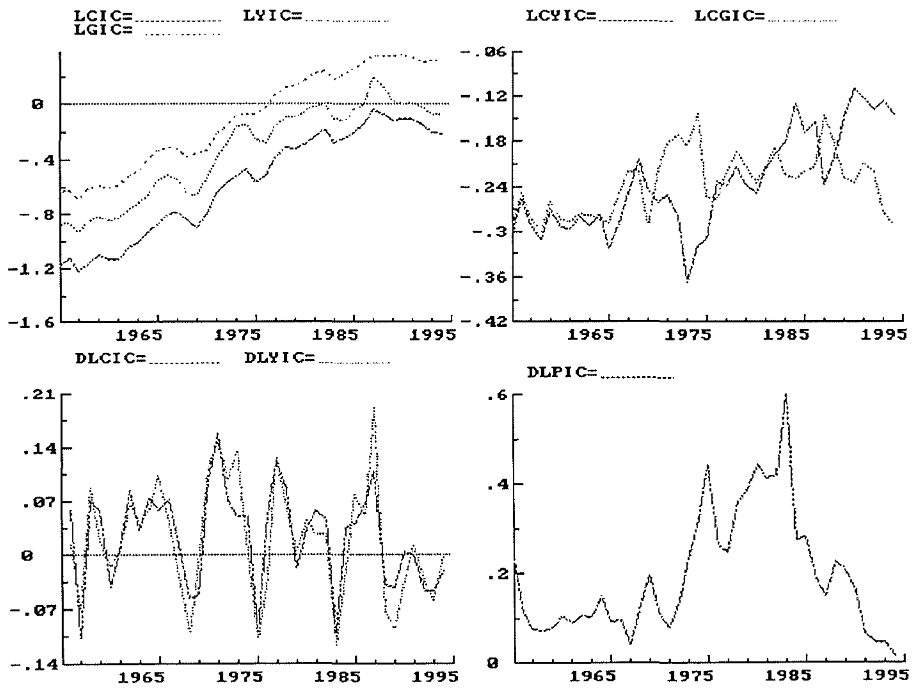
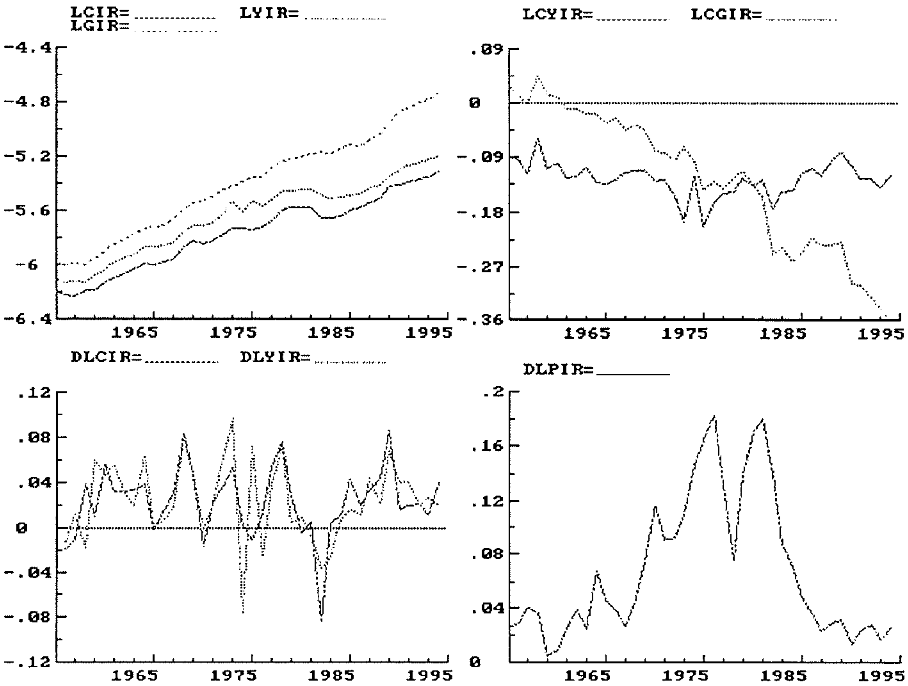


FIGURE 3.1 (Continued): Ireland



Italy

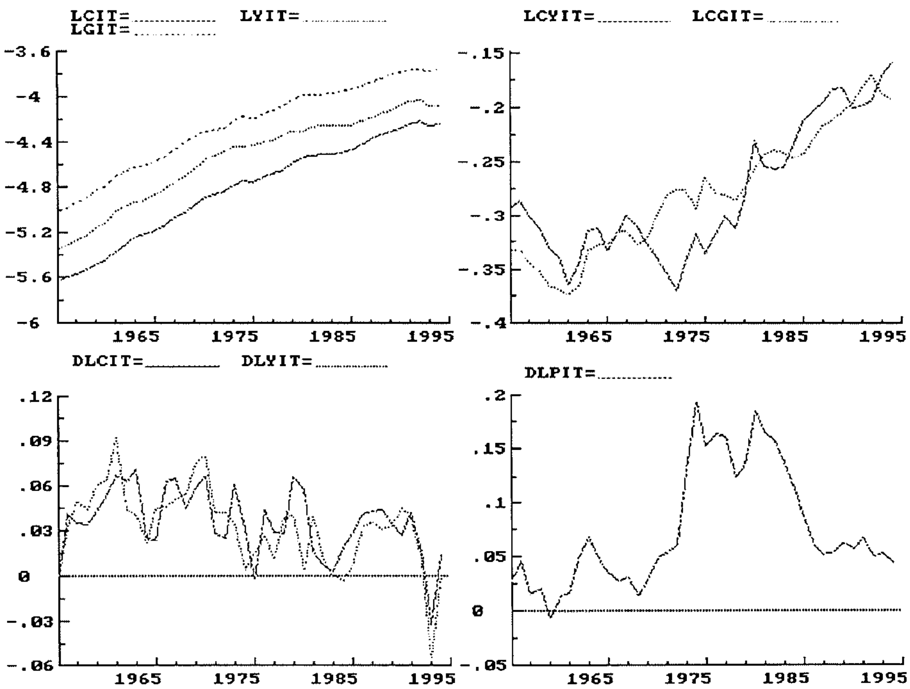
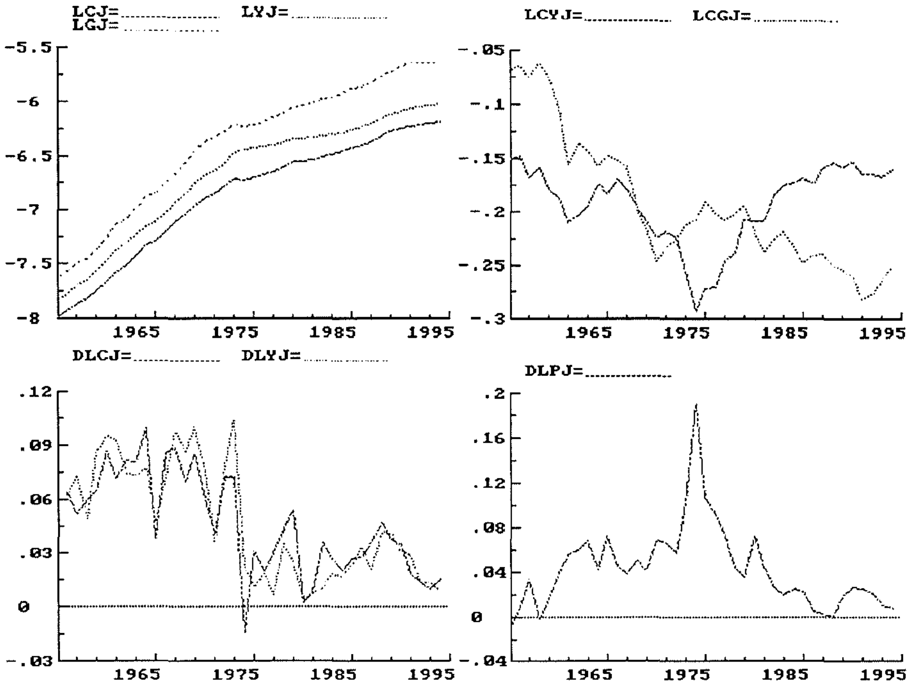


FIGURE 3.1 (Continued): Japan



Netherlands

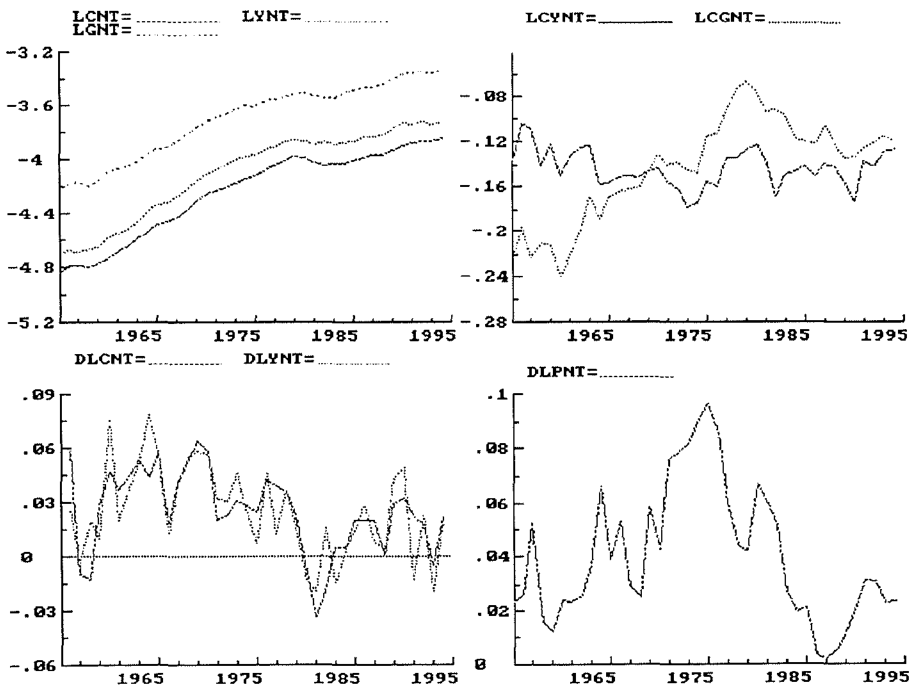
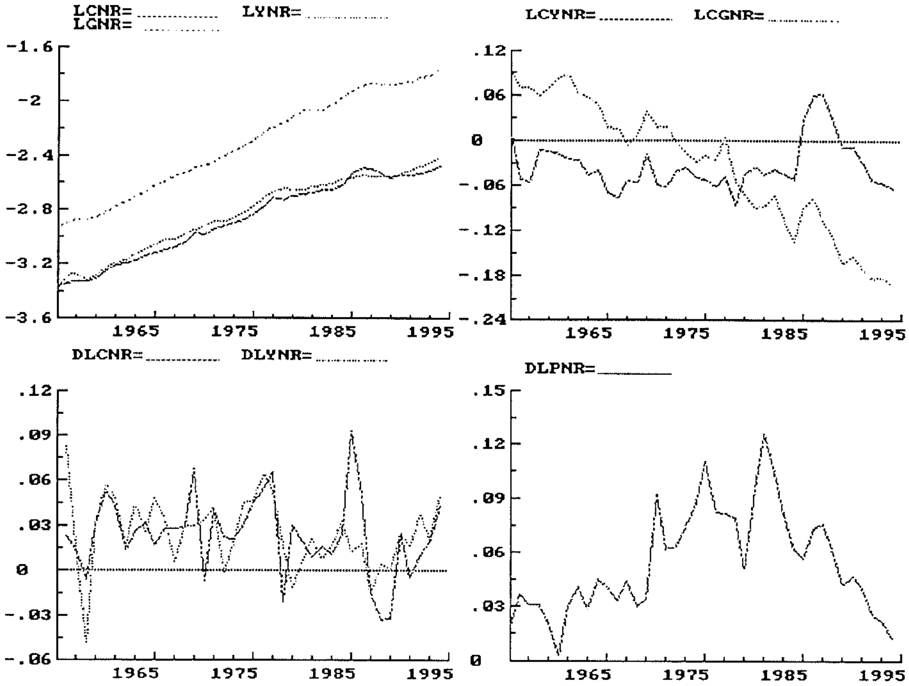


FIGURE 3.1 (Continued): Norway



Spain

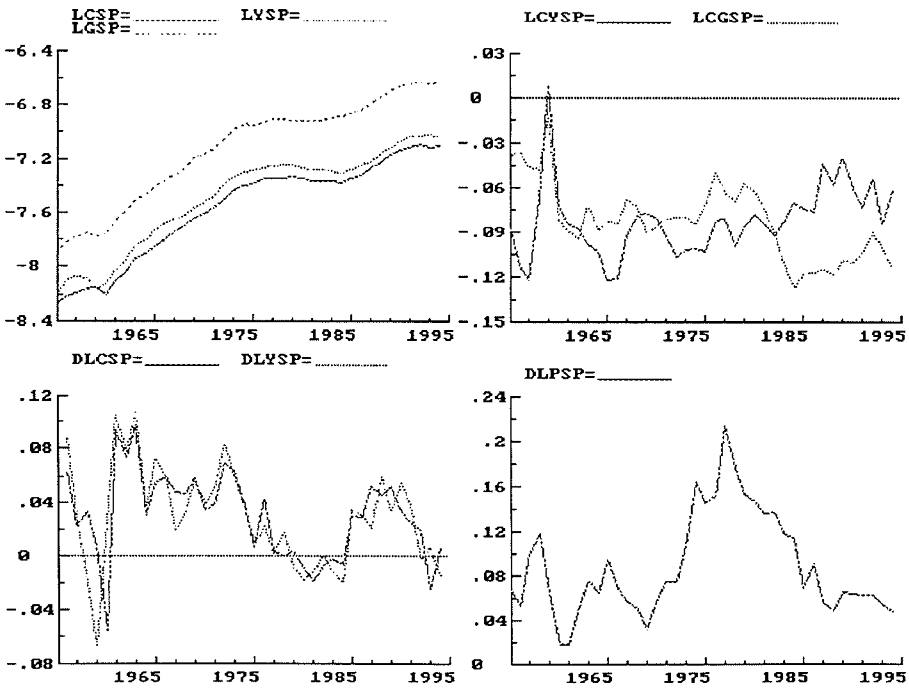
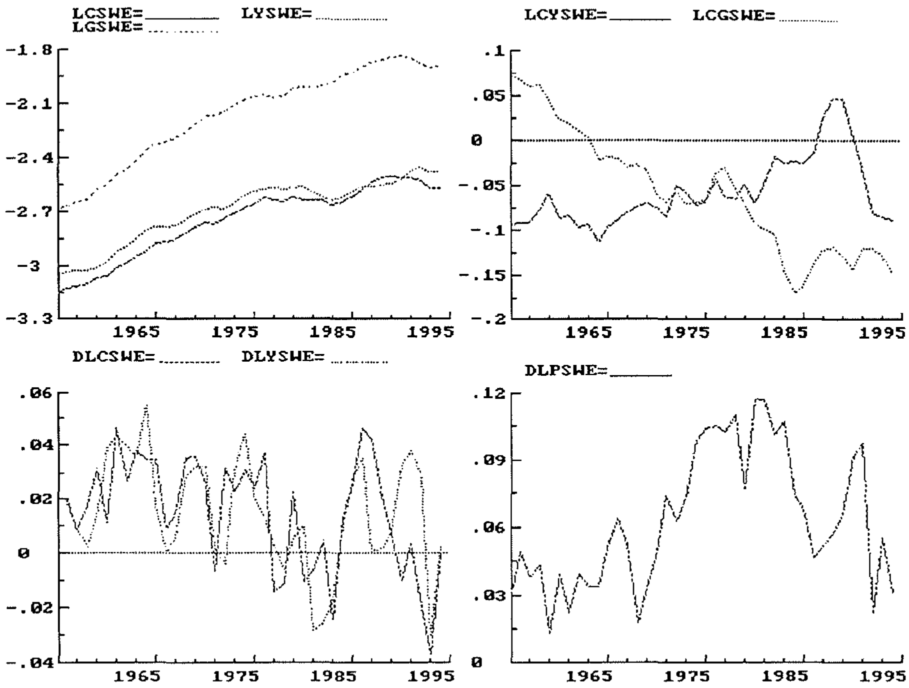


FIGURE 3.1 (Continued): Sweden



Switzerland

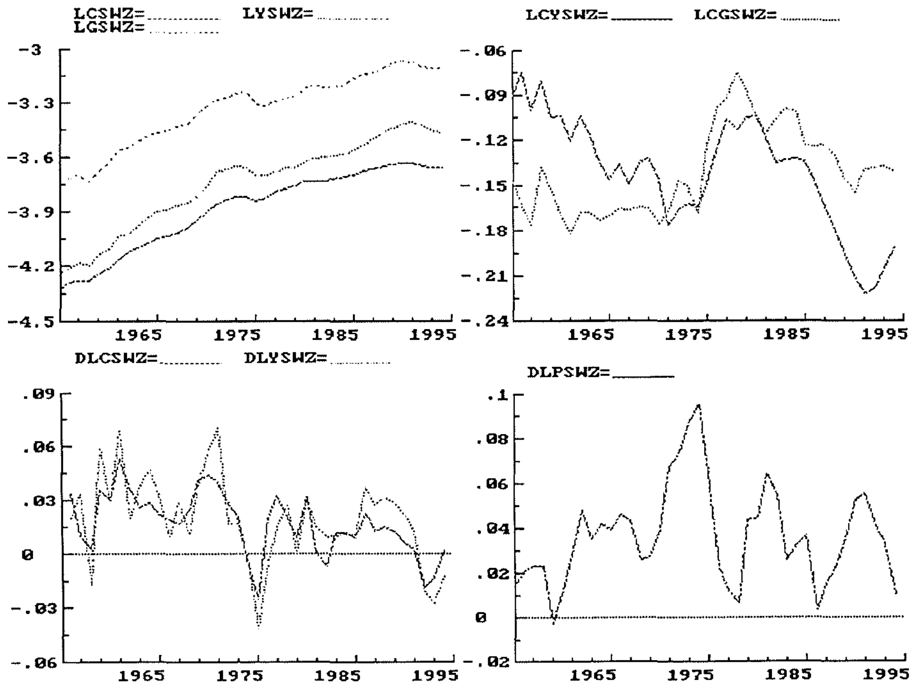
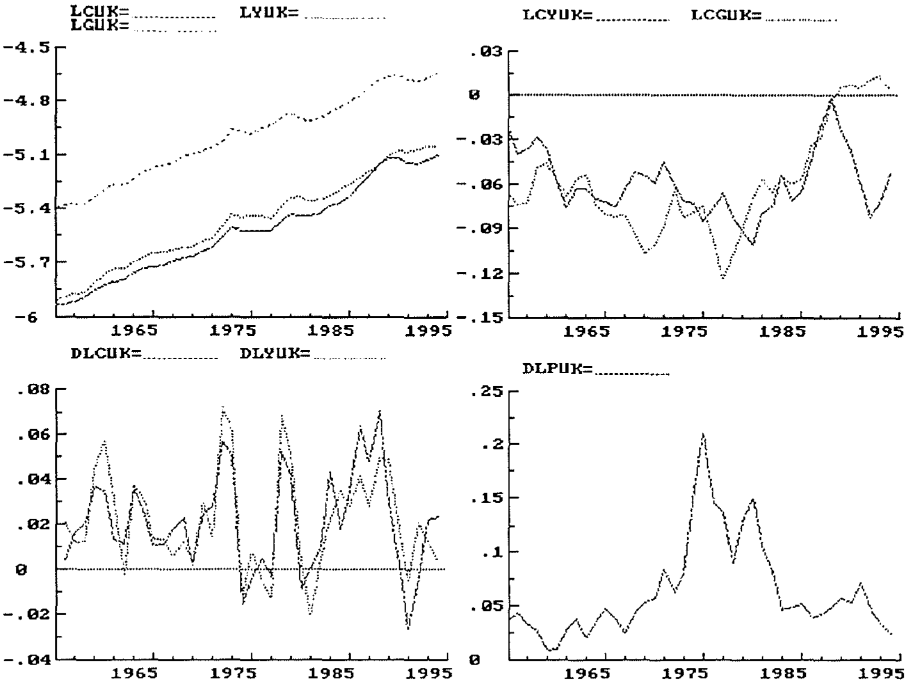
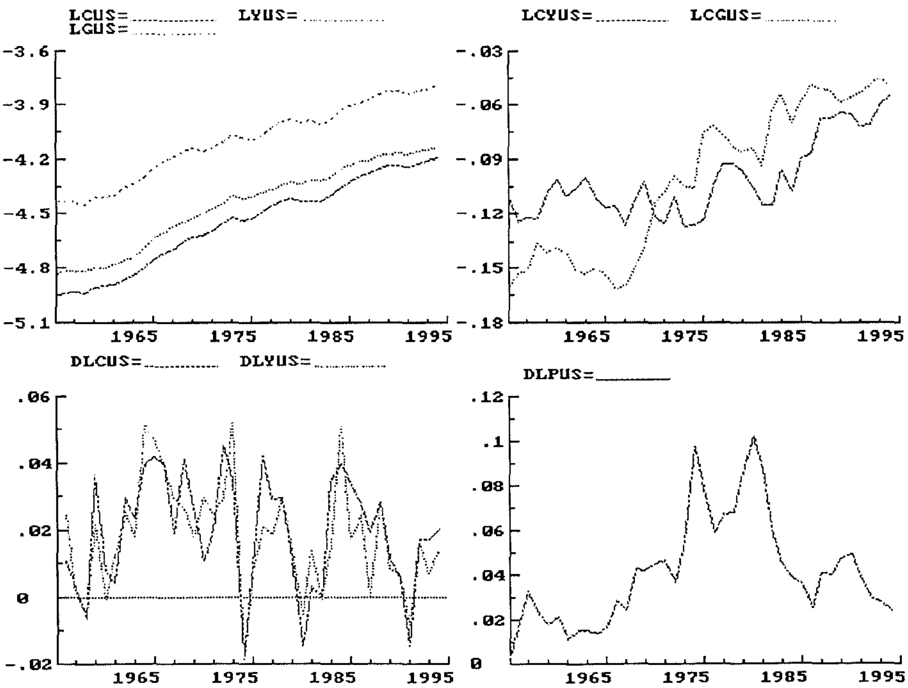


FIGURE 3.1 (Continued):UK



USA



is divided into four quadrants, is provided for each country. Quadrant 1 features the logs of real per-capita GDP, income and consumption. Quadrant 2 plots the logs of the consumption-income and consumption-GDP ratios. The mean of the latter is equalised to that of the former to provide easy comparison of their relative variations. Quadrant 3 plots the growth rates of consumption and income while the fourth quadrant graphs inflation.

Abnormal observations are apparent in many of the countries' series throughout the 1950s, which probably reflects post-war adjustment. With all models estimated over the period 1960-1994 these should not affect modelling too much.

Inflation rises sharply for all countries during the mid-1970s and generally falls in the 1980s and 1990s, suggesting outliers and/or breaks in this data.[5] This general feature may *explain* (either directly or as a proxy) patterns in consumption, income and the average propensity to consume (APC), which is our primary concern. Indeed, the rate of increase of consumption, income and GDP slows for most OECD countries after 1973/4 and again in the 1980s - being either split or damped trends - and are consistent with the oil shocks of 1973 and 1979. This is compatible with Barrell and Magnussen's (1996) analysis of the effects of oil shocks upon the world economy. The latter is also consonant with the adoption of monetary policies by many developed countries during this period.

With German reunification, consumption, income and GDP all shift down in 1991 and consumption and income growths correspondingly feature outliers in 1991. To the extent that income (and inflation) cannot explain the shift/outlier in consumption one may need to employ a *shift* and or *spike* dummy variable to remove evident misspecification from a German consumption function.

Financial deregulation in the Nordic countries' (Denmark, Finland, Norway and Sweden) during

[5] Unlike most other countries, Greek inflation shifts up in 1974 without subsequently falling. Consumption and income *growths* correspondingly shift down. This *permanent* shift may be caused by the oil shock *and* the conflict with Turkey over Cyprus *and* the restoration of democracy [I am grateful to Costas Milas (University of Warwick) for this explanation].

the 1980s (see Berg 1994), similar to the UK, manifests itself as a surge in consumption with APCs nearing or exceeding unity. Consumption subsequently collapses as households experience difficulty in repaying loans (possibly exacerbated by tax reforms). This boom and bust cycle is particularly acute for Finland.

Of course, outliers and/or breaks in income may *explain* corresponding outliers and/or breaks in consumption, making accommodation with additional explanatory and/or dummy variables unnecessary when building consumption functions. Thus, it is interesting to examine co-movements in consumption and income (and inflation).

3.3.3 Do Consumption and Income Move Together Through Time?

Keynes's (1936) absolute income hypothesis (AIH) suggests that consumption and income move together through time, forming an equilibrium relationship. DHSY suggests that consumption forms a long run dynamic relationship with income, inflation (and income growth), while consumption, income and wealth form an equilibrium according to Ando and Modigliani's (1963) LCH. Although these theories indicate the importance of current income in the long run determination of consumption they also specify other explanatory factors, suggesting that current income may not, on its own, form an equilibrium relationship with consumption. Friedman's (1957) PIH suggests that consumption only responds to current income to the extent that the latter causes permanent income to change. This implies that consumption need not follow current income at all.[6] However, binding liquidity constrains, income uncertainty, myopia and bounded rationality may mean that consumption follows current income more closely than predicted by the PIH-LCH. The theory is, therefore, mixed on the issue of whether consumption and income move together through time.

Through visual inspection of Figure 3.1 we obtain some initial insights into the movements of consumption and income. Inspection of the plots suggest that consumption and income do, for

[6] Jin (1993) shows that Campbell's (1987) PIH-LCH formulation implies cointegration between consumption and disposable income.

many countries, feature similar time paths. This includes the German data where consumption and income appear to shift down together in 1991 with reunification - which is an example of contemporaneous mean cobreaking. Similarly, Japanese consumption and income both feature split trends in 1974 and may, therefore, also cobreak. In contrast, Greek income shifts upwards in 1974 without a corresponding shift in consumption. At first sight this does not appear to be consistent with an equilibrium relationship between consumption and income, however, it may be that consumption adjusts slowly to the income shock (which would be consistent with agents smoothing consumption in the face of current income changes) retaining the long run relation - Raj (1995) makes a similar point regarding money demand. Such adjustment may be appropriately modelled in a dynamic modelling framework such as an error correction model (ECM). In Spain, income surges in the late 1950s which is followed by a sharp rise in consumption in the early 1960s. This might be interpreted as either intertemporal cobreaking or the absence of a complete equilibrium. Further, in Denmark, Finland, Norway, Sweden and the UK, consumption surges in the mid/late 1980s without a corresponding rise in income and is typically followed by plummeting consumption with a notably less severe decline in income. This suggests that consumption and income may not, on their own, cointegrate, there is some omitted effect, such as financial deregulation, which needs accommodation, possibly using wealth.

Both theory and visual inspection of data plots suggest that consumption and income move together through time, however, whether they form an equilibrium relation on their own, or if other factors need to be added to secure a long run relationship needs to be resolved. We will consider this issue using cointegration techniques in Chapter 4.

3.3.4 Is GDP a Good Proxy for Disposable Income?

Models of consumption based upon the PIH-LCH typically specify labour income as the appropriate income measure while ECMs are designed for Hicksian income. Disposable income, being a close approximation to both measures, is often employed as it is available for the countries generally studied, mainly the UK and the USA.[7] In contrast, recent cross-country comparisons of consumer behaviour have employed GDP to proxy income - see, for examples, Bayoumi and MacDonald (1995) and Carruth *et al* (1996). GDP is a broader measure of income incorporating both public and private sectors and without deducting tax payments and other transfers. It is, therefore, far further removed from labour and Hicksian income than disposable income. We therefore suggest that private disposable income is a far more appropriate measure of income than GDP, in the current context. Thus, if GDP is not highly correlated with income, one might be cautious in drawing too firm inferences from studies employing this measure.[8]

Table 3.1 gives the simple correlation coefficients for LY with LG, for the APCs LCY with LCG and for the growth rates DLY with DLG. Average correlations over the twenty countries are also reported (AVE20). As expected, due to shared trends, the levels of income and GDP are highly correlated with an average correlation of 99%. The growth rates of income and GDP feature a reasonable correlation, approximately averaging 71% - only three countries feature a correlation below 60%. However, the average correlation for the two APCs is 17%, with six countries exhibiting negative associations. This implies that models using the APC (or the log-levels of consumption and income) as dependent, explanatory or instrument variables, will potentially yield highly different inferences when using disposable income rather than GDP.[9] This includes both ECMs and some rational expectations models.

Disposable income, being far closer to the theorised measures of income for consumer behaviour than GDP, should be considered the more appropriate measure. Given the correlations of, in particular, the two APCs (and to some extent the growth rates) it is probable that quite different

[7] Labour income is disposable income less capital income while Hicksian income is disposable income less inflation induced losses on assets.

[8] Similarly we would argue that inferences drawn from studies using national disposable income minus government consumption expenditure, such as Jin's (1994), will provide inferior inference to those employing disposable income. Although this measure would be expected to provide superior inference relative to GDP, being closer to the theorised concepts of income.

[9] Unreported initial experiments testing the rational expectations PIH-LCH hypothesis and Granger non-Causality in a stationary vector autoregression show some divergence of inference when using these two different measures of income.

inferences may be drawn using the two different measures. One would be advised to always use disposable income, if available, rather than GDP for analysing consumer behaviour.

3.3.5 The Series' Orders of Integration

We wish to assess whether the series are second-order stationary, that is, whether their mean, co-variances and variance are constant through time. Although formal, augmented Dickey-Fuller (ADF) tests will be employed, these have low power, particularly in small samples and if incorrect lag lengths are employed (see Ayat *et al* 1995 and Linden 1995). Their inference may be sensitive to data breaks and outliers (see Perron 1989 and Zivot and Andrews 1992) and the presence of a moving average error process. Further, as Harvey (1997) points out, testing a series using a univariate random walk model with linear trend may be a too restrictive way of characterising a series. Tests for more flexible forms of stochastic nonstationarity are provided by fractional unit root tests (see Sowell 1992) and for nonlinear deterministic trends by Bierens (1997). Since our primary aim is to obtain initial insights into series' univariate properties for a broad range of countries we confine our analysis to identification of stochastic versus linear trends and the order of integration. Further, identification of trends, outliers and breaks through visual inspection of the plots will be used to augment inferences drawn from these basic ADF tests.

Inspection of each country's plot, see Figure 3.1, shows that LC, LY, and LG exhibit generally increasing means, as expected in growing industrial economies, and so cannot be stationary. ADF tests indicating these series are stationary will be rejected.

In contrast, consumption and income growth show no discernable trend which is consistent with stationarity. One exception is that Japanese consumption and income growth both shift down in 1974, apparently signalling the end of Japan's exceptional post war-growth. Such a large shock can may make Japan's consumption and income growth *appear* nonstationary, indeed, ADF tests indicate that consumption and GDP growth are trend stationary while income growth is nonstationary. Over the longer term these Japanese growth series may turn out to be I(0), as the 1974 shock becomes less pronounced relative to the longer sample. Bearing this caveat in mind, we will assume that these Japanese growth series are stationary to be consistent with the other OECD countries. Such consistency of inference is desired to facilite the application of a single model to all countries.

All countries' price growths follow a similar pattern. Inflation is low throughout the 1950s and 1960s, it rises sharply in the early/mid 1970s (with the oil price shock) and declines throughout the 1980s and 1990s (as OECD economies generally emphasise anti-inflation policies). Although the mean of inflation does not generally increase or decrease the variance is not obviously constant, so *may* be nonstationary. Greece is the exception because inflation shifts up sharply in 1973/4 without a subsequent decline suggesting that inflation is nonstationary, due to a nonconstant mean. Thus, the log of prices is probably I(1) or I(2).[10]

DHSY employ the microeconomic homogeneity postulate that consumption is homogenous of degree one in income: LC and LY cointegrate with a unit coefficient implying that (the log of) the APC is stationary. Bollerslev and Hylleberg (1985), for example, find evidence favouring the alternative hypothesis of a below unit income elasticity (for the UK), suggesting the APC is nonstationary (trended), which may be justified by Engel's law or, perhaps, a Keynesian postulate. Of course, the financial deregulation and tax reform of the 1980s and 1990s, particularly in Denmark, Finland, Norway, Sweden, the UK and USA, may have shifted the propensity to consume during this period, removing the APC's trend (though increasing its variance). Thus, our inferences may be sample specific, especially given Bayoumi's (1993) conclusion that the impact of financial deregulation in the UK is temporary. Visual inspection of Figure 3.1 suggests that the log of the consumption-income ratio features a reasonably constant mean for Australia, Denmark, Finland, Germany, Iceland, Ireland, Japan, the Netherlands, Norway, Spain, Sweden and the UK. Although this may suggest stationarity, these series' high volatility (variances) means they may be nonstationary. The other countries'

[10] Noticeably inflation is exceptionally high in Iceland which may undermine the difference of the log to approximate its growth rate. This approximation is employed to facilitate the same logarithmic model specifications for all countries.

consumption-income ratios are trended (nonstationary).[11]

More formal inference is obtained from the ADF test which is based upon the following equation:

$$\Delta X_t = a_1 + a_2 X_{t-1} + a_3 t + \sum_{i=0}^{L} b_i \Delta X_{t-i} + u_t \qquad \textbf{(3.1)}$$

where X_t is the series of interest, t is a time trend and u_t an error term. L is specified to ensure u_t is *white noise*, or at least non-autocorrelated.[12] The number of lagged dependent variables included in the test equations (L) is determined using Schwartz's Bayesian Information Criterion (SBIC) which is given by:

$$SBIC = k\,(\ln T/T) + \ln[(1/T)\hat{\mathbf{u}}'\hat{\mathbf{u}}] \qquad \textbf{(3.2)}$$

where k is the number of estimated coefficients and T is the sample size. The minimum value of SBIC gives the value of L which optimises the fit versus parsimony trade-off.

Since the SBIC may not ensure *white noise* residuals we gauge the validity of inference using the modified lagrange multiplier test first-order serial correlation (LMSC1) - see Spanos (1986) p. 521. This statistic is automatically produced by Microfit 3.22 and its $\chi^2(1)$ critical value is 3.84.

After the appropriate lag length has been chosen, the following hypotheses can be tested using **(3.1)**:

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[11] For the average propensity to consume (APC) measured as the consumption-GDP ratio, the following appear to exhibit a constant mean: Australia, Canada, Finland, France, Iceland, and Switzerland. The other fourteen countries' consumption-GDP ratios appear to be trended (nonstationary). Fewer of the APC series appear stationary using GDP compared to those using income.

[12] We could have incorporated dummy variables into many countries' test equations, requiring the simulation of new critical values for many different test equations. To avoid such complications we only consider the introduction of (reunification) dummy variables in Germany's ADF equation and employ standard critical values.

a) Excluding the trend ($a_3 = 0$) we can test for difference stationarity with: H_0: $a_2 = 0$, X_t is nonstationary; against H_A: $a_2 < 0$, X_t is stationary by comparing the test statistic (t-ratio), reported as TAU in Table 3.2, with the 5% critical value of -2.947 (see MacKinnon 1990).
b) Including the trend ($a_3 \neq 0$) we can test H_0: $a_2 = 0$ and/or $a_3 = 0$; X_t is nonstationary against, H_A: $a_2 \neq 0$ and $a_3 \neq 0$; X_t is stationary around a linear trend. The F-test statistic, reported as PHI in Table 3.2, is compared with the critical value 7.036 (see Dickey and Fuller 1981) to assess this hypothesis. We contend that, strictly, $a_2 < 0$ for stationarity around a linear trend. Further, H_A may not be rejected if $a_2 < 0$ but $a_3 = 0$, which would suggest the series is simply stationary rather than being stationary around a linear trend. We report the t-ratio of the time trend's coefficient (a_3) - denoted t-trend in Table 3.2 - to assess this.

Although the data is available over the period 1955 to 1994, all ADF tests use the standard sample of 1960-1994 to allow for differencing and lagged variables. This allows one to test the second difference series with L=2, which should be sufficient to remove autocorrelation in the annual data presently employed. The small sample size (35 observations) may affect the tests' reliability, however, longer data series for such a broad selection of countries is not available. The ADF test results, in conjunction with visual inspection, should allow us to gauge the orders of integration of our data - Horioka (1996 and 1997) and Jin (1993) conduct ADF tests for similarly small samples.

The ADF test results are reported in Table 3.2. There appears to be some heterogeneity across countries, however, the following generalisations can be drawn.[13] The logarithms of consumption, income and GDP generally appear to be I(1). Although both APCs (using income and GDP) might generally be I(0) they are more likely nonstationary, probably being I(1). Sarantis and Stewart (1998a) find both APCs to be nonstationary using panel unit root tests (see Im *et al* 1997 and Taylor and Sarno 1997). Prices are probably I(2), which is what we assume, but could be I(1), it is very difficult to determine a clear order of integration. Real interest rates

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[13] The SBIC rarely indicated L>1 and autocorrelation was generally not evident.

TABLE 3.2: ADF Tests

ADF Tests (Australia)

	LCAUL	DLCAUL	LYAUL	DLYAUL	LGAUL	DLGAUL	LPAUL	DLPAUL	DDLPAUL
TAU (L)	-1.222 (0)	-5.585 (1)	-1.603 (0)	-5.837 (0)	-1.557 (0)	-5.157 (0)	-1.391 (1)	-1.855 (1)	-5.143 (1)
LMSC1	0.282	4.286	0.088	0.973	0.051	0.141	2.604	3.280	0.170
SBIC	-8.437	-8.411	-7.423	-7.348	-7.540	-7.479	-7.752	-7.808	-7.860
PHI	1.342	17.819	1.657	18.415	2.028	14.004	3.819	1.408 (2)	15.081
t-trend	1.088	-1.658	0.872	-1.367	1.261	-1.108	2.332	-1.156	-1.577

	LCYAUL	DLCYAUL	LCGAUL	DLCGAUL	RIAUL	DRIAUL	UAUL	DUAUL	
TAU (L)	-0.862 (0)	-7.251 (0)	-2.637 (0)	-6.188 (0)	-1.323 (0)	-4.690 (0)	-0.576 (0)	-4.877 (0)	
LMSC1	1.235	1.314	0.325	1.128	2.243	1.323	0.873	1.370	
SBIC	-8.080	-8.112	-8.282	-8.095	-7.795	-7.782	0.150	0.143	
PHI	2.439	27.418	3.474	19.280	1.497	10.733	3.360	11.679	
t-trend	2.016	1.218	0.414	0.812	1.110	0.291	2.517	0.412	

TABLE 3.2 (continued): ADF Tests (Austria)

	LCAUT	DLCAUT	LYAUT	DLYAUT	LGAUT	DLGAUT	LPAUT	DLPAUT	DDLPAUT
TAU (L)	-3.661 (0)	-4.715 (0)	-3.239 (0)	-1.491 (2)	-4.044 (0)	-1.425 (2)	-0.933 (1)	-2.673 (0)	-6.968 (0)
LMSC1	0.286	2.689	1.215	0.292	0.002	0.065	0.349	0.122	0.054
SBIC	-8.189	-7.891	-7.810	-7.722	-8.014	-7.710	-8.370	-8.445	-8.291
PHI	6.598	19.694	5.277	2.731	7.985	19.419 (0)	1.781	3.788	24.295
t-trend	0.375	-3.264	0.537	-1.758	0.272	-3.462	1.627	-0.731	-0.782

	LCYAUT	DLCYAUT	LCGAUT	DLCGAUT	RIAUT	DRIAUT	UAUT	DUAUT	
TAU (L)	-1.507 (0)	-6.260 (1)	-2.350 (0)	-6.367 (0)	-2.474 (0)	-7.065 (0)	-0.256 (0)	-4.438 (0)	
LMSC1	0.105	0.399	0.435	1.002	0.747	0.026	2.091	0.029	
SBIC	-8.229	-8.203	-8.446	-8.296	-8.339	-8.246	-1.957	-2.029	
PHI	4.189	18.990	2.698	20.330	3.982	24.472	5.302	10.810	
t-trend	-2.404	0.100	0.183	0.778	1.309	0.466	3.243	1.256	

TABLE 3.2 (continued): ADF Tests (Belgium)

	LCB	DLCB	LYB	DLYB	LGB	DLGB	LPB	DLPB	DDLPB
TAU (L)	-2.374 (0)	-3.884 (0)	-2.271 (0)	-4.604 (0)	-3.755 (0)	-4.433 (0)	-0.592 (1)	-2.319 (0)	-6.869 (0)
LMSC1	3.442	0.041	2.041	0.047	0.002	2.969	0.036	0.124	1.373
SBIC	-7.909	-7.891	-7.345	-7.254	-7.905	-7.612	-7.658	-7.749	-7.618
PHI	2.019 (1)	9.635 (1)	2.703	13.039	6.916	17.114	2.828	2.651	23.252
t-trend	0.587	-1.785	0.591	-1.835	0.334	-3.084	2.293	-0.273	-0.553

	LCYB	DLCYB	LCGB	DLCGB	RIB	DRIB	UB	DUB	
TAU (L)	-2.287 (0)	-8.134 (0)	-3.469 (0)	-6.932 (0)	-2.769 (0)	-9.059 (0)	-1.116 (1)	-2.990 (0)	
LMSC1	1.697	0.353	2.455	0.904	0.979	0.023	1.455	2.095	
SBIC	-8.234	-8.190	-8.508	-8.231	-7.384	-7.361	-0.444	-0.507	
PHI	2.985	32.270	6.064	28.333	6.094	39.790	3.704	4.466 (1)	
t-trend	-0.881	0.360	0.577	2.025	1.964	-0.031	2.443	0.457	

TABLE 3.2 (continued): ADF Tests (Canada)

	LCC	DLCC	LYC	DLYC	LGC	DLGC	LPC	DLPC	DDLPC
TAU (L)	-1.374 (1)	-3.528 (0)	-2.091 (1)	-3.533 (0)	-1.784 (1)	-3.819 (0)	-1.498 (1)	-2.020 (1)	-4.410 (0)
LMSC1	0.061	0.046	0.391	0.206	0.114	0.005	2.502	0.013	0.510
SBIC	-7.339	-7.383	-7.197	-7.171	-7.167	-7.174	-8.278	-8.345	-8.326
PHI	1.255	6.992	5.323 (0)	9.374	1.774	8.893	3.506	2.170	10.687
t-trend	0.803	-1.180	-1.708	-2.196	0.650	-1.590	2.126	-0.586	-1.258

	LCYC	DLCYC	LCGC	DLCGC	RIC	DRIC	UC	DUC	
TAU (L)	-2.163 (0)	-5.271 (0)	-3.020 (1)	-4.662 (0)	-2.144 (0)	-8.560 (0)	-1.607 (1)	-4.283 (0)	
LMSC1	0.327	0.174	1.922	0.043	1.584	0.220	1.111	2.453	
SBIC	-8.192	-8.067	-9.139	-8.990	-8.048	-8.043	0.046	0.022	
PHI	2.578	16.686	5.195	11.327	4.349	35.749	4.654	8.901	
t-trend	0.736	1.868	1.100	0.974	1.930	0.373	2.510	0.087	

TABLE 3.2 (continued): ADF Tests (Denmark)

	LCD	DLCD	LYD	DLYD	LGD	DLGD	LPD	DLPD	DDLPD
TAU (L)	-2.196 (0)	-4.610 (0)	-0.944 (0)	-5.588 (1)	-2.918 (0)	-5.048 (0)	-1.805 (1)	-1.566 (0)	-6.483 (0)
LMSC1	0.229	0.090	0.028	0.709	0.211	0.427	0.853	0.062	0.760
SBIC	-7.044	-6.943	-6.301	-6.280	-7.650	-7.445	-7.919	-7.923	-7.866
PHI	5.113	11.298	5.670	15.503	7.074	16.538	2.044	2.670	23.735
t-trend	2.200	-1.101	3.194	-0.619	2.164	-2.173	0.920	-1.661	-1.718
	LCYD	DLCYD	LCGD	DLCGD	RID	DRID	UD	DUD	
TAU (L)	-2.701 (1)	-4.485 (0)	-1.174 (0)	-4.882 (0)	-2.244 (0)	-7.889 (0)	0.223 (0)	-4.469 (0)	
LMSC1	0.012	1.519	0.759	0.011	2.486	0.502	2.298	0.045	
SBIC	-6.809	-6.705	-7.967	-7.943	-7.363	-7.380	-0.100	-0.188	
PHI	3.545	10.279	5.570 (1)	11.758	6.043	30.755	3.501	10.184	
t-trend	-0.133	-0.809	-3.033	0.488	2.499	0.632	2.635	0.790	

TABLE 3.2 (continued): ADF Tests (Finland)

	LCFI	DLCFI	LYFI	DLYFI	LGFI	DLGFI	LPFI	DLPFI	DDLPFI
TAU (L)	-2.215 (1)	-3.210 (0)	-2.795 (1)	-3.107 (0)	-2.239 (2)	-4.235 (1)	-1.313 (1)	-2.502 (1)	-5.589 (1)
LMSC1	1.366	0.139	0.666	0.152	0.081	0.114	1.313	0.490	0.464
SBIC	-7.010	-6.969	-7.020	-6.904	-7.036	-6.988	-7.217	-7.241	-7.171
PHI	2.744	7.497	8.244 (0)	9.221	2.428	12.400	2.187	2.624 (0)	16.563
t-trend	0.799	-1.952	-0.664	-2.651	0.063	-2.182	1.601	-1.071	-1.205
	LCYFI	DLCYFI	LCGFI	DLCGFI	RIFI	DRIFI	UFI	DUFI	
LAGS	0	0	1	0	0	0	2	1	
TAU (L)	-2.226 (0)	-6.180 (0)	-3.218 (1)	-4.780 (0)	-2.375 (0)	-6.341 (0)	-0.167 (2)	-5.643 (1)	
LMSC1	0.080	0.329	0.179	3.696	0.305	0.028	0.274	0.181	
SBIC	-7.673	-7.545	-8.454	-8.275	-7.234	-7.078	0.230	0.129	
PHI	6.193	18.765	5.275	11.375	3.205	19.653	13.453 (1)	16.551	
t-trend	2.567	0.480	-0.625	-0.591	0.895	0.374	4.259	1.062	

TABLE 3.2 (continued): ADF Tests (France)

	LCFR	DLCFR	LYFR	DLYFR	LGFR	DLGFR	LPFR	DLPFR	DDLPFR
TAU (L)	-6.751 (0)	-2.873 (0)	-6.546 (0)	-3.391 (0)	-6.207 (0)	-3.294 (0)	-1.060 (1)	-1.344 (0)	-5.671 (0)
LMSC1	0.270	0.079	1.135	0.009	0.352	0.912	0.385	0.790	0.348
SBIC	-9.018	-8.568	-8.399	-7.851	-8.557	-8.080	-7.816	-7.883	-7.852
PHI	22.461	15.785	20.871	18.999	18.697	16.561	3.632	1.118	16.103
t-trend	-0.550	-4.342	-0.287	-4.462	-0.133	-4.124	2.442	-0.677	-0.720
	LCYFR	DLCYFR	LCGFR	DLCGFR	RIFR	DRIFR	UFR	DUFR	
TAU (L)	-1.364 (0)	-5.921 (0)	-2.995 (0)	-6.113 (0)	-1.871 (0)	-6.694 (1)	0.522 (1)	-3.559 (0)	
LMSC1	0.580	4.243	0.782	9.264	0.045	0.177	0.004	0.035	
SBIC	-8.691	-8.637	-9.198	-8.964	-8.197	-8.267	-1.339	-1.432	
PHI	3.211	19.821	5.411	19.288	2.639	21.712	4.915	7.474	
t-trend	2.091	1.655	1.293	1.047	1.304	0.082	3.080	1.387	

TABLE 3.2 (continued): ADF Tests (Germany)

	LCGER*	DLCGER*	LYGER*	DLYGER*	LGGER*	DLGGER*	LPGER	DLPGER	DDLPGER
TAU (L)	-2.104 (1)	-4.768 (0)	-2.247 (0)	-4.985 (0)	-2.655 (0)	-7.178 (0)	-0.828 (2)	-3.413 (1)	-4.264 (0)
LMSC1	5.148	2.410	3.418	0.184	0.468	0.001	0.032	0.002	2.168
SBIC	-7.876	-7.844	-7.209	-7.188	-7.610	-7.493	-8.986	-9.066	-8.857
PHI	2.549	13.763	2.620	13.468	3.521	29.087	1.387	5.942	9.252
t-trend	0.844	-1.794	0.550	-1.269	0.419	-1.778	1.437	-0.662	-0.750
	LCYGER	DLCYGER	LCGGER*	DLCGGER*	RIGER	DRIGER	UGER	DUGER	
TAU (L)	-1.726 (0)	-6.724 (0)	-1.342 (0)	-6.671 (0)	-1.487 (2)	-8.729 (1)	-1.181 (1)	-3.954 (0)	
LMSC1	0.385	0.799	0.265	0.667	0.125	0.044	0.001	0.066	
SBIC	-8.557	-8.496	-8.390	-8.336	-8.252	-8.284	-0.497	-0.556	
PHI	1.568	22.415	1.560	22.218	3.142	36.947	4.430	7.649	
t-trend	0.474	-0.645	1.141	-0.744	1.966	0.151	2.682	0.302	

TABLE 3.2 (continued): ADF Tests (Greece)

	LCGRE	DLCGRE	LYGRE	DLYGRE	LGGRE	DLGGRE	LPGRE	DLPGRE	DDLPGRE
TAU (L)	-3.262 (1)	-2.950 (0)	-3.939 (0)	-4.342 (0)	-4.782 (0)	-3.765 (0)	0.077 (1)	-1.730 (0)	-5.436 (0)
LMSC1	2.381	1.002	0.107	0.414	0.009	2.335	0.532	0.389	1.712
SBIC	-7.710	-7.524	-6.507	-6.198	-7.003	-6.640	-6.587	-6.689	-6.604
PHI	12.693 (0)	10.877	7.774	18.057	12.143	17.762	3.377	1.990	14.860
t-trend	-1.304	-3.246	-0.585	-3.369	-1.117	-3.903	2.598	0.995	-0.749
	LCYGRE	DLCYGRE	LCGGRE	DLCGGRE	RIGRE	DRIGRE	UGRE	DUGRE	
TAU (L)	-2.871 (0)	-6.689 (0)	-2.272 (0)	-7.056 (0)	-2.065 (0)	-5.770 (1)	-1.667 (1)	-3.603 (1)	
LMSC1	1.919	0.187	0.622	0.007	0.681	0.017	2.847	0.007	
SBIC	-7.482	-7.281	-7.836	-7.732	-6.899	-6.814	-0.863	-0.915	
PHI	4.148	26.502	5.092	32.643	2.428	19.845	2.905	7.356	
t-trend	-0.493	2.020	2.117	2.603	0.799	1.909	1.694	1.232	

TABLE 3.2 (continued): ADF Tests (Iceland)

	LCIC	DLCIC	LYIC	DLYIC	LGIC	DLGIC	LPIC	DLPIC	DDLPIC
TAU (L)	-1.586 (0)	-4.933 (1)	-1.724 (2)	-4.757 (1)	-1.446 (1)	-3.984 (0)	-0.930 (1)	-1.840 (0)	-7.040 (0)
LMSC1	1.734	0.010	1.870	0.989	0.002	0.112	1.061	0.421	0.324
SBIC	-5.441	-5.447	-5.115	-5.126	-6.324	-6.363	-4.624	-4.699	-4.643
PHI	1.247	14.112	3.181 (2)	12.672	1.640	8.764	2.095	1.804	25.966
t-trend	0.224	-1.624	1.804	-1.416	1.084	-1.200	1.807	-0.543	-1.244
	LCYIC	DLCYIC	LCGIC	DLCGIC	RIIC	DRIIC	UIC	DUIC	
TAU (L)	-1.483 (0)	-6.637 (0)	-2.802 (0)	-6.170 (1)	-2.928 (1)	-5.982 (1)	0.646 (0)	-4.394 (0)	
LMSC1	0.059	0.231	0.291	1.797	0.539	1.768	2.167	3.137	
SBIC	-6.479	-6.434	-6.612	-6.458	-5.162	-5.051	-1.079	-1.137	
PHI	5.185	21.522	3.981	19.783	4.056(0)	17.827	1.950 (1)	12.577(1)	
t-trend	2.779	0.376	0.533	-1.109	-0.072	0.685	1.914	1.589	

TABLE 3.2 (continued): ADF Tests (Ireland)

	LCIR	DLCIR	LYIR	DLYIR	LGIR	DLGIR	LPIR	DLPIR	DDLPIR
TAU (L)	-1.017 (0)	-4.431 (0)	-1.288 (0)	-6.411 (0)	-0.194 (0)	-4.459 (0)	-1.584 (1)	-1.655 (0)	-4.955 (0)
LMSC1	2.532	1.785	0.591	0.074	1.753	0.621	1.141	2.303	1.577
SBIC	-6.804	-6.839	-6.549	-6.510	-7.362	-7.414	-7.070	-7.096	-7.045
PHI	4.439 (1)	9.669	2.749	20.752	2.742	9.648	3.013	2.112	12.935
t-trend	2.778	-0.431	1.924	-0.856	2.336	0.090	1.826	-1.204	-1.086
	LCYIR	DLCYIR	LCGIR	DLCGIR	RIIR	DRIIR	UIR	DUIR	
LAGS	0	0	0	0	0	2	1	0	
TAU (L)	-3.602 (0)	-10.358 (0)	0.073 (0)	-6.969 (0)	-2.133 (0)	-5.597 (2)	-1.187 (1)	-3.309 (0)	
LMSC1	4.356	0.469	1.252	0.116	0.322	0.265	0.850	1.543	
SBIC	-7.331	-7.304	-7.237	-7.272	-6.912	-6.822	0.209	0.151	
PHI	3.073 (1)	53.181	4.588	24.331	3.488	15.880	3.703	5.310	
t-trend	0.680	0.738	-3.028	-0.797	1.501	0.846	2.406	-0.010	

TABLE 3.2 (continued): ADF Tests (Italy)

	LCIT	DLCIT	LYIT	DLYIT	LGIT	DLGIT	LPIT	DLPIT	DDLPIT
TAU (L)	-3.946 (0)	-3.532 (0)	-2.745 (1)	-2.928 (0)	-3.756 (0)	-4.185 (0)	-1.439 (1)	-1.762 (0)	-5.030 (0)
LMSC1	1.493	1.379	0.405	0.170	0.124	0.028	1.001	2.051	1.572
SBIC	-7.699	-7.520	-7.547	-7.437	-7.677	-7.431	-7.237	-7.276	-7.209
PHI	7.754	10.801	3.956	7.693	7.254	14.193	3.407	2.029	13.878
t-trend	0.527	-2.627	0.705	-2.369	0.761	-2.729	2.125	-0.978	-1.350
	LCYIT	DLCYIT	LCGIT	DLCGIT	RIIT	DRIIT	UIT	DUIT	
TAU (L)	-0.020 (0)	-5.369 (0)	-0.936(0)	-5.575(0)	-1.663 (0)	-6.126 (0)	-1.031 (1)	-3.667 (0)	
LMSC1	0.152	0.014	0.058	0.072	0.278	1.303	0.034	0.017	
SBIC	-7.778	-7.785	-8.716	-8.693	-7.788	-7.709	-1.395	-1.464	
PHI	3.058	15.715 (1)	7.955 (1)	15.200	3.151	19.037	8.519	7.099	
t-trend	2.473	1.404	3.809	-0.365	1.828	0.889	3.936	0.909	

TABLE 3.2 (continued): ADF Tests (Japan)

	LCJ	DLCJ	LYJ	DLYJ	LGJ	DLGJ	LPJ	DLPJ	DDLPJ
TAU (L)	-6.866 (0)	-2.880 (0)	-3.109 (1)	-2.135 (0)	-6.535 (0)	-2.681 (0)	-2.047 (1)	-2.262 (0)	-6.490 (0)
LMSC1	0.162	3.138	0.590	0.259	1.571	1.750	0.017	0.008	2.056
SBIC	-7.770	-7.302	-7.700	-7.537	-7.244	-6.887	-7.168	-7.146	-7.016
PHI	23.257	12.522	4.879	5.524	21.705	9.524	2.093	4.562	21.303
t-trend	0.576	-3.686	0.551	-2.413	0.935	-3.149	0.338	-1.899	-0.878
	LCYJ	DLCYJ	LCGJ	DLCGJ	RIJ	DRIJ	UJ	DUJ	
TAU (L)	-1.037 (0)	-4.808 (0)	-2.936 (0)	-4.962 (0)	-2.865 (0)	-8.526 (0)	-0.881 (1)	-3.608 (0)	
LMSC1	1.988	0.807	0.301	0.005	2.633	1.491	0.617	0.892	
SBIC	-8.260	-8.272	-8.160	-7.945	-7.610	-7.529	-3.096	-3.174	
PHI	1.765	11.820	6.458	13.691	4.393	36.557	8.773	8.100	
t-trend	1.552	0.850	-1.901	1.417	0.813	0.904	4.049	1.601	

TABLE 3.2 (continued): ADF Tests (the Netherlands)

	LCNT	DLCNT	LYNT	DLYNT	LGNT	DLGNT	LPNT	DLPNT	DDLPNT
TAU (L)	-2.608 (1)	-2.486 (0)	-4.649 (0)	-4.155 (0)	-3.576 (0)	-4.097 (0)	-1.762 (1)	-1.903 (0)	-6.492 (0)
LMSC1	0.530	0.090	0.179	0.188	0.812	0.542	0.284	0.025	0.522
SBIC	-8.180	-8.089	-7.621	-7.224	-7.881	-7.667	-8.246	-8.255	-8.166
PHI	3.970	5.232	10.504	17.368	6.853	12.113	2.306	2.981	21.605
t-trend	1.057	-1.941	0.173	-3.437	0.971	-2.295	1.209	-1.486	-1.014
	LCYNT	DLCYNT	LCGNT	DLCGNT	RINT	DRINT	UNT	DUNT	
TAU (L)	-3.482 (0)	-7.426 (0)	-1.927 (0)	-5.738 (0)	-2.190 (0)	-6.922 (0)	-0.944 (2)	-3.819 (1)	
LMSC1	0.864	0.458	0.144	4.358	0.165	1.738	0.012	0.007	
SBIC	-8.526	-8.270	-8.335	-8.229	-7.537	-7.439	-0.408	-0.481	
PHI	6.346	28.064	1.816	17.508	4.404	23.238	5.041 (1)	7.068	
t-trend	0.826	0.997	0.164	-1.244	1.905	0.083	2.863	-0.062	

TABLE 3.2 (continued): ADF Tests (Norway)

	LCNR	DLCNR	LYNR	DLYNR	LGNR	DLGNR	LPNR	DLPNR	DDLPNR
TAU (L)	-1.850 (0)	-4.546 (0)	-2.504 (0)	-3.661 (0)	-2.118 (0)	-3.773 (0)	-0.896 (1)	-2.101 (0)	-5.946 (1)
LMSC1	0.950	0.081	2.904	0.144	2.912	0.006	0.317	0.028	0.062
SBIC	-7.148	-7.100	-7.850	-7.841	-8.098	-8.118	-7.637	-7.714	-7.615
PHI	2.224	11.272	3.265 (1)	7.543	2.348 (1)	8.135	1.805	2.401	20.422
t-trend	1.011	-1.241	1.860	-1.218	1.529	-1.313	1.662	-0.679	-1.770
	LCYNR	DLCYNR	LCGNR	DLCGNR	RINR	DRINR	UNR	DUNR	
TAU (L)	-2.182 (0)	-5.740 (0)	-0.352 (0)	-6.446 (0)	-1.550 (0)	-6.882 (0)	2.898 (2)	-3.731 (1)	
LMSC1	0.759	0.021	0.558	3.023	0.450	2.222	4.060	0.002	
SBIC	-7.366	-7.232	-7.536	-7.544	-7.342	-7.303	-1.046	-0.907	
PHI	2.502	15.986	7.125	20.419	2.710	23.778	7.219	14.228	
t-trend	0.580	-0.103	-3.752	-0.492	1.697	0.819	2.229	3.231	

TABLE 3.2 (continued): ADF Tests (Spain)

	LCSP	DLCSP	LYSP	DLYSP	LGSP	DLGSP	LPSP	DLPSP	DDLPSP
TAU (L)	-2.208 (1)	-3.597 (0)	-4.307 (1)	-3.749 (0)	-3.605 (1)	-3.248 (0)	-0.876 (1)	-1.504 (0)	-5.833 (0)
LMSC1	8.783	2.530	0.306	1.566	0.002	2.531	0.120	0.001	3.390
SBIC	-6.847	-6.807	-7.315	-6.960	-7.670	-7.431	-7.165	-7.243	-7.179
PHI	2.838	8.351	10.757	13.753	8.698	9.815	4.086	1.215	17.049
t-trend	0.910	-1.728	1.498	-3.120	1.849	-2.669	2.694	-0.472	-0.736
	LCYSP	DLCYSP	LCGSP	DLCGSP	RISP	DRISP	USP	DUSP	
TAU (L)	-4.863 (0)	-9.121 (1)	-3.796 (0)	-7.628 (0)	-2.458 (0)	-8.157 (0)	-0.452 (1)	-2.460 (0)	
LMSC1	2.820	5.941	4.562	11.424	2.248	0.857	2.709	2.844	
SBIC	-8.181	-7.900	-8.435	-8.124	-6.424	-6.359	0.193	0.097	
PHI	27.902	47.099	8.126	29.603	3.530	32.269	4.201	3.286	
t-trend	4.376	1.944	-1.260	1.002	1.009	-0.066	2.855	0.771	

TABLE 3.2 (continued): ADF Tests (Sweden)

	LCSWE	DLCSWE	LYSWE	DLYSWE	LGSWE	DLGSWE	LPSWE	DLPSWE	DDLPSWE
TAU (L)	-2.436 (1)	-3.479 (0)	-2.796 (2)	-3.970 (1)	-4.681 (0)	-3.206 (0)	-0.384 (1)	-2.547 (0)	-8.148 (0)
LMSC1	0.162	0.184	0.153	0.011	2.850	0.361	1.553	1.084	0.052
SBIC	-7.792	-7.724	-7.919	-7.796	-8.067	-7.935	-7.456	-7.553	-7.489
PHI	3.014	8.888	5.816 (1)	10.238	10.717	9.078	1.814	3.154	34.982
t-trend	0.484	-2.102	1.922	-1.869	-0.339	-2.499	1.863	-0.121	-1.361
	LCYSWE	DLCYSWE	LCGSWE	DLCGSWE	RISWE	DRISWE	USWE	DUSWE	
TAU (L)	-1.380 (0)	-4.825 (0)	-1.620 (0)	-4.369 (0)	-1.559 (1)	-9.864 (0)	-3.111 (1)	-4.622 (1)	
LMSC1	3.343	0.822	2.277	1.489	0.158	0.057	1.627	0.841	
SBIC	-7.676	-7.656	-8.203	-8.198	-7.222	-7.251	-1.080	-1.043	
PHI	2.544 (1)	11.449	5.354 (1)	10.879 (1)	2.123	47.798	10.845	11.220	
t-trend	1.357	-0.434	-2.791	0.778	1.326	0.561	3.075	1.022	

TABLE 3.2 (continued): ADF Tests (Switzerland)

	LCSWZ	DLCSWZ	LYSWZ	DLYSWZ	LGSWZ	DLGSWZ	LPSWZ	DLPSWZ	DDLPSWZ
TAU (L)	-2.713 (1)	-2.800 (0)	-2.119 (1)	-3.281 (0)	-3.319 (0)	-3.703 (0)	-1.255 (2)	-3.831 (1)	-4.819 (0)
LMSC1	2.837	0.649	0.004	0.163	3.127	0.261	1.389	1.146	1.698
SBIC	-8.491	-8.385	-7.631	-7.601	-7.508	-7.440	-8.141	-8.193	-7.917
PHI	3.713	7.237	3.080	7.016	4.781 (1)	8.516	3.751 (1)	8.026	12.320
t-trend	0.490	-2.356	1.260	-1.646	1.871	-1.624	2.243	-1.121	-1.115
	LCYSWZ	DLCYSWZ	LCGSWZ	DLCGSWZ	RISWZ	DRISWZ	USWZ	DUSWZ	
LAGS	0	0	0	1	0	0	2	1	
TAU (L)	-1.118 (0)	-4.573 (0)	-1.411 (0)	-5.278 (1)	-2.626 (0)	0.115 (0)	0.136 (2)	-6.132 (1)	
LMSC1	3.634	0.160	1.428	0.059	0.583	44.967	0.105	0.050	
SBIC	-8.478	-8.501	-8.425	-8.426	-7.717	-7.531	0.253	0.152	
PHI	1.886 (1)	10.141	1.157	13.782	4.186	19.220	1.969	20.887	
t-trend	-1.274	-0.058	0.602	-0.556	1.180	0.951	1.979	1.568	

TABLE 3.2 (continued): ADF Tests (the UK)

	LCUK	DLCUK	LYUK	DLYUK	LGUK	DLGUK	LPUK	DLPUK	DDLPUK
TAU (L)	-0.299 (1)	-3.835 (0)	-0.567 (2)	-5.007 (1)	-1.156 (2)	-5.091 (1)	-1.131 (1)	-1.900 (0)	-5.419 (0)
LMSC1	2.487	2.376	0.356	0.550	0.031	0.085	0.345	0.841	0.481
SBIC	-7.446	-7.545	-7.315	-7.407	-7.585	-7.645	-7.050	-7.113	-7.013
PHI	5.400	7.143	7.708 (1)	12.223	6.929 (1)	13.281	2.617	2.075	15.256
t-trend	3.269	0.133	3.847	-0.300	3.538	-0.896	1.960	-0.764	-1.039
	LCYUK	DLCYUK	LCGUK	DLCGUK	RIUK	DRIUK	UUK	DUUK	
TAU (L)	-3.193 (1)	-4.372 (0)	-0.449 (0)	-4.710 (0)	-2.179 (0)	-7.173 (0)	-1.003 (2)	-4.962 (1)	
LMSC1	0.084	0.123	1.817	1.002	0.458	1.288	0.203	0.247	
SBIC	-8.502	-8.327	-8.424	-8.451	-7.172	-7.086	0.014	-0.056	
PHI	5.884	9.563	3.388	12.327	2.769	25.042	1.786	11.954	
t-trend	1.199	0.611	2.557	1.370	0.902	0.277	1.587	-0.180	

TABLE 3.2 (continued): ADF Tests (the USA)

	LCUS	DLCUS	LYUS	DLYUS	LGUS	DLGUS	LPUS	DLPUS	DDLPUS
TAU (L)	-1.5212 (0)	-4.1475 (0)	-2.1653 (0)	-4.7725 (0)	-1.4315 (0)	-4.6764 (0)	-0.8693 (1)	-1.7486 (0)	-5.097 (0)
LMSC1	3.3246	3.6065	0.154	0.1608	1.4942	3.6241	1.0362	1.5521	2.5549
SBIC	-8.0487	-8.0967	-8.1267	-8.0279	-7.6017	-7.5926	-8.449	-8.5272	-8.453
PHI	2.1934 (1)	8.9302	2.3424	13.6346	2.5226	11.1128	3.3112	1.6165	13.353
t-trend	1.7314	-0.8807	0.348	-1.7511	1.6969	-0.783	2.3989	-0.4954	-0.921
	LCYUS	DLCYUS	LCGUS	DLCGUS	RIUS	DRIUS	UUS	DUUS	
TAU (L)	-0.719 (0)	-6.4449 (0)	-0.7325 (0)	-5.5636 (0)	-2.3824 (0)	-6.2831 (0)	-1.9758 (0)	-5.0091 (0)	
LMSC1	0.2349	1.1221	0.1033	0.0979	0.3022	0.1468	3.0215	1.5384	
SBIC	-8.9645	-8.9586	-8.9833	-8.9684	-8.375	-8.2223	-0.0084	0.0781	
PHI	2.8091	21.3331	2.5444	15.008	2.9061	19.2068	3.78198 (1)	12.2325	
t-trend	2.2445	1.0283	2.1203	0.0061	0.5129	-0.2462	1.126	-0.2758	

TABLE 3.2 notes. "TAU" is the test statistic for the difference stationarity hypothesis without a trend in the test equation (critical value is -2.95), where L, given in brackets after this statistic, denotes the number of lagged dependent variables in the test equation. "PHI" represents the test for stationarity around a linear trend (critical value is 7.306), where "t-trend" denotes the t-ratio of the time trend incorporated in the test equation. When L is different when the trend is included to when it is excluded, it is reported in brackets after "PHI". "LMSC1" denotes the lagrange multiplier test for first order autocorrelation (critical value is 3.84) and "SBIC" is Schwartz's Bayesian Information Criterion. In the German tests an asterix denotes the use of a *spike* dummy variable, which is unity in 1991 and zero otherwise, in the ADF test equation. We do not adjust critical values to account for this variable.

are unambiguously I(1) for all countries while unemployment is generally I(1).[14]

Approximately fifteen percent of the results deviate from the general inference that the logs of consumption, income and GDP are I(1), prices are I(2) and the APCs are nonstationary.[15] Such deviations may be due to the low power of the ADF test and the small sample size used - these anomalies may be explained when we move from univariate to multivariate modelling. Thus, our general inference, which is consistent with prior beliefs, visual inspection of the data and previous empirical findings, is presented as the starting point for our subsequent multivariate analysis.

3.4 Conclusion

This chapter has addressed five issues using data for twenty OECD countries. Whether consumption is smooth relative to income; if there outliers and breaks in the data; whether consumption and income appear to move together through time; If GDP a good proxy for disposable income; and the series' orders of integration.

The evidence suggests that consumption is smooth relative to income for some, but not all, countries. Thus, the smoothing postulate of the LCH does not hold for some countries. This may be due to the durable component in the total consumption expenditure measure used and/or the need to relax some of the stronger assumptions of the pure LCH such as perfect capital markets,

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[14] Strictly, the unemployment rate is trend stationary for Finland, Italy, Japan and Sweden and at least I(2) for Spain. However, this finding of trend stationarity is for the sample and one would not expect linear trends to continue forever because this implies that unemployment will breach one of its boundaries of 0% or 100%. Given this, and that its use is restricted to being a potential instrument, we assume all unemployment rates are I(1).

[15] The deviations from the general inference are that consumption *growth* is trend stationary for France, Greece, Japan and Switzerland, while the Netherlands' consumption is at least I(2). Income is at least I(2) for Austria and Japan while income is trend stationary for the UK and income *growth* is trend stationary for Italy. GDP is at least I(2) for Austria and trend stationary for Sweden while its *growth* rate is trend stationary for Japan. Prices are I(1) for Switzerland and Germany. The consumption-income ratio is stationary for Ireland, the Netherlands, Spain and the UK. The consumption-GDP ratio is stationary for Canada, Finland and Spain.

forward looking behaviour, no income uncertainty and the needs of families. In later chapters we will investigate the role of durability, liquidity constraints and forward and backward looking behaviour in formal models.

The individual series for each country are subject to a wide range of outliers and structural breaks. Most series suffer outliers and breaks during the 1950s, probably reflecting post-war adjustment. Since all regressions are estimated over the period 1960/1 to 1994 these should not have a substantive impact upon our modelling. In general, inflation soars in the 1970s, coinciding with a fall in many countries' consumption, income and GDP log-levels and growth rates. This motivates the consideration of a direct (or indirect) relationship between consumption and inflation in our modelling. A boom and bust cycle during the mid-late 1980s and early 1990s appears to have occurred in the UK and Nordic countries with their APCs nearing or exceeding unity.

Inspection of data plots suggest that consumption, income and GDP tend to move together through time indicating the possibility of cointegration and/or cobreaking. Cointegration between consumption, income and inflation is tested in Chapter 4.

Private disposable income is a much closer approximation of the concepts of income (labour income and Hicksian income) typically employed in consumption functions than GDP. The former is, therefore, regarded as more appropriate than the latter. Evidence suggests a low correlation between basic data transformations using income and GDP. In particular, the simple correlation between the consumption-income and consumption-GDP ratios is generally very low and sometimes negative. Thus, consumption functions using disposable income and GDP may yield quite divergent inferences. One may, therefore, be advised to prefer disposable income (when available) to GDP when building consumption functions of the household or private sector. Unlike many previous studies considering a broad range of countries we employ the most appropriate available proxy for income, disposable income, in our subsequent empirical analyses.

The series' orders of integration are found to be heterogeneous across countries. This may be due to the low power of the ADF test employed, the small sample size, the presence of moving average errors and/or incorrect choice of order of augmentation in the test equation and outliers/breaks in the data. Nevertheless, the general inference is consistent with the conventional wisdom that the logs of consumption, income and GDP are I(1). The APCs are nonstationary for most countries but may be I(0) for some. The order of integration of the (log of) prices is I(2) for the majority of countries but I(1) for some. For consistency of model specification across countries we assume the APC is I(1) and prices are I(2) for all countries. These general orders of integration will provide the starting point for our multivariate analysis where, it is hoped, these anomalies will be resolved.

APPENDIX 3.1A: Data Definitions, Construction, Sources and Coverage

The data to be analysed are annual observations (1955-1994) for country * on the following series:

C* Total private consumers' expenditure in current (C*) and 1990 (RC*) prices. The primary source is OECD National Accounts, volume 1 (Main Aggregates) with some of the earlier observations (1950s) obtained from UN National Accounts and later data (1990s) from, mainly, OECD Quarterly National Accounts 96/1, volume 1, and also OECD Economic Outlook 6/96.

Y* Private (households and non-profit institutions) disposable income in current prices is total current receipts minus property income disbursements minus direct taxes minus transfers to central government minus other transfers. The main source is UN National Accounts with latest observations taken from OECD Economic Outlook: this is derived from the net household saving ratio, $S^*=\{[(Y^*-C^*)/Y^*]x100\}$, using the formula $Y=\{C^*/[1-(S^*/100)]\}$. Missing observations for Greece, Japan and Switzerland are taken from OECD National Accounts, for Denmark from OECD Economic Outlook and for Ireland from Ireland's National Accounts 1975-1981. For Iceland the disposable income series were unavailable in any of these publications, however, two real per-capita household disposable income series, official and model series, were very kindly provided by Thorarinn Petursson (Bank of Iceland) - official source is National Economic Institute: *Historical Statistics*, Reykjavik, September 1995. We use the model series, which is preferred by the Bank of Iceland. This series was only available from 1962 onwards - GDP was used for earlier observations. For Norway a real (1970) private disposable income series was kindly provided by Kari H Eika (Bank of Norway), which was available over the period 1962-1978. Nominal income was available from UN National Accounts and, indirectly, from OECD Economic Outlook over the period 1975-1994. The implied consumer price index was used to make the latter data real, 1990 prices, and so the former was spliced to it. The series was then made nominal by multiplying it with the price index and spliced to nominal GDP to obtain the pre-1962 data.

G* Current (G) and 1990 (RG) price GDP was taken from OECD National Accounts, volume 1 (Main Aggregates). Latest observations were taken from either OECD Quarterly National Accounts, volume 1, or OECD Economic Outlook.

POP* Population (millions of people) were obtained from International Monetary Fund International Financial Statistics (IMFIFS) line 99z. Most of the data, to 1990, was obtained using Manchester's Data Archive. For Belgium the 1990-1994 observations were obtained from Eurostatistics July 1996 edition.

I* Short term nominal interest rates, measured in percent per annum. The majority of the data comes from IMFIFS. The various measures available are the discount rate (line 60), the money market rate (line 60b), the Treasury bill rate (line 60c), the deposit rate (line 60l) and the long term government bond yield (line 61). The definitions used for any particular country are subject to data constraints, however, our order of preference is 60b, 60c, 60l, 60 and, in the absence of short term rates, 61. For Belgium (60b), Canada (60c), France (60b), Germany (60b), the UK (60c) and the USA (60b), the single measure (specified in brackets) is available over the full sample period: 1955-1994. The money market rate (60b) is spliced with the discount rate (60) for Austria (1955-1966), Denmark (1955-1971), Finland (1955-1977), Ireland (1955-1970), Italy (1955-1968), Japan (1955-1956), the Netherlands (1955-1959), Spain (1955-1973), and Switzerland (1955-1968). The period where the discount rate is used is specified in brackets. For Spain the money market rate is taken from OECD Main Economic Indicators (MEI). This is the call market rate and is the same as line 60b in IMFIFS but is available for a few more years in the former source. The Treasury bill rate is spliced with the discount rate for Sweden (line 60 is used for the period 1955-1962) and with the government bond yield for Australia (line 61 is used for the period 1955-1968). The deposit rate (60l) is spliced with the discount rate for Greece (1955-1960), Iceland (1955-1972) and Norway (1955-1978). The period where the discount rate is used is given in brackets.

U* Standardised unemployment rates, measured in percentages, are available annually from 1955-1992 for all 20 countries, except Greece and Iceland, from Layard, Nickell and Jackman (1994). OECD MEI is used to obtain observations for 1993 and 1994 (and 1991 for some countries). Data for Greece, over the period 1960-1994, was kindly provided by Costas Milas. Data for 1951 and 1961 is available in OECD Labour Force Statistics. Linear interpolation is used to obtain data in the *low unemployment* 1951 to 1960 era. This provides data for the pre-1960 period, which is reserved for transformations and lags. Iceland's unemployment rate is only available over the period 1961-1994 and is obtained from various editions of MEI, ILO Yearbook of Labour Statistics, OECD Labour Force Statistics and UN Statistical Yearbook. It is assumed that Iceland's unemployment rate is 0.1% each year over the period 1955-1960 because it is this value from the period 1961-1966 and one would expect low unemployment post-war. Once again the majority of these observations are used for lags and transformations. The data for Austria, Denmark, Ireland and Sweden are unstandardised rates.

3.1A.1 Transformed Series

The raw data is transformed for economic analysis. The transformations applied are:

P* The implied (total) consumption price deflator, 1990~100. P= [(C/RC)x100].

LP* = ln(P*/100).

DLP* Consumer price inflation: DLP*=LP*-LP*(-1)

LC* Natural logarithm of real per-capita total consumer expenditure:

ln[(RC*)/(POP*x1000000)]. Multiplying population by one million gives population in persons rather than million of persons.

DLC* Consumption growth: DLC*=LC*-LC*(-1)

LY* Natural logarithm of real per-capita private disposable income: LY*=ln{[Y*/(P*/100)]/[POP*x1000000]}.

DLY* Income growth: DLY*=LY*-LY*(-1)

LG* Natural logarithm of real per-capita GDP: LG* = ln [(RG*)/(POP*x1000000)].

LCY* Natural logarithm of the consumption-income ratio: ln (C*/Y*).

LCG* Natural logarithm of the consumption-GDP ratio: ln (RC*/RG*). Since the price deflators of real consumption and real GDP are different, unlike consumption and income, real values are used to obtain the *real* consumption-GDP ratio.

DU* = U*-U*(-1)

RI* The natural logarithm of one plus the short term real interest rate is exactly defined as ln[1+(I*/100)]-DLP*. This is the measure we use.[16]

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[16] The variable defined in our rational expectations model, loosely termed the real interest rate, is $\ln(1+r_t)$, **(3.1A.1)**

$$\text{where, } (1+r_t) = \{1+(I_t/100)\}/\{1+[(P_t-P_{t-1})/P_{t-1}]\}$$

$$= \{1+(I_t/100)\}/\{(P_{t-1}+P_t-P_{t-1})/P_{t-1}\}$$

$$= \{1+(I_t/100)\}/\{P_t/P_{t-1}\}. \qquad \textbf{(3.1A.2)}$$

Substitution of **(3.1A.2)** into **(3.1A.1)** gives:

$$\ln(1+r_t) = \ln\{1+(I_t/100)\}/\{P_t/P_{t-1}\}$$

3.1A.2 OECD Countries

Prior to events in eastern bloc countries there were twenty-four OECD countries. We consider the twenty for which disposable income data of reasonable coverage is available. The particular OECD countries being analysed are: Australia ($m), Austria (Schillings m), Belgium (Francs m), Canada ($m), Denmark (Kroner m), Finland (Markkaa m), France (Francs m), Germany (Deutschmarks m), Greece (Drachma m), Iceland (Kroner b), Ireland (£m), Italy (Lire b), Japan (Yen b), Netherlands (Guilders m), Norway (Kroner m), Spain (Pesetas b), Sweden (Kroner m), Switzerland (Francs m), the UK (£m), and the USA ($m). The units of measurement of the monetary variables are given in brackets after the country they relate to; m denotes millions and b, billions. Due to data constraints regarding disposable income, Luxembourg, New Zealand, Portugal and Turkey are omitted from the analysis. Yugoslavia, which briefly joined the OECD, and recent members, Mexico, Korea and three transition economies from eastern Europe, are not

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$$= \ln[1+(I_t/100)] - \ln[P_t/P_{t-1}]$$

$$= \ln[1+(I_t/100)] - \Delta\ln P_t. \qquad \textbf{(3.1A.3)}$$

(3.1A.3) is the exact measure which we use in our study, denoted **RI***. If we define the nominal rate of interest as a proportion, thus:

$$R_t = (I_t/100), \qquad \textbf{(3.1A.4)}$$

then substitution of **(3.1A.4)** into **(3.1A.3)** yields:

$$\ln(1+r_t) = \ln(1+R_t) - \Delta\ln P_t. \qquad \textbf{(3.1A.5)}$$

The approximation,

$$R_t \approx \ln(1+R_t),$$

which is valid for small R_t, when substituted into **(3.1A.5)** gives:

$$\ln(1+r_t) \approx R_t - \Delta\ln P_t. \qquad \textbf{(3.1A.6)}$$

(3.1A.6) is a very convenient alternative approximation to **(3.1A.3)**, though is only valid for small values of R_t. For example, when $R_t = 0, 0.1, 0.2$ and 0.3, $\ln(1+R_t) = 0, 0.095, 0.182$ and 0.262. It generally provides an underestimate so, in general, reduces the variation in the proxy variable **(3.1A.6)** relative to its exact counterpart **(3.1A.3)**.

considered due to difficulties in obtaining data of reasonable coverage.

For Australia all series refer to the Fiscal Year: April 1 - March 31, rather than January 1 - December 31 (as is the case for all other countries).

APPENDIX 3.2A: Outliers and Breaks in Data

We provide a country by country overview of outliers and breaks, observed from Figure 3.1, in this appendix.

Australian income (growth) sharply falls then rises in 1967 and 1968.
Austrian consumption growth plummets in 1978, while income growth falls sharply in 1981.
Belgian consumption and income features a split trend from 1981 to 1985 (growths are negative), with a similar downward shift in 1993.
Canadian consumption (1982, 1991) and income (1983, 1991) series shift sharply downwards.
Danish consumption and income growth plummet in 1974.
Finnish consumption and income dramatically decline between 1991-1993, following the exceptional consumption growth in the late 1980s (APC exceeds unity in 1987) probably due to financial liberalisation - see Berg 1994.
French data seems quite stable. Income and consumption follow damped trends (these are often hard to distinguish from split trends).
German consumption and income feature clear downward shifts in their means due to reunification.
Greek inflation shifts up in 1974. Consumption, income and GDP exhibit split trends with their growth rates shifting down at the same time. Both measures of the APC decline, levelling off in 1974. The oil shock, conflict with Turkey over Cyprus and restoration of democracy may explain the change in behaviour from 1974.
Icelandic consumption and income are very volatile but follow very similar patterns. Both plummet in 1975 and 1983, and income soars in 1987.
Irish consumption and income shift downwards in 1982.
Italian consumption and income shift down in 1993.

Japanese consumption and income exhibit split trends with the series' growth slowing after 1974 - the end of this tiger's exceptional post-war growth.

The **Netherlands'** consumption and income shift down in 1981.

Norwegian consumption sharply rises in 1986 (the APC exceeds unity) and plummets in the period 1987-1989: being similar to the events in the other Nordic countries and the UK, probably all due to financial deregulation - see Berg 1994.

Spanish consumption and income's trends slow down in 1975 and speed up after 1985 only to slow again in 1993.

Swedish Consumption and income growth declines from 1977 to 1984 only to surge between 1986 and 1989, with consumption outstripping income so the APC exceeds unity. From 1990 consumption falls, despite strong income growth, and bottoms out in 1993, when income falls as well. Once again the turbulent pattern in the 1980s is likely due to financial deregulation - see Berg 1994.

Swiss consumption and income both shift down in 1975 and again in 1992.

The **UK's** consumption and income follow similar trends with pronounced cycles: a large (Barber) boom in the early 1970s followed by a deep slump between 1974 and 1977, then another accentuated boom over the period 1978-1979 and from 1980-1982 there is a deep slump. Since the trend in the data remains unchanged and consumption and income generally move together through time these are probably pronounced cycles. A new monetarist administration obtained office in 1979 with a deregulatory mandate. The main effects of financial deregulation occurred in the mid/late 1980s with a consumer boom (an economic miracle) between 1986 and 1988 (with the APC nearing unity). This was followed by a plummet in (particularly) consumption and income which bottomed out in 1991, as those now heavily exposed to debt found difficulty with repayments.

The **USA's** consumption and income follow similar paths. There are three deep slumps which shift consumption and income down in 1974, 1980 and 1991 - these may alternatively be interpreted as pronounced cycles around a damped trend.

CHAPTER 4

TESTING FOR COINTEGRATION BETWEEN CONSUMPTION, INCOME AND INFLATION

4.1 Introduction

This Chapter tests for cointegration between the natural logarithms of private total consumption (lnC) and private disposable income (lnY) and the rate of inflation (ΔlnP), for twenty OECD countries. The use of private disposable income, rather than GDP or national disposable income, for all twenty countries should allow enhanced inference relative to the few previous studies of OECD countries' consumer behaviour which use these broader measures. Consumption, income and inflation form the long run relationship utilised in Davidson *et al*'s (1978) [DHSY hereafter] pioneering work, and may, for example, be interpreted as approximating Ando and Modigliani's (1963) Life Cycle Hypothesis (LCH) formulation with naive income expectations and inflation proxying wealth effects. Given the prominence of the LCH which, at present, appears to be the starting point for the majority of modern analysis into consumer behaviour, it would seem desirable to employ a model based upon the three major variables underlying this theory: consumption, income and wealth. However, because data of reasonable coverage on wealth is unavailable for the majority of countries being considered, we use inflation to proxy these important asset effects. (Any non-wealth inflation effects will also be captured.) Thus, this standard consumption function provides a model that can be estimated for our sample of twenty OECD countries and can be thought of as approximating the fundamental features of the dominant LCH, affording valuable insights into the comparative consumer behaviours of the OECD economies.

There are other important influences on aggregate consumption that it would be desirable to investigate, such as, demography (see Berloffa 1997 and Horioka 1997), income uncertainty (see Carroll 1994, Church *et al* 1994 and Merrigan and Normandin 1996), interest rates (see Hall 1988 and Hahm 1998) and liquidity constraints (see Miles 1992, Berg 1994, Muellbauer 1994 and Bacchetta and Gerlach 1997). However, these factors are not considered in our analysis for three reasons. First, because there is a lack of reliable long consistent time series on such factors, especially age structure and liquidity constraints.[1] Second, whilst allowing for a certain degree of heterogeneity in model specification we also wish to consider similar models to facilitate meaningful cross-country comparison. This could be hindered if we considered a wide range of variables for each economy, especially if certain factors were retained in some countries' models and not others. Third, because we intend to apply the Johansen (1988, 1991 and 1995) procedure to each country using thirty-five time-series observations. Since this method is based upon a vector autoregression (VAR), degrees of freedom become increasingly scarce with the proliferation of variables entered endogenously in the equilibrium relation and, therefore, the efficiency of parameter estimates and the reliability of inference can be undermined.[2] Hence, in the face of data constraints, we focus on the three most important variables in the consumption relation (with inflation acting as a wealth proxy) to minimise the dimension of the VAR, so maximising the reliability of inference, and to maintain a standard set of variables to help the comparison of models across countries.[3]

The novelty of our investigation is in the use of *private* sector disposable income to measure income over the estimation period 1960-1994 for all twenty OECD countries and the development of country-specific models which are free from evident misspecification to identify each economy's long run consumption function. Both should yield superior estimates and inference relative to previous studies. We ensure that each country's model is free from evident

[1] We are only aware of analyses of four OECD countries using reliable time series on wealth, being Australia, Japan, the UK and the USA. Although we use demographic and liquidity constraints proxies in Chapter seven, these are representative averages for each country over the period 1960-1994. For example, the age structure variables are only available at five year intervals for most economies so do not constitute an annual time series.

[2] Greenslade, Hall and Henry (1998) highlight the potential efficiency and testing gains from reducing the dimension of the VAR in the context of the Johansen procedure. In particular, they emphasise the use of exogeneity assumptions.

[3] Horioka (1996), investigating Japanese consumption applied the Johansen procedure using a *maximum* of thirty-eight observations to a VAR system with three variables. Our specification features almost identical degrees of freedom so should provide valid inference.

misspecification by using country specific short run dynamics and dummy variables. We also allow for some degree of flexibility in the long run consumption function's specification by considering whether an intercept should be included in the cointegrating vector or not and, where appropriate, examining whether inflation should be omitted (by use of overidentification restrictions). We are not aware of any previous study which estimates equilibrium consumption functions for so many countries, from models facilitating such flexibility in specification or using as good a proxy for income as we do.

Using each country's preferred specification we investigate whether consumption, income and inflation form plausible long run consumption functions, whether the variables in any identified cointegrating relationships are statistically significant and what the sizes of the estimated long run coefficients are. We also test whether consumption is homogeneous of degree one in income and assess the issue of weak exogeneity.

The next section elaborates upon the theoretical interpretations of the postulated equilibrium. Section 4.3 tests for cointegration using the Johansen procedure. Hypothesis tests on the revealed cointegrating relations are conducted in section 4.4 and enable the selection of favoured long run consumption functions. Section 4.5 compares and contrasts these favoured long run consumption functions and, where possible, identifies general features of consumer behaviour across the OECD. Section 4.6 draws conclusions.

4.2 Specification of the Long Run Consumption Function

Although Keynes (1936) argued that there was a "psychological law" which made current income the primary determinant of consumption, subsequent theories which have come to underpin the majority of contemporary analysis of consumption emphasise the role of expected, rather than current, income - for example, the permanent income hypothesis (PIH) and LCH. However, relaxation of some of the stronger assumptions of the pure PIH-LCH to recognise, for examples, the presence of liquidity constraints, income uncertainty and information constraints limiting expectation formation, suggest that current income remains the primary determinant of consumption. Even under the rational expectations PIH-LCH (REPIH/RELCH), Campbell

(1987) demonstrates that consumption should cointegrate with current *disposable* income assuming that *labour* income is covariance stationary - see Jin 1993 p. 4. Further, the form of the LCH specified by Ando and Modigliani (1963), which assumes that expected income is proportional to current income, suggests that current consumption is a function of current income and wealth. Thus, one might expect current income to form a cointegrating relationship with consumption, either on its own or in conjunction with other factors. We therefore consider income as a prime candidate in forming an equilibrium with consumption.

In addition to income, we consider whether inflation influences equilibrium consumption. After the second World War, inflation was reasonably low in many of the OECD economies. However, inflation rapidly rose throughout the OECD with the OPEC oil shocks of the 1970s. This stimulated interest in the effects of inflation upon the macroeconomy and inflation emerged as a major argument in consumption functions. Three broad justifications for the influence of inflation upon consumption can be identified: anticipated effects, unanticipated effects and as a proxy for wealth.

Anticipated inflation effects include Juster and Watchel's (1972a, 1972b) argument that higher rates of inflation are associated with more volatile inflation which, assuming slow adjustment of *nominal* incomes, leads to increased instability (and therefore uncertainty) of *real* incomes, raising precautionary savings. Bulkley (1981) argues that discrete annual *nominal* wage increases leads to a sawtooth income profile which requires consumers to save more at the beginning of the wage contract period to compensate the inflation induced fall in *real* income throughout the year. When anticipated inflation is rising, the increased saving of those at the beginning of their contracts exceeds the dissaving of those at the end of their contracts, causing aggregate consumption to fall. Carruth *et al* (1996) argue that rising inflation may reduce consumption by approximating nominal interest rate movements through the Fisher effect. In contrast to these theories, which suggest a negative association between consumption and inflation, Springer (1977) argues that anticipated inflation induces upward revisions in expected inflation, causing an intertemporal substitution of planned future consumption to the present: consumption and inflation are positively associated. Similarly, Carruth and Henley (1992) suggest that there may be such a positive relation if increased inflation lowers real rates of return so inducing an intertemporal substitution of consumption to the present.

Hadjimatheou (1987) argues that inflation during the 1970s was mostly unanticipated. Deaton (1977) postulated a mechanism through which consumption is negatively related to *unanticipated* inflation. Consumers who make intermittent infrequent purchases and whose expectations are formed (adaptively) from recent experience mistake *unanticipated* absolute price rises for relative price increases during periods of *accelerating* inflation. With all goods being perceived as *relatively* more expensive, consumption of all goods falls. Many researchers have appealed to Deaton's (1977) hypothesis to justify the inclusion of inflation in consumption functions.

Inflation has also been used to proxy wealth effects. Hendry and Ungern-Sternberg (1981) - HUS hereafter - use inflation to adjust the conventional measure of *disposable* income for *expected* capital gains and losses on assets. This is because conventionally measured income incorporates increased nominal interest payments arising during inflationary periods but does not exclude the corresponding inflation induced capital gains and losses on assets. This violates the widely accepted Hicksian measure of income which is "the amount of accruals that an individual can spend (in real terms) leaving the real value of its wealth constant." (Rossi and Schiantarelli 1982, p. 374). Although this has often been implemented as a direct adjustment of income one could alternatively incorporate inflation as a separate regressor (see HUS p. 248). Indeed, inflation induced capital gains and losses need not be confined to income adjustments but may be incorporated as separate regressors on particular assets. For example, Pesaran and Evans (1984) enter inflation induced capital gains and losses on monetary holdings, equities and bonds as three separate variables in their saving function. Further, adjustments for capital gains and losses need not be confined to liquid, financial assets, as was common with the studies in the early/mid 1980s. Patterson (1984) argues that HUS's theoretical framework suggests one also needs to consider capital gains on illiquid wealth. The use of liquid assets in these early studies was probably due to data constraints rather than a belief that capital gains on liquid assets were of sole importance. Indeed, the financial deregulation that has been noticeable in the UK and the Nordic countries during the 1980s suggest that illiquid wealth effects are of increasing importance in consumption functions.

Bean (1978) presents evidence to support the argument that the statistical significance of inflation in consumption functions is best interpreted as capturing the impact of wealth, because the presence of high inflation causes a significant deterioration in the *real* value of money-fixed assets and so reduces consumption. This is argued to be of far greater importance than income uncertainty or money illusion effects because of the potential severity of the erosion of *real* wealth and its effect in causing consumers' to adjust their behaviour. Hadjimatheou (1987) points out that inflation is generally found to be negatively related to consumption, which is consistent with its use as a proxy for wealth effects. Recently, good series on assets have become available for a small number of countries which has allowed the testing of inflation effects separate from wealth effects. Lattimore (1994) finds no role for inflation in an annual Australian consumption function when well defined wealth effects are incorporated. Church *et al* (1994) report consumption functions from the major UK macroeconomic models. Wealth variables feature in all of these models while inflation has a role in none (except for the London Business School model where income is adjusted for inflation induced capital gains and losses). Thus, the emerging evidence appears to suggest that inflation has little direct influence on consumption, rather it acts as a proxy for wealth when data on assets is unavailable.

Given data constraints and the importance of asset effects, we employ inflation as a proxy for various wealth effects in our long run consumption functions.[4] To the extent that inflation influences consumption beyond acting as a wealth proxy, such effects will also be captured. We therefore base our empirical analysis upon the dynamic long run solution to DHSY's model, relaxing the unit income elasticity and ignoring the income growth term, thus:[5]

$$\ln C_t = b_0 + b_1 \ln Y_t + b_2 \Delta \ln P_t. \qquad \textbf{(4.1)}$$

[4] Pesaran, Shin and Smith (1997) utilise a similar argument to justify the use of inflation in their consumer function.

[5] Strictly, the DHSY model's dynamic long run solution (with the unit income elasticity postulate relaxed) is: $\ln C_t = b_0 + b_1 \ln Y_t + b_2 \Delta \ln P_t + b_3 \Delta \ln Y_t$. We exclude the income growth term, $\Delta \ln Y_t$, because it is not typically regarded as a primary argument in long run time-series consumption functions, because it would increase the dimension of the VAR and because its inclusion with the short run counterparts of $\ln Y_t$ can cause perfect multicollinearity when using the Johansen method.

Various hypotheses of interest can be tested using **(4.1)**. First, one can test whether there is statistical support for the existence of an equilibrium relationship (cointegration). Second, and given support for cointegration, one can obtain estimates for the long run elasticities to determine whether they are consistent with theoretical prior beliefs and, therefore, if the estimated equilibrium is economically sensible. For example, one might expect $b_1 \leq 1$ and $b_2 < 0$. Third, one can test which variables are statistically significant and whether consumption is homogeneous of degree one in income (the long run unit income elasticity, $b_1 = 1$), being the microeconomic homogeneity postulate *imposed* by DHSY.

We are aware of only three recent analyses of this model for a number of OECD countries. Firstly, Carruth *et al* (1996) estimate the dynamic DHSY model, which *implies* the equilibrium **(4.1)**, for a panel of the fifteen European Union (EU) countries over the period 1955-1990 using the Seemingly Unrelated Regression Estimation (SURE) procedure. They find implicit evidence favouring cointegration (by considering whether the adjustment coefficient on the error correction term is negative and statistically significant) for eight of the fifteen countries. The implicit rejection of cointegration for seven of the EU countries is found to manifest itself in the imposition of the long run unit income elasticity - a weak test suggests the rejection of the unit income elasticity for all EU countries except Ireland.[6] They also find inflation effects are statistically significant for only seven countries and in all these cases the influence is negative.[7] However, Stewart (1998) argues that to interpret inflation effects from the dynamic DHSY model as part of the equilibrium relationship, they would have to enter as lagged, rather than contemporaneous, terms: it is lagged nonstationary terms that are typically considered to define long run relations in error correction models. Since Carruth *et al* (1996) enter the inflation term contemporaneously this implies that this period's equilibrium consumption is determined by current income and *next* period's inflation. This is not the equilibrium of interest, especially if inflation is to proxy wealth effects, and we are cautious in interpreting these inflation effects as

[6] Carruth *et al* (1996) implicitly find evidence against valid error correction behaviour for Denmark, France, Ireland, Italy, the Netherlands, Sweden and the UK, by adding the lagged value of the log of income to the DHSY model.

[7] Carruth *et al* (1996) find statistically significant and negative inflation effects for Belgium, Finland, France, Greece, Ireland, Luxembourg and Sweden.

providing the desired long run information. Carruth (1996) *et al* also find that the consumption function is not similar across EU countries and suggest that "there may be a case for adopting a country specific search for the best empirical model of consumer spending" (Carruth *et al* 1996, p. 12), though recognise that this would require better data than they employ. For example, they use GDP at factor cost to approximate income which will likely produce different, and possibly misleading, results relative to disposable income.[8] Seven of the countries' dynamic models suffer from some form of misspecification indicating that these initial inferences need to be treated with caution.[9]

Pesaran, Shin and Smith (1997) investigate **(4.1)** for twenty-four OECD countries over the period 1962-1992/3 (thirty-one/thirty-two observations). They draw inference using twenty-four time-series regressions based upon the same dynamic autoregressive distributed lag model with the intercept restricted into the long run component of the model. With reference to the adjustment coefficient on the error correction term they find implicit evidence of cointegration for twenty countries. The four countries where there is no evident long run relationship are common to our sample and are Denmark, France, Germany and Switzerland. The estimated long run elasticity is clearly variable across countries being significantly less than unity in nine countries, significantly greater than unity in three and insignificantly different from one in the remaining twelve.[10] The long run inflation coefficient is more variable across countries than the long run income elasticity but is statistically significant in only ten countries, however, when it is significant it is negative.[11] Seven of the twenty-four countries' regressions are subject to

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[8] In Chapter 3 we note the different inference regarding stationarity when using the consumption-GDP ratio rather than the consumption-income ratio.

[9] Carruth *et al* (1996) find evident misspecification for Denmark, Italy, Luxembourg, Portugal, Spain, Sweden and the UK.

[10] Pesaran, Shin and Smith (1997) find evidence of a below unit income elasticity for Austria, Finland, Ireland, Japan, Luxembourg, Norway, Spain, Switzerland and Turkey, and an above income elasticity for Italy, the UK and the USA.

[11] Pesaran, Shin and Smith (1997) find the inflation coefficient to be statistically significant and negative for Australia, Belgium, Canada, Finland, Japan, the Netherlands, New Zealand, Sweden, the UK and the USA.

evident misspecification suggesting inference is invalid in these cases and possibly reflecting omitted explanatory factors.[12] Estimating the DHSY model for the whole panel of countries, imposing homogeneity of the long run elasticities across countries but allowing the short run dynamics to vary from country to country, they find that the hypothesis of common long run coefficients across countries is rejected. Since we are not aware of the availability of a method for estimating heterogeneous *long run* relations using panel data, this implies that, at present, the evident heterogeneity of long run consumption functions across countries requires the use of separate time-series regressions for each country. Thus, Pesaran, Shin and Smith's (1997) time-series results provide useful initial insights into the heterogeneity of OECD countries' long run consumption functions. However, the regression results discussed above use national disposable income, which incorporates general government disposable income, which could yield different inference from an income measure based solely upon the private sector.[13]

Larsson, Lyhagen and Lothgren (1998) illustrate the application of their panel cointegration test by applying it to **(4.1)** for twenty-three OECD countries using the *maximum* sample period 1960-1994.[14] The definitions of the variables are the same as those employed by Pesaran, Shin and Smith (1997). In particular, we emphasise their use of national, rather than the more appropriate private, disposable income measure. Inference from their time-series tests suggests one cointegrating vector for seventeen countries, two cointegrating vectors for four countries (Australia, France, Japan and Portugal) and three cointegrating vectors for two economies

[12] There is evidence of misspecification for Austria, Germany, the Netherlands, Norway, Portugal, the UK and the USA.

[13] Pesaran, Shin and Smith (1997) do run time-series regressions, therefore allowing for heterogeneous long run estimates, using private disposable *labour* income, which is more appropriate than disposable income. However, the sample only exceeds twenty (twenty-five) observations for eight (six) of the OECD countries suggesting, at present, that valid inference can only be obtained for a small number of countries.

[14] Larsson, Lyhagen and Lothgren's (1998) panel cointegration test is based upon the average of, in their application, the twenty-three individual countries' trace test statistics (from the standard Johansen Vector Error Correction Model; see equation **(4.2)** below), adjusted using a standard normal transformation. The use of each lagged *level* variable reduces their time-series sample size by one observation. The length of time-series used for each country is always shorter than that employed here.

(Austria and Greece). Their panel cointegration test indicates that the *largest* number of common cointegrating vectors across the panel is two. They then consider applying overidentification restrictions on the potential two cointegrating vectors inferred. The cointegration space is restricted so that consumption and income consitute one vector and inflation a separate vector - inflation is stationary, or cointegrates with itself. They report time series tests for each individual economy suggesting that these restrictions cannot be rejected for all but two countries (Portugal and Turkey). A panel version of this overidentification test confirms that consumption and income cointegrate and inflation is stationary across the twenty-three countries.

These recent investigations indicate that there exist one or two cointegrating vectors between consumption, income and inflation for OECD countries. It also appears that there is no common consumption function for these countries suggesting the need to develop country-specific models. Since we are not aware of any panel estimation methods that allow both the specification of short run dynamics and estimates of long run elasticities to be different from country to country this means that, at present, the most flexible country-specific models will be secured through time-series estimation. Another common feature of these three studies is that their inferences relating to the majority of OECD countries with reasonable length samples are drawn using income measures which incorporate government income. Superior inferences regarding private sector behaviour may be obtained using income measures solely based upon the private sector. Both the Carruth *et al* (1996) and Pesaran, Shin and Smith's (1997) investigations include a proportion of models subject to some form of misspecification. Future attempts to build models free from evident misspecification would be desirable. One issue upon which the above studies are not in clear agreement regards the validity of the long run unit income elasticity. The study by Carruth *et al* (1996) suggests that the relaxation of the unit elasticity substantially improves virtually all EU countries' models whereas Pesaran, Shin and Smith (1997) cannot reject the unit income elasticity for twelve of the twenty-four OECD countries. These studies also cast doubt over the relevance of inflation as a determinant of long run consumption. Further clarification of these issues is desirable.

The aim of the present Chapter is to build on the results of these recent studies by *explicitly* testing for cointegration. Where cointegration is evident we estimate country specific long run consumption functions using time-series data to reflect the evident heterogeneity of consumer behaviour across countries in as flexible a manner as is currently possible. We also seek to obtain superior parameter estimates by using income measures solely based upon the private sector. We have data for the private sector which will allow us to estimate **(4.1)** for twenty OECD countries using 1960-1994 as our estimation period (thirty-five observations). To our knowledge, there is no previous study which estimates consumption functions for so many countries, using such a long time-series of data based solely upon the private sector. This study also seeks to ensure valid inference by selecting models to be free from evident misspecification. These estimated models will enable us to further clarify whether the long run unit income elasticity postulate is valid or not and whether inflation constitutes part of the long run consumption function.

4.3 Testing for Cointegration Using the Johansen Procedure

Engle and Granger (1987) introduced a means for estimating and testing for cointegrating relationships. They specify the necessary condition for cointegration as all variables being integrated of the same order as the dependent variable (y_t).[15] The order of integration is the number of times (d) a variable (x_t) must be differenced to induce second-order stationarity: denoted $x_t \sim I(d)$. Typically (the logs of) variables are I(1). The sufficient condition for cointegration is that the linear combination of variables ($u_t = y_t - \beta_{EG} x_t$ where β_{EG} is the cointegrating vector) exhibits a reduced order of integration ($u_t \sim I(d-b)$, b>0, typically d=b=1, implying $u_t \sim I(0)$). If there is evidence of cointegration there exists, according to the Granger representation theorem, an error correction representation of the variables. That is, one can obtain a parsimonious relationship between the variables, following the general-to-specific methodology, where $-\gamma$ is negative and statistically significant in a dynamic error correction model such as: $\Delta y_t = \delta_0 + \Sigma\delta_{1i}\Delta y_{t-i} + \Sigma\delta_{2i}\Delta x_{t-i} - \gamma u_{t-1} + v_t$, where v_t is a white noise error term.

[15] This necessary condition may turn out to be too stringent. What one requires is that the linear combination of regressors explain the nonstationarity of the dependent variable. Thus, provided at least one regressor is integrated of the same order as the dependent variable, variables integrated of a lower order may also enter the cointegrating relation - see Hall and Patterson (1992). Indeed, regressors with higher orders of integration than the dependent variable may enter the equilibrium if their linear combination is integrated of an order no greater than the dependent variable - see Charemza and Deadman (1997).

The Engle and Granger method has been criticised for reasons which include the following. Ordinary Least Squares (OLS) estimation of the cointegrating vector does not allow for the possible endogeneity of regressors. Inference regarding the existence of an equilibrium is sensitive to the variable upon which the model is normalised. The method is potentially invalid for equilibriums comprised of more than two variables because it will be unable to identify any multiple cointegrating relations that may exist. The method yields inefficient and, as Malley and Moutos (1996) suggest, inconsistent estimates of the long run parameters in finite samples. It fails to account for short run dynamics when estimating the long run relation so is subject to omitted variable bias. Further, Gerrard and Godfrey (1998) argue that the typical diagnostic checks for heteroscedasticity and functional form are invalid for the first stage regression making it impossible to test the cointegrating relation for these forms of misspecification. Johansen's method for testing for cointegration overcomes these problems and Horioka (1996) cites Shintani's (1994) finding that this procedure has greater power than the Engle and Granger method. We therefore use the standard Johansen procedure to test for cointegration. For a set of K, I(1) endogenous variables, $\mathbf{X}_t$, the Johansen procedure is based upon the Vector Error Correction Model (VECM):

$$\Delta\mathbf{X}_t = \boldsymbol{\Gamma}_1\Delta\mathbf{X}_{t-1} + \ldots + \boldsymbol{\Gamma}_{L-1}\Delta\mathbf{X}_{t-L+1} + \boldsymbol{\Pi}\mathbf{X}_{t-1} + \mathbf{B}\mathbf{D}_t + \mathbf{A} + \mathbf{u}_t, \qquad \textbf{(4.2)}$$

where $\boldsymbol{\Gamma}_i$, i=1,...,L-1, are the coefficients on the lagged stationary terms; $\boldsymbol{\Pi}$ ($=\boldsymbol{\alpha\beta}'$) is the coefficient matrix on the nonstationary terms, with $\boldsymbol{\alpha}$ being a matrix of adjustment coefficients and $\boldsymbol{\beta}$ is a matrix of r distinct cointegrating vectors.[16] $\mathbf{D}_t$ denotes a matrix of J-1 contemporaneous exogenous variables which, in this case, are dummy variables, where **B** represents the coefficient matrix corresponding to these dummies; **A** is the intercept matrix and $\mathbf{u}_t$ is a vector of error terms. The number of cointegrating vectors is determined by the rank of $\boldsymbol{\Pi}$ using the maximum eigenvalue and/or trace test statistics.

Dummy variables may be included to remove evident misspecification (primarily departures

[16] $\boldsymbol{\Pi}$ is the long run solution to the *levels* VAR system $\mathbf{X}_t = \boldsymbol{\Pi}_1\mathbf{X}_{t-1} + \boldsymbol{\Pi}_2\mathbf{X}_{t-2} + \ldots + \boldsymbol{\Pi}_L\mathbf{X}_{t-L}$ + deterministic terms + $\mathbf{u}_t$, corresponding to the error correction form, **(4.2)**. Deterministic terms enter the cointegrating vector only when they are restricted into the cointegration space.

from normally distributed residuals: outlying errors) which may arise due to many factors, including omitted variables.[17] There are many country specific events that may cause misspecification due to omitted variables such as German reunification in 1990/1991, the dramatic slowdown in Japan's remarkable post-war growth in 1973/1974 and the financial deregulation that occurred in the UK (see, for example, Miles 1992) and the Nordic countries (see, for examples, Lehmussaari 1990 and Berg 1994) during the 1980s. There may also be other country-specific events of which we are currently unaware. Further, we do not use a direct measure of wealth so inflation may be unable to fully approximate all asset effects for all countries. Therefore, any large non-synchronised movements of consumption and income will cause a large change in savings and, therefore, wealth. With wealth omitted from our explanatory factors, any such large changes may manifest themselves as outlying errors. Further, because the VECM is a system of equations, outliers in non-consumption equations could also cause misspecification. We also note that use of dummy variables to remove misspecification may be more desirable than continuously extending the lag length of the VECM because, as Hall (1991) points out, choosing too large a lag length when degrees of freedom are likely to be scarce will cause bias in the tests for cointegration. Indeed, a parsimonious means of removing residual autocorrelation *and* departures from normality is desirable because the Johansen procedure has been shown to be very sensitive to the independent normal errors assumption (see, Huang and Yang, 1996).[18] The use of dummy variables as an alternative to increasing the size of the VAR system is advocated by Clements and Mizon (1991).

For comparative purposes one standard specification will be used to test for cointegration for all countries. Given the annual frequency of the data one may have the prior belief, based upon previous researchers' findings, that the VAR lag length, L, is equal to two (one lagged stationary term). Pre-empting the results below, the system version of Schwartz's Bayesian Information

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[17] Pesaran, Shin and Smith (1997) argue that there are likely to be omitted factors from consumption functions including only consumption, income and inflation.

[18] Maddala and Kim (1998) cite Huang and Yang's (1996) finding that when "the errors are not independent normal... the Johansen method has a greater probability (than least squares methods) of rejecting the null of no cointegration even when there are no cointegrating relations." (Maddala and Kim 1998, p. 173).

Criterion (denoted SBIC, hereafter) tends to indicate a VAR lag length of one or two.[19] A lag length of one may be considered overly parsimonious because the model excludes all short run dynamics and, as Hall (1991) observes, choosing too small a lag length can make the test statistics for cointegration unreliable by biasing the residuals of the VAR. Considering a standard model with L=2 helps us defend ourselves against such potential problems. The standard unrestricted specification, in terms of the variables we use, is:[20]

$$\Delta \ln C_t = a_{10} + \gamma_{11}\Delta \ln C_{t-1} + \gamma_{12}\Delta \ln Y_{t-1} + \gamma_{13}\Delta\Delta \ln P_{t-1} + \pi_{11}\ln C_{t-1} + \pi_{12}\ln Y_{t-1} + \pi_{13}\Delta \ln P_{t-1} + u_{1t} \quad \textbf{(4.3)}$$

$$\Delta \ln Y_t = a_{20} + \gamma_{21}\Delta \ln C_{t-1} + \gamma_{22}\Delta \ln Y_{t-1} + \gamma_{23}\Delta\Delta \ln P_{t-1} + \pi_{21}\ln C_{t-1} + \pi_{22}\ln Y_{t-1} + \pi_{23}\Delta \ln P_{t-1} + u_{2t}$$

$$\Delta\Delta \ln P_t = a_{30} + \gamma_{31}\Delta \ln C_{t-1} + \gamma_{32}\Delta \ln Y_{t-1} + \gamma_{33}\Delta\Delta \ln P_{t-1} + \pi_{31}\ln C_{t-1} + \pi_{32}\ln Y_{t-1} + \pi_{33}\Delta \ln P_{t-1} + u_{3t}$$

However, because a different specification may be preferred for any particular country model selection criteria need to be applied. Whether the long run matrix includes an intercept or not does not affect the VAR in unrestricted reduced form (VECM) so model selection criteria applied to the VECM will yield the same results however the intercept is specified. Therefore model selection criteria are applied to **(4.2)** to gauge if the favoured model deviates from the standard form **(4.3)**.

4.3.1 VECM Model Selection

The favoured VECM for any particular country is determined by estimating **(4.2)** for L=1,2,3 and 4 and choosing the favoured VECM as that with the lowest SBIC from those which are free

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[19] In Larsson, Lyhagen and Lothgren's (1998) application the lag length, L, chosen for the Johansen equation, **(4.2)**, is one for all countries, except Iceland, the Netherlands and Switzerland, where it is two. This is chosen using the SBIC and yields "reasonable fit in terms of the test statistics for normality and autocorrelation." (Larsson, Lyhagen and Lothgren, 1998, p. 11).

[20] Our general inference upon the orders of integration of the logs of consumption and income and inflation, based upon ADF tests and visual inspection of the data, were that they are all I(1) - see Chapter 3.

from evident autocorrelation and non-normality (according to both system and unreported individual equation tests). This yields up to two models for each country, the standard model, **(4.3)**, and, if different, the favoured model.

Table 4.1 summarises the model selection results for each country. The first column specifies the country to which the results relate. The second column specifies whether and which dummy variables have been employed. The system SBIC, system test for second-order serial correlation (denoted "SC") and system test for departures from normality (denoted "N") are reported for the VECM with lag lengths (L=) 1, 2, 3 and 4.[21]

A lag length greater than two is only favoured for one country, Australia (where L=3).[22] For

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[21] The vector version of the SBIC (see Doornik and Hendry 1995, p. 286) is defined as:

$$SBIC = \ln|\hat{\Omega}| + p[(\ln T)/T]$$

where $|\hat{\Omega}|$ is the determinant of $\hat{\Omega} = [(\hat{\mathbf{u}}_t\hat{\mathbf{u}}_t')/T]$. The residual variance/covariance matrix in **(4.2)**, $\mathbf{u}_t\mathbf{u}_t'$, is 3x3. p is the number of parameters in the system, which is K(KL+J) for **(4.2)**. T is the sample size. The smallest SBIC indicates the model with the optimal fit versus parsimony trade-off. The vector error serial correlation test considers the null hypothesis H_0: $\mathbf{R}_1 = \mathbf{R}_2 = \ldots = \mathbf{R}_S = 0$, in the auxiliary system:

$$\Delta\mathbf{X}_t = \sum_{i=1}^{L-1}\mathbf{\Gamma}^*_i\Delta\mathbf{X}_{t-i} + \mathbf{\Pi}^*\mathbf{X}_{t-1} + \mathbf{B}^*\mathbf{D}_t + \mathbf{A}^* + \sum_{s=1}^{S}\mathbf{R}_s\hat{\mathbf{u}}_{t-s} + \mathbf{e}_t$$

where $\mathbf{R}_s$ is a KxK matrix of coefficients on autoregressive error terms (from equation **(4.2)**) for s lags and $\mathbf{e}_t$ is a Kx1 vector of white noise disturbance terms. Significant S-order autocorrelation is rejected if the approximate F-statistic is below its critical value (with K^2S restrictions) - see p. 216 in Doornik and Hendry (1995) for further details. We use S=2, being the statistic automatically produced by PcFiml 8.0 in its test summary (for annual data). The vector normality test employed is a direct analogue of the standard single equation skewness and kurtosis test statistic - see Doornik and Hendry (1995) pp. 216-217 for details. If the test statistic is below its critical value (defined by a χ^2 distribution with 2K degrees of freedom) one can reject evident departures from normality.

[22] We also report a model with L=3 for Denmark, Sweden and the UK. As we shall see below, this is the only specification for Sweden which did not reject the hypothesis of cointegration. For Denmark and the UK this is the only specification which yields a plausible, unique cointegrating vector.

TABLE 4.1: VECM Model Selection

LAGS (L) →		1			2			3			4		
Cntry	Dummy	SBIC	SC	N	SBIC	SC	N	SBIC	SC	N	SBIC	SC	N
AUL	NONE	-24.460	**2.901**	7.823	-24.180	**2.615**	8.333	**-23.820**	1.616	7.002	-23.570	0.813	6.504
AUT	NONE	-24.910	0.790	**18.120**	-24.260	0.582	8.314	-23.490	1.515	4.588	-23.130	1.455	5.493
	74;78	**-25.090**	0.684	4.475									
BEL	NONE	-23.540	**3.407**	10.539	**-23.690**	1.701	9.895	-23.450	1.054	7.967	-23.260	1.018	12.219
CAN	NONE	-23.740	**2.615**	12.409	-24.050	0.686	**16.216**	-23.350	1.060	**21.889**	-22.890	1.126	**20.377**
	76;82;91				**-24.960**	1.203	1.923						
DEN	NONE	-21.690	**4.409**	3.493	**-22.430**	0.757	4.255	-21.960	0.835	2.867	-21.240	**2.165**	5.030
FIN	NONE	-21.950	**1.822**	8.665	-21.900	1.130	7.639	-21.340	0.883	10.474	-20.900	0.774	9.616
	69;72;74				**-22.260**	0.941	2.468						
FRA	NONE	-25.720	1.082	9.681	-25.120	0.731	6.224	-24.470	1.357	6.700	-23.820	**2.868**	7.194
	74	**-26.130**	1.221	2.179									
GER	NONE	-24.950	1.171	9.848	-24.450	0.895	**14.828**	-23.940	1.062	**15.650**	-23.470	1.511	9.989
	91	-25.260	**1.792**	7.577	**-24.970**	1.370	7.515						
GRE	NONE	**-22.710**	1.430	2.210	-22.110	1.342	1.802	-21.620	1.456	1.254	-21.030	0.762	4.396
ICE	NONE	-16.810	**2.212**	**15.210**	**-16.850**	0.529	9.851	-16.400	0.576	8.192	-15.720	0.549	6.747
IRE	NONE	-21.300	0.929	**14.598**	-20.830	1.468	**18.132**	-20.180	1.473	**20.532**	-20.040	0.676	**16.105**
	73;82	**-21.600**	1.280	5.471	-21.350	0.607	8.087						
ITA	NONE	-22.330	**2.220**	5.727	-22.980	**1.863**	4.679	-22.710	1.284	11.222	-22.060	1.776	10.250
	93	-23.690	**1.951**	4.955	**-23.400**	1.297	3.058						
JAP	NONE	-24.030	**2.667**	**25.843**	-24.390	1.168	6.130	-23.740	1.287	7.640	-23.600	1.057	7.107
	74				**-24.770**	0.962	6.559						
NET	NONE	**-24.380**	1.742	2.137	-24.220	0.838	2.425	-23.670	1.155	4.815	-23.280	1.026	4.894
NOR	NONE	-22.570	1.608	**16.070**	-22.190	1.478	**20.253**	-21.550	1.452	11.840	-21.180	1.321	2.303
	708081; 78;8586	**-23.690**	0.888	8.680	-23.220	0.837	12.259						
SPA	NONE	-23.020	1.574	10.930	-22.660	0.924	6.499	-22.260	1.016	6.794	-21.760	0.602	6.730
	74;77	**-23.160**	1.279	5.563	-22.810	1.246	3.197						
SWE	NONE	-23.010	**1.886**	8.775	-22.650	1.261	**13.998**	-22.420	1.107	8.321	-21.820	1.707	8.993
	92				**-23.650**	1.532	5.886						
SWZ	NONE	-24.900	**1.789**	2.544	-25.100	1.234	6.070	-24.500	1.407	8.498	-24.210	**1.899**	5.701
	6386;71; 79				**1.560**	1.556	3.274						
UK	NONE	-22.820	**2.194**	**16.560**	-22.730	0.952	12.465	-22.470	0.518	9.863	-21.930	0.511	11.784
	74;75				**-23.740**	1.008	5.929						
USA	NONE	-25.720	**1.843**	7.607	**-25.510**	0.964	5.097	-24.900	1.198	8.074	-24.450	0.640	3.690
Distribution			$F18,65$	$\chi^2(6)$		$F18,57$	$\chi^2(6)$		$F18,48$	$\chi^2(6)$		$F18,40$	$\chi^2(6)$
5% Critical Values			1.779	12.59		1.799	12.59		1.838	12.59		1.872	12.60

Table 4.1 Notes: SBIC is the system version of Schwartz's fit versus parsimony criteria, SC is a system test of second order serial correlation while N is a system version for testing for departures from normally distributed residuals. Distributions and critical values are given at the bottom of the table - when the model includes dummies the SC tests

use different degrees of freedom on the denominator and are accounted for in drawing inferences. Bold emphasis indicates misspecification at the 5% level of significance. *Spike* dummy variables are indicated by the year which takes on the unit value, for example, 74;78 indicates two dummies, the first being unity in 1974 and zero otherwise and the second being unity in 1978 and zero otherwise. Similarly, *single* dummy variables with more than one non-zero value are indicated by, for example, 8586, where in 1985 and 1986 the variable is unity but otherwise zero. Bold emphasis of the SBIC indicates the favoured model for a particular country: minimum SBIC for models with no evidence of misspecification.

seven of the twenty countries the favoured lag length is one (Austria, France, Greece, Ireland, the Netherlands, Norway and Spain). The remaining twelve countries are consistent with our prior belief because they favour a lag length of two.[23] However, for only four of the twenty countries (Belgium, Denmark, Iceland and the USA) is the favoured model the same as the standard model. In those countries where the favoured and standard models are different, both will be subject to the hypothesis tests outlined below.

4.3.2 Testing for cointegration

The existence, and number, of cointegrating vectors will be tested assuming both that an intercept enters (restricted) and does not enter (unrestricted) the cointegration space. As is well known and outlined in, for examples, Johansen (1988), Johansen and Juselius (1990), Johansen (1991) and Johansen (1995), the existence and number of cointegrating vectors depends upon the rank of the matrix, $\mathbf{\Pi}$, containing the coefficients on the lagged nonstationary terms which, for example, in **(4.3)**, are: $\pi_{11}, \pi_{12} \ldots, \pi_{33}$. If $\mathbf{\Pi}$ is of full rank, in our case r=3, this implies that the terms entered as nonstationary in the VECM are stationary, which contradicts the assumption that all variables in $\mathbf{X}_t$ are I(1). If $\mathbf{\Pi}$ is of zero rank, r=0, this implies there is no cointegration and the nonstationary terms need to be removed from the VECM to ensure both sides of the VAR are stationary. Finally, if $\mathbf{\Pi}$ is of reduced rank, r=1 or r=2 in our case, this implies $\mathbf{\Pi X}_{t-1}$

[23] One frequently employed procedure for selecting the favoured lag length in a Johansen VECM is to choose the number of lagged variables as that which minimises the SBIC. Had we applied this criteria to those models which exclude dummy variables we would have selected models subject to evident autocorrelation and/or non-normality for eight countries (Australia, Austria, Canada, Finland, Ireland, Italy, Norway and Sweden). This suggests that ignoring misspecification tests in favour of using selection criteria, such as the SBIC, can lead to choosing models which provide invalid inference.

~ I(0) and hence the existence of r cointegrating vectors. The value of r is determined using a likelihood ratio test, either based upon the maximum eigenvalue of the stochastic matrix or its trace.[24] When either test statistic exceeds its 5% critical value, the null that r equals zero, one or two, is rejected in favour of a larger number of cointegrating vectors. Ford and Morris (1995) suggest that Johansen (1988) prefers the maximum eigenvalue version of the test while Cheung and Lai (1993) indicate that the trace statistic is more robust in the face of skewness and excess kurtosis. Most of the literature appears to prefer the trace version. Both test statistics are reported here.

Our prior belief is that there will be one cointegrating vector defining the equilibrium (log) level of consumption. However, drawing inference from the Johansen test may not be straightforward, especially when using small samples, due to its low power, the possibility of spurious cointegration, its sensitivity to how restricted the VECM specification is and the chosen lag length of the VAR.[25] We will consider the impact of each of these factors upon inference in turn.

First, there is a general trade off between the size and power of a test: reducing the probability of a type I error (size) will also reduce the probability of correctly rejecting a false null hypothesis (power). Setting the size at the usual five percent level means that the power of the test will also be low which, given the intrinsic low power of the Johansen test (especially in small samples), means that the probability of correctly rejecting cointegration will be small.

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[24] Both test statistics are likelihood ratio tests; the first is based upon the maximum eigenvalue (EIG) of the stochastic matrix while the latter is based upon this matrix's trace (TRA), that is:

$$LR_{TRA} = T \sum_{i=r+1}^{K} \ln (1-\lambda_i); \qquad \text{and,} \qquad LR_{EIG} = T \ln (1-\lambda_{r+1}).$$

Where $\lambda_{r+1}, \ldots, \lambda_n$ are the K-r smallest squared canonical correlations (see Johansen 1988 p. 233), and K is the number of endogenous variables in the system (K=3 in the present analysis). The test statistics defined above are compared to the 95% and 99% critical values, reported in Osterwald-Lenum (1990) and reproduced in Banerjee *et al* (1993), to determine the value of r.

[25] Degrees of freedom are scarce in the present study because we only use thirty-five observations in our estimation period. However, since Horioka (1996) obtained useful inference when employing the Johansen procedure to a three variable VAR using a *maximum* of 38 observations, we believe our inference will also be useful.

Thus, the Johansen test may not indicate as much cointegration as actually exists.

Second, Gonzalo and Lee (1998) suggest that the Johansen procedure has a tendency to infer spurious cointegration when variables included in the model are not pure I(1) processes, but cannot be distinguished from being I(1) using standard tests.[26] They suggest the need for a deeper data analysis than is provided by standard unit root tests to properly implement the Johansen method. Spurious cointegration also arises if singularity in the VECM is caused by the error covariance matrix and not just the long run impact matrix, $\mathbf{\Pi}$. This suggests that the Johansen procedure may indicate too much cointegration.

Third, it is known that the more restricted is the VECM's specification the more favourable is the test towards finding cointegration. For example, specifying the intercept as restricted into the cointegration space will more likely uncover long run relations compared to those with an unrestricted intercept. We report cointegration results for both the restricted and unrestricted intercept specifications.

Fourth, Hall (1991) points out that if the lag length of the VAR is too large and the sample size is small, the canonical correlations will approach unity as degrees of freedom fall, biasing the test statistics upwards and making inference more favourable towards cointegration. We conduct cointegration tests for four countries with a lag length greater than two (Australia, Denmark, Sweden and the UK) so these countries' models will be the most likely to be subject to such bias. On the other hand, Hall (1991) notes that if the lag length is so small that the residuals of the VAR are serially correlated, the cointegration test statistics will become unreliable. In the present study, only two of the standard models (L=2) show signs of residual autocorrelation (Australia and Italy) and there is no evidence of misspecification in any of the favoured models. Hence, the test statistics are unlikely to be affected by serial correlation but may, through small degrees of freedom, show some bias towards cointegration.

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[26] It is suggested that variables with long-memory properties and a trending behaviour which are not pure I(1) processes may be difficult to differentiate from being I(1).

Overall, there are potential biases which can cause one to infer too much or too little cointegration. However, it is difficult to evaluate which type of bias will be present for any particular VECM and, when there is more than one, which will dominate. To the extent that such biases may exist in our analysis, we show pragmatism when interpreting our statistical results.

Given the possible sensitivity of inference to specification and that, theoretically speaking, we have a strong prior belief that only one cointegrating vector exists, our aim is see whether we can uncover statistical support for a unique cointegrating relation. That is, to see which VECM specifications cannot reject the finding of a single cointegrating vector. If for any particular country cointegration cannot be secured using a 5% significance level, we will consider whether results more favourable to our prior can be obtained using critical values at the 10% level of significance.[27] Similarly, if more than one cointegrating vector is all that can be justified using the 5% level for any particular country, we will consider use of 1% critical values. Thus, we look for statistical support for our economic prior of a unique cointegrating vector. Pesaran and Pesaran (1997), p. 297, argue that one may appeal to economic priors when, *in any particular instance*, it is feared that such statistical procedures are uniformative regarding cointegration rank, especially when employing small samples - as we do here. Indeed, Greenslade *et al* (1998) provide a Monte Carlo experiment which demonstrates how, when using small samples, an unrestricted VECM with eight endogenous variables, based upon asymptotic results, can easily underestimate or overestimate the true number of cointegrating vectors. They suggest that "a thorough use of economic theory at an early stage, rather than treating a model as a pure statistical artefact, can yield enormous benefits." (Greenslade *et al* 1998, p. 1). The emphasis of their work is on the imposition of exogeneity restrictions. In the present study we do not have a strong prior belief that income and/or inflation can be treated as exogenous. However, in an analogous manner to Greenslade *et al* (1998) we seek to use our economic prior of a unique cointegrating vector to guide our choice of r in the face of potentially misleading inference when using small samples.

[27] Bewley and Yang (1998) note that applied researchers commonly resort to using the 10% level of significance when employing the Johansen procedure.

TABLE 4.2: Testing For Cointegration in the VECM

		Null Hypotheses -		r=0		r=1		r=2	Inference (r=)	
Country	Dummies	Lags (L)	Intercept	Max Eig	Trace	Max Eig	Trace	Eig/Trace	5%	1%
AUL	NONE **M**	2	U	13.550	17.290	3.495	3.736	0.241	0	0
	NONE **M**	2	R	32.650	46.210	10.670	13.560	2.891	1	1
	NONE	3	U	12.320	18.330	6.010	6.010	0.001	0	0
	NONE	3	R	25.950	37.950	10.230	12.000	1.769	1	1
AUT	NONE	2	U	16.350	31.340	9.535	14.990	5.450	1	0
	NONE	2	R	25.500	44.690	13.730	19.180	5.450	1	1
	74;78	1	U	34.090	52.990	12.560	18.910	6.349	1	1
	74;78	1	R	79.900	105.600	15.690	25.700	10.010	3	1
BEL	NONE	2	U	9.681	17.400	5.399	7.716	2.317	0	0
	NONE	2	R	21.030	32.290	7.469	11.260	3.794	1*	0
CAN	NONE **M**	2	U	19.680	33.170	8.937	13.490	4.554	1	0
	NONE **M**	2	R	28.160	47.470	14.730	19.310	4.577	1	1
	76;82;91	2	U	32.520	40.570	6.363	8.052	1.690	1	1
	76;82;91	2	R	42.370	70.480	26.390	28.110	1.720	2	2
DEN	NONE	2	U	29.300	42.370	13.030	13.070	0.040	1	1
	NONE	2	R	41.780	59.550	13.270	17.770	4.497	1	1
	NONE	3	U	24.760	39.150	14.240	14.390	0.154	1	1
	NONE	3	R	28.070	46.940	15.160	18.860	3.701	1	1
FIN	NONE	2	U	16.520	29.150	8.242	12.630	4.391	0	0
	NONE	2	R	16.620	34.240	9.389	17.620	8.234	0	0
	69;72;74	2	U	17.380	34.360	13.090	16.980	3.884	3	0
	69;72;74	2	R	17.530	39.150	16.580	21.620	5.034	2	0
FRA	NONE	2	U	23.020	35.270	10.020	12.240	2.227	1	0
	NONE	2	R	27.540	43.400	10.310	15.870	5.555	1	1
	74	1	U	41.980	60.970	16.550	18.990	2.439	2	1
	74	1	R	89.500	114.300	18.840	24.760	5.919	2	1
GER	NONE **M**	2	U	17.580	29.290	8.157	11.710	3.556	0	0
	NONE **M**	2	R	17.760	37.350	15.000	19.590	4.590	1	0
	91	2	U	17.530	32.740	13.700	15.210	1.509	1	0
	91	2	R	29.120	51.430	17.500	22.310	4.809	2	1
GRE	NONE	2	U	19.890	31.560	8.643	11.670	3.030	1	0
	NONE	2	R	22.750	38.070	9.307	15.320	6.017	1	0
	NONE	1	U	52.900	72.630	14.960	19.730	4.771	3	1
	NONE	1	R	82.590	110.600	18.400	27.970	9.577	3	1
ICE	NONE	2	U	19.100	32.410	10.640	13.320	2.672	1	0
	NONE	2	R	22.010	38.570	13.180	16.560	3.384	1	0

Table 4.2 continued

		Null Hypotheses –		r=0		r=1		r=2	Inference (r=)	
Country	Dummies	Lags (L)	Intercept	Max Eig	Trace	Max Eig	Trace	Eig/Trace	5%	1%
IRE	NONE M	2	U	14.490	21.880	7.364	7.385	0.021	0	0
	NONE M	2	R	18.630	34.040	11.950	15.410	3.467	0	0
	73;82	1	U	34.360	41.690	7.330	7.331	0.001	1	1
	73;82	1	R	41.780	71.310	28.140	29.530	1.391	2	2
ITA	NONE M	2	U	12.100	23.000	8.787	10.900	2.111	0	0
	NONE M	2	R	16.050	32.820	10.850	16.770	5.918	0	0
	93	2	U	15.450	27.720	11.140	12.270	1.135	0	0
	93	2	R	27.660	45.810	12.030	18.150	6.116	1	1
JAP	NONE	2	U	25.040	46.490	18.600	21.450	2.847	2	2
	NONE	2	R	36.020	60.080	18.840	24.060	5.219	2	1
	74	2	U	32.540	47.010	12.470	14.480	2.008	1	1
	74	2	R	45.840	64.430	15.910	18.590	2.684	1	1
NET	NONE	2	U	12.890	27.220	10.500	14.320	3.826	0	0
	NONE	2	R	15.170	33.350	12.390	18.180	5.786	0	0
	NONE	1	U	22.630	38.730	12.620	16.090	3.470	1	1
	NONE	1	R	54.000	70.880	12.690	16.880	4.190	1	1
NOR	NONE M	2	U	9.945	16.320	5.715	6.379	0.664	0	0
	NONE M	2	R	13.230	25.180	6.317	11.950	5.630	0	0
	708081;78;8586	1	U	33.270	48.710	14.520	15.440	0.923	2	1
	708081;78;8586	1	R	55.830	85.910	17.470	30.080	12.610	3	1
SPA	NONE	2	U	39.250	46.930	7.498	7.679	0.181	1	1
	NONE	2	R	43.950	56.620	9.860	12.670	2.810	1	1
	74;77	1	U	45.400	75.750	28.050	30.360	2.307	2	2
	74;77	1	R	51.960	107.600	44.860	55.670	10.800	3	2
SWE	NONE M	2	U	13.900	20.200	4.013	6.120	2.108	0	0
	NONE M	2	R	15.420	23.100	4.249	7.679	3.430	0	0
	92	2	U	13.510	18.610	3.638	5.100	1.461	0	0
	92	2	R	15.190	21.850	4.240	6.680	2.482	0	0
	NONE	3	U	16.900	25.940	4.947	9.043	4.096	0	0
	NONE	3	R	23.830	34.680	6.533	10.850	4.317	1	0
SWZ	NONE	2	U	15.750	30.220	12.830	14.470	1.635	1	0
	NONE	2	R	24.890	43.670	15.500	18.790	3.286	1	1
	6386;71;79	2	U	22.020	39.010	12.070	16.990	4.915	1	1
	6386;71;79	2	R	30.160	59.270	22.020	29.100	7.089	2	1

Table 4.2 continued

UK	NONE	2	U	16.290	20.170	3.879	3.885	0.006	0	0
	NONE	2	R	17.050	31.710	10.790	14.670	3.873	0	0
	74;75	2	U	16.200	**31.100**	**14.820**	14.900	0.085	2	0
	74;75	2	R	**30.790**	**52.250**	15.600	**21.470**	5.866	1	1
	NONE	3	U	**24.090**	27.850	3.293	3.759	0.465	1	0
	NONE	3	R	**26.410**	**49.360**	**19.680**	**22.950**	3.269	2	1
USA	NONE	2	U	14.100	19.730	4.458	5.639	1.181	0	0
	NONE	2	R	**25.490**	**35.820**	7.307	10.330	3.021	1	0
5% Critical Values			U	*21.00*	*29.70*	*14.10*	*15.40*	*3.80*		
			R	*22.00*	*34.90*	*15.70*	*20.00*	*9.20*		
1% Critical Values			U	*25.52*	*35.65*	*18.63*	*20.04*	*6.65*		
			R	*26.81*	*41.07*	*20.20*	*24.60*	*12.97*		

Table 4.2 Notes: Dummy variables are as specified in Table 4.1. The status of the intercept is indicated as unrestricted (U) or restricted (R). Max Eig (Trace) is the maximum eigenvalue (trace) test statistic for cointegration for the null hypotheses that the number of cointegrating vectors (r) equal 0, 1 and 2. For the null of r=2 the trace and maximum eigenvalue statistics are the same. A bold test statistic indicates rejection of the null hypothesis at the 5% level of significance. The 5% and 1% critical values are given at the bottom of the table (no account has been made when dummy variables are included). The number of cointegrating vectors favoured at the 5% and 1% levels are given in the last columns, headed 5% and 1%, respectively. An asterix (*) denotes significance and number of cointegrating vectors indicated at the 10% level. The 10% critical values for r=1, r=2 and r=3 are, with unrestricted intercept, 18.60 (26.79), 12.07 (13.33) and 2.69; and with restricted intercept, 19.7, (32.0) 13.75 (17.85) and 7.53 - trace critical values are given in brackets when different from those based upon the maximum eigenvalue.

Table 4.2 reports the Johansen cointegration test results. The first column of the Table indicates the country to which the results relate while the dummy variables employed are specified in the second column - an ("M") in this second column indicates that the model suffers from evident misspecification according to the tests reported in Table 4.1. The third and fourth columns denote, respectively, the number of lags ("L") included in the VECM and whether the intercept in the cointegration space is unrestricted ("U") or restricted ("R"). The next five columns of the Table specify the maximum eigenvalue ("Max Eig") and trace ("Trace") test statistics for the null hypotheses that r=0, r=1 and r=2. A bold test statistic indicates rejection of the null hypothesis at the five percent level. The number of cointegrating vectors inferred by these tests at the 5% and 1% levels of significance are given in the last two columns (denoted "Inference r="). If r=1 can be inferred by *either* the trace or maximum eigenvalue statistics then we will infer the presence of one cointegrating vector, as suggested by our prior economic belief. Three cointegrating vectors will only be inferred if the tests for the null hypotheses of r=0 and r=1 and r=2 are *all* rejected and that r equals one or two cannot be supported. This is because r=3 suggests all the variables are stationary, which is inconsistent with the OECD countries' observed consumption and income series being clearly trended - see the data plots and ADF tests reported in Chapter 3.

We report the cointegration results for both standard and favoured models for all countries. However, misspecification, according to the system tests reported in Table 4.1, is evident in the standard model for seven countries (Australia, Canada, Germany, Ireland, Italy, Norway and Sweden). The invalid inference that one would obtain from these misspecified models is that there is little or no cointegration (see Table 4.2). We only draw inference from those models where misspecification is not evident.

For all twenty countries evidence of *at least* one cointegrating vector can be uncovered using either the standard or favoured model. For Belgium one cointegrating vector can only be inferred for the standard model with restricted intercept using the ten percent level of significance. In the case of Sweden we had to search for a specification to secure the inference of cointegration (this model features no dummy variables, L=3 and the intercept is restricted into the cointegration equation).

For nineteen of the twenty countries there is at least one form of VECM where *exactly* one cointegrating vector can be inferred. The exception is Finland, where the standard model indicates no cointegration and the favoured specification suggests two or three equilibrium relations exist. Berg (1994) argues that Finland, in particular, suffered severely from financial deregulation and subsequent re-regulation in the late 1980s and early 1990s. After the deregulation fuelled boom which caused consumption to exceed income there was a damagingly deep slump. The dramatic nature of events in Finland may mean that the omission of wealth and deregulation variables from its VECM has an especially detrimental impact upon inference, explaining the difficultly in securing a unique cointegrating vector.

In the countries where the impact of dummy variables can be isolated (when the lag length is the same in the favoured and standard models) their inclusion causes the inferred number of cointegrating vectors to rise for five countries (Canada, Finland, Germany, Italy, and the UK), to fall for one country (Japan) and to stay the same for one country (Sweden). Rejecting the null of fewer cointegrating vectors appears to increase with the addition of dummy variables. However, one might expect that the addition of variables to the VECM, when using small samples, would reduce the power of the test (lower the probability of rejecting the null). Since the reverse has generally happened, low power would not seem to be a major factor influencing the results. Further, one could argue that the introduction of dummy variables yields results more consistent with our prior beliefs. For example, the introduction of dummies clearly changes the inference from no cointegration to evidence of long run relations for Finland, Italy and the UK. A similar, if less clear change in inference occurs for Germany. For Japan, introducing dummy variables causes the general inference of the number of cointegrating vectors to fall from 2 to 1, which is consistent with a move toward our prior beliefs.[28]

In addition, for seven countries (Finland, Ireland, Italy, the Netherlands, Norway, Sweden and the UK) one can reject the existence of a cointegrating relationship using the standard model (L=2 and no dummy variables). For four of these countries (Finland, Ireland, Italy and Norway), evidence cannot be presented to *support* cointegration without the incorporation of dummy variables in the VECM. Indeed, the standard models are misspecified for all four of these countries.

The introduction of dummy variables generally appears to remove misspecification, account for important unmodelled events and yield inference more consistent with our prior belief of the existence of one unique cointegrating relationship. This is consistent with the intuitively appealing view that improving model specification yields more *sensible* results.

4.4 Selecting Favoured Cointegrating Vectors for Each Country

Evidence supporting cointegration has been presented for all twenty countries and for nineteen,

[28] We note that critical values of these cointegration tests would be expected to alter with the introduction of dummy variables, which we do not account for when drawing inference.

exactly one long run relation can be justified. However, dummy variables are required to secure these results for some countries. For many of these countries we have a choice of cointegrating vectors arising from the various different specifications of lag length, whether the intercept is restricted into the cointegrating vector or not and whether dummy variables are included in the system or not. A choice also arises when there is evidence that more than one cointegrating vector exists or not. To draw inferences about a particular country's long run consumption function, to compare consumer behaviour across countries and to build error correction models, we need to select a *single* favoured long run consumption function for each country. To this end, we first outline the criteria used to select the favoured long run consumption functions and then we apply these criteria to the cointegrating vectors uncovered for each country.

4.4.1 Criteria for Selecting Favoured Long Run Consumption Functions

We employ both statistical and economic criteria to select our favoured long run consumption functions - the need to employ both statistical and economic criteria arises because of the difficulty in uniquely identifying the cointegrating vector and adjustment coefficients when using the Johansen procedure.[29] The statistical criteria are hypothesis tests placed on both the identified cointegrating relations (β) and corresponding adjustment coefficients (α). To consider

[29] Hall and Patterson (1992) note that although Π is uniquely estimated using the Johansen procedure the partition into α and β is not unique. However, they argue that this does not affect the *test* for cointegration because it is the rank of Π which is important for revealing the number of cointegrating vectors. Typically the Johansen procedure is employed to identify the number of cointegrating vectors and, if unique, the estimated long run relations are checked for robustness using alternative estimation methods which control for endogeneity and serial correlation (see for examples, Phillips and Hansen, 1990, Phillips and Loretan, 1991 and Stock and Watson, 1993) and, in addition, short run dynamics, though not necessarily endogeneity (see, Cuthbertson and Gasparo, 1993 and 1995). Swamy and Tavlas (1992) argue that "It might be tempting to propose that one is not really interested in Π, *per se*, the main focus being the discovery of the existence of some co-integrated relationships, whatever the value of Π. But if the posited relationship, being non-unique, cannot describe the real world, then any suppositions concerning the existence of equilibrium relationships, based on it, also do not describe the real world." (Swamy and Tavlas, 1992, pp. 21-22). We follow the standard Johansen procedure adopting both statistical and economic criteria to check for the plausibility of uncovered long run consumption functions, and, in Chapter 5, develop error correction models. Given the historical success of the error correction methodology we maintain the standard interpretation of our models in the absence of alternative explanations, whilst noting the above caveats.

tests on the cointegrating relations we outline the various potential forms of these vectors. First, when a constant term does not enter in the cointegrating vector, called an unrestricted intercept and denoted with a "U" subscript, the general form of cointegrating relation is expressed as equation **(4.4a)**. When the intercept is restricted into the equilibrium relation, denoted with an "R" subscript, the cointegrating vector is specified as **(4.4b)**.

$$Z_{U,rt} = -\beta_{U,r1}\ln C_t + \beta_{U,r2}\ln Y_t + \beta_{U,r3}\Delta \ln P_t, \qquad \textbf{(4.4a)}$$

$$Z_{R,rt} = -\beta_{R,r1}\ln C_t + \beta_{R,r0} + \beta_{R,r2}\ln Y_t + \beta_{R,r3}\Delta \ln P_t. \qquad \textbf{(4.4b)}$$

The subscript "r", denotes the cointegrating vector upon which the tests are conducted, that is, whether it is the first (r=1) or second (r=2) vector.[30] The first of the three general sets of hypotheses to be considered are tests of zero restrictions on the parameters in these cointegrating vectors. For this we use the standard likelihood ratio (LR) statistic to test the statistical significance of each parameter in the error correction term, $Z_{h,rt}$, where h=U,R. When there is a unique cointegrating vector (r=1) this involves testing the single hypothesis:

$$H_0: \beta_{h,1i} = 0. \qquad \textbf{(4.5a)}$$

When r=2 we are only able to conduct tests of the significance of a single variable on *both* cointegrating vectors. We therefore test the following joint hypothesis:

$$H_0: \beta_{h,1i} = \beta_{h,2i} = 0. \qquad \textbf{(4.5b)}$$

If **(4.5a)** or **(4.5b)** can be rejected this suggests that the variable in question is statistically significant in the cointegrating vector(s).[31] Given cointegration, the statistical significance of the

[30] Although our prior belief is that there exists one cointegrating vector we also consider the second cointegrating vector if the first is inconsistent with plausible equilibrium consumption behaviour.

[31] When r=2 this suggests the variable features joint statistical significance in both cointegrating vectors. From tests on the first cointegrating vector one can determine the

parameters in the cointegrating vector indicate which variables adjust to make the long run relationship hold. Because the focus of our interest is in the determination of consumption, this variable's coefficient must be statistically significant for the cointegrating vector to represent a long run consumption function. Since income is theorised to be the main determinant of consumption its parameter would also be expected to be statistically significant for a well defined long run relationship to exist. If inflation enters the long run relationship with statistical significance one would expect its coefficient to be negative. Whether the intercept should be included is treated as a statistical matter, however, if there is no inflation effect and consumption were homogeneous of degree one in income, a significant negative intercept would be required to allow the long run average propensity to consume (APC) to be below one, so ensuring consistency with observed positive aggregate saving.

For r=1, consumption is homogeneous of degree one in income (unit income elasticity), in the long run, if the following hypothesis cannot be rejected.

H_0: $\beta_{h,11} + \beta_{h,12} = 0.$ **(4.6a)**

When r=2 the corresponding joint hypothesis is:

H_0: $\beta_{h,11} + \beta_{h,12} = 0$ and $\beta_{h,21} + \beta_{h,22} = 0.$ **(4.6b)**

Rejection of **(4.6a)** or **(4.6b)** indicates that this homogeneity postulate is inconsistent with the data. This might be expected for the reasons given in Bollerslev and Hylleberg (1985),[32] and/or

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statistical significance of a variable in that first vector, however, one cannot always deduce whether such a variable is significant in the second.

[32] Bollerslev and Hylleberg (1985) outline four potential explanations for consumption exhibiting a below unit income elasticity in the long run. They are summarised by the following quote. "The causes of the downward sloping *APC* are more difficult to find even if one may resort to an explanation based on a variation of Engel's law, i.e. postulating that the consumption expenditures on non-durables have an income elasticity below 1 at the income level for the estimation period. Another explanation put forward by Deaton (1977) is that the savings ratio increases during periods of accelerating inflation due to a mass illusion as to the absolute price level which is caused by the inability on the side of the consumers to separate relative and

if consumption is homogeneous of degree one in lifetime resources (income and wealth) - see Molana (1989).[33] The unit income elasticity hypothesis can be rejected in two ways. The coefficient on income can be less than one (below unit income elasticity) or above unity (above unit income elasticity). Since, in aggregate, consumers cannot spend more than they earn in the long run, an above unit income elasticity is considered implausible. Thus, if the unit income elasticity hypothesis is rejected we expect the coefficient on income to be less than one to reflect defensible consumer behaviour.

For valid error correction behaviour the coefficient on the target variable must be negative. Since our focus is on consumer behaviour we require the estimated parameter on the log of consumption to be negative ($\pi_{h,r1}<0$). We choose to normalise upon consumption in the cointegrating vectors, **(4.4a)** and **(4.4b)**, by setting $-\beta_{h,r1} = -1$, so yielding directly interpretable coefficients on the other parameters in the cointegrating vector. Given the coefficient on the log of consumption is, $\pi_{h,r1} = (\alpha_{h,r1})(-\beta_{h,r1})$, in the restricted VECM, this implies that the adjustment coefficient, $\alpha_{h,rk}$, must be *positive* for valid error correction behaviour in the consumption function. That is, in the *standard* form of the restricted VECM (L=2, r=1 and excluding dummy variables), given as equation **(4.7)**, we require $\alpha_{h,11} > 0$.

$$\Delta \ln C_t = a_{10} + \delta_{11}\Delta \ln C_{t-1} + \delta_{12}\Delta \ln Y_{t-1} + \delta_{13}\Delta\Delta \ln P_{t-1} + \alpha_{h,11}Z_{h,rt-1} + u_{1t} \quad \textbf{(4.7)}$$

$$\Delta \ln Y_t = a_{20} + \delta_{21}\Delta \ln C_{t-1} + \delta_{22}\Delta \ln Y_{t-1} + \delta_{23}\Delta\Delta \ln P_{t-1} + \alpha_{h,12}Z_{h,rt-1} + u_{2t}$$

$$\Delta\Delta \ln P_t = a_{30} + \delta_{31}\Delta \ln C_{t-1} + \delta_{32}\Delta \ln Y_{t-1} + \delta_{33}\Delta\Delta \ln P_{t-1} + \alpha_{h,13}Z_{h,rt-1} + u_{3t}$$

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absolute price rises. This explanation is supported by the decreasing trend in the total consumption expenditure-income relation. A third and somewhat different explanation is given by HUS, who postulate that perceived income is not Y_t but Y_t minus a fraction of the change in the value of net liquid assets. The fourth and final explanation considered here is that there has been a shift towards durables in the long run relation due to the increase in the relative prices of non-durables to durables." (Bollerslev and Hylleberg 1985, pp. 155-156; my italics).

[33] If the coefficients on the logs of consumption and wealth equal unity (homogeneity of degree one in lifetime resources), making the innocuous assumption that the wealth elasticity is positive, suggests that the income elasticity is below one, so rejecting the unit income elasticity hypothesis.

We also require the adjustment coefficient, $\alpha_{h,11}$, to be statistically significant, which can tested by the hypothesis specified by **(4.8)**. This test for long run (or error correction) Granger non-causality (LRGNC) can also be applied to the adjustment coefficients in the non-consumption equations of **(4.7)**. Therefore the general hypothesis for LRGNC is specified as:

$$H_0: \alpha_{h,rk} = 0. \tag{4.8}$$

If **(4.8)** holds the error correction term, $Z_{h,rt-1}$, should not enter the kth equation in the restricted VECM, **(4.7)**. That is, the kth equation is not Granger caused by the long run information incorporated in the error correction term. For the consumption equation we expect **(4.8)** to be rejected for valid error correction behaviour. If the LRGNC hypothesis can be rejected in the income and/or inflation equations of **(4.7)** this implies that consumption has a feedback effect upon these two variables, suggesting violation of weak exogeneity (see, for example, Charemza and Deadman, 1997), and the need to allow for the simultaneous determination of these three variables.[34] For some countries we conduct an analogous test for LRGNC using a non-standard form of restricted VECM which, in particular, jointly tests for the statistical significance of the adjustment coefficients on two cointegrating vectors (when r=2).[35]

4.4.2 Selecting a Favoured Long Run Consumption Function For Each Country

We have a strong economic prior that there exists a *unique* cointegrating vector between the log of consumption, the log of income and inflation. When we cannot reject statistical support for

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[34] We discuss weak exogeneity and its implications in greater detail in Chapter 5.

[35] We also conduct LRGNC tests in non-standard forms of the restricted VECM for certain countries. The most general form of the restricted VECM may be specified as:

$$\Delta \ln C_t = a_{10} + \Sigma\delta_{11i}\Delta \ln C_{t-i} + \Sigma\delta_{12i}\Delta \ln Y_{t-i} + \Sigma\delta_{13i}\Delta\Delta \ln P_{t-i} + \Sigma\phi_{1j}D_{1jt} + \alpha_{h,11}Z_{h,1t-1} + \alpha_{h,21}Z_{h,2t-1} + u_{1t}$$
$$\Delta \ln Y_t = a_{20} + \Sigma\delta_{21i}\Delta \ln C_{t-i} + \Sigma\delta_{22i}\Delta \ln Y_{t-i} + \Sigma\delta_{23i}\Delta\Delta \ln P_{t-i} + \Sigma\phi_{2j}D_{2jt} + \alpha_{h,12}Z_{h,1t-1} + \alpha_{h,22}Z_{h,2t-1} + u_{2t}$$
$$\Delta\Delta \ln P_t = a_{30} + \Sigma\delta_{31i}\Delta \ln C_{t-i} + \Sigma\delta_{32i}\Delta \ln Y_{t-i} + \Sigma\delta_{33i}\Delta\Delta \ln P_{t-i} + \Sigma\phi_{3j}D_{3jt} + \alpha_{h,13}Z_{h,1t-1} + \alpha_{h,23}Z_{h,3t-1} + u_{3t}$$

where i=1,...,L and j=1,...,J-1. The general form of joint hypothesis for the test for LRGNC is:

H_0: $\alpha_{h,1k} = \alpha_{h,2k} = 0$.

a single cointegrating relation between these variables we consider that four of the criteria outlined above need to be satisfied for a plausible long run consumption function to have been revealed. They are, first, that the adjustment coefficient in the consumption equation is positive and statistically significant and, second, that consumption and income are statistically significant. A third criteria that we apply is that *when* inflation is significant it should have a negative coefficient. A fourth criteria which we believe should hold is that the equilibrium APC should be less than one to reflect the persistence of positive aggregate saving observed for OECD countries. That is, there is either a below unit income elasticity or, if the income elasticity is not significantly different from one, there should be a statistically significant and negative intercept or inflation term.

When, for any particular country, we are unable to find statistical *support* for a *unique* cointegrating vector, we examine the possibility that there exist two distinct cointegrating relations. Pesaran and Shin (1994) demonstrate that the necessary condition for the exact identification of cointegrating vectors in a Johansen system, analogous to the order condition, is that one needs to impose (m=) r^2 restrictions on the long run coefficients. When there is only one cointegrating vector this typically involves imposing a normalisation restriction, which may be interpreted as choosing which variable constitues the dependent variable. However, when $r>1$ one will need to apply other (typically exclusion) restrictions. For example, when $r=2$, one will need to impose (m=) 4 restrictions to exactly identify the two separate cointegrating relations of which, only two, can be normalisation restrictions.[36] Within the context of the DHSY model Larsson *et al* (1998) suggest one possible set of (over) identification restrictions when $r=2$. They suggest that consumption and income may constitute one cointegrating vector and that inflation may be stationary, so constituting a second cointegrating relation. This involves placing two normalisation retrictions (one on each vector) and three exclusion restrictions: $\boldsymbol{\beta}_{h,13}=0$, $\boldsymbol{\beta}_{h,21}=0$ and $\boldsymbol{\beta}_{h,22}=0$, on **(4.4a)** or **(4.4b)**. Since the number of restrictions, m (=5), exceeds r^2 (=4), this produces an overidentified long run matrix which, following Pesaran and Shin (1994), can be tested using using an LR test which follows a χ^2 distribution with m-r^2 (in this case one) degrees

[36] Pesaran and Pesaran (1997) further point out that one must apply "at least r independent restrictions on each of the r cointegrating vectors." (Pesaran and Pesaran, 1997, p. 439).

of freedom. Therefore, if we find that r=2 (and there is no plausible unique cointegrating vector for consumption) for any particular country then, if the above overidentification retriction cannot be rejected, whilst the estimated parameters on the long run consumption equation are plausible, the overidentified consumption equation will represent our favoured long run consumption function.[37]

We will first consider if these criteria are satisfied for any of the OECD countries examined here. Table 4.3 reports the identified adjustment coefficients, cointegrating vectors and test statistics for the above hypotheses, for selected VECM specifications where cointegration was indicated. We do not report this information for all possible specifications where cointegration cannot be rejected, both to save space and because there are many cointegrating relations which obviously cannot constitute long run consumption functions. When there is support for cointegration we report the standard specification (L=2, r=1 and excluding dummy variables) and the best non-standard specification. Further, if the best non-standard model's cointegrating relation appears to be the second of two vectors (when r=2 can be justified from the results reported in Table 4.2) we also report the first vector for comparative purposes. For some countries the standard model either suffers from evident misspecification (according to the system tests reported in Table 4.1) or does not produce cointegration. In these cases the phrase "Standard model is misspecified" indicates a model suffering from misspecification; "No cointegration without dummy variables" signifies those countries where cointegration is rejected if dummy variables are excluded from the VECM; and "No cointegration when L=2" indicates those countries where cointegration is rejected when there are two lags in the VECM.

The first column of Table 4.3 indicates the country to which the reported results refer and the second specifies the dummy variables included in the VECM - the favoured long run consumption function is denoted by an **F**. The third column gives the number of lagged variables (L) used in the VAR, the fourth stipulates whether the intercept is restricted (R) into the

[37] It is not obvious that any other form of (over) identification restriction would provide an economically sensible combination of cointegrating vectors, for example, we would not expect consumption to form a long run relationship solely with inflation. Therefore, we do not consider a different form of overidentification restriction when r=2.

TABLE 4.3: Tests on the VECM's Long Run Matrix ($\Pi=\alpha\beta'$)

VECM Specification						α			β				
			Lags	Int		$\alpha_1=0$	$\alpha_2=0$	$\alpha_3=0$	$\beta_0=0$	$\beta_1=0$	$\beta_2=0$	$\beta_3=0$	$\beta_1+\beta_2=0$
Country	Dummies		(L)	U/R	r=	ΔlnC	ΔlnY	$\Delta\Delta$lnP	Int	lnC	lnY	ΔlnP	Unit
AUL	NONE		2			STANDARD MODEL IS MISSPECIFIED							
	NONE	F	3	R	1	+0.281			0.526	-1.000	1.107	-1.011	
						(14.455)	(3.304)	(0.056)	(1.624)	(2.188)	(2.072)	(3.412)	(1.135)
AUT	NONE		2	U	1	+0.548				-1.000	0.869	-0.154	
						(4.745)	(0.988)	(0.839)		(4.752)	(4.444)	(0.374)	(6.605)
	74;78	F	1	R	1	+0.611			-0.502	-1.000	0.864	-0.209	
						(57.698)	(40.554)	(0.880)	(14.786)	(13.167)	(12.109)	(1.084)	(21.808)
BEL	NONE	F	2	R	1	+0.127			-0.226	-1.000	0.866	0.908	
						(6.168)	(1.655)	(9.460)	(1.206)	(0.953)	(0.797)	(1.259)	(2.950)
CAN	NONE		2			STANDARD MODEL IS MISSPECIFIED							
	76;82;91	F	2	R	1	+0.595			-0.353	-1.000	0.931	-0.703	
						(15.128)	(3.887)	(10.061)	(10.851)	(8.533)	(8.160)	(4.531)	(12.115)
DEN	NONE		2	U	1	-0.004				-1.000	27.960	63.290	
						(7.717)	(14.984)	(0.571)		(0.002)	(1.334)	(4.441)	(16.029)
	NONE		2	R	1	+0.011			-20.040	-1.000	-8.003	-19.640	
						(17.016)	(27.408)	(0.351)	(20.112)	(0.021)	(1.113)	(4.086)	(21.355)
	NONE	F	3	U	1	+0.047				-1.000	1.464	1.926	
						(0.249)	(0.242)	(6.324)		(4.151)	(7.036)	(10.225)	(8.374)
FIN	NONE		2			NO COINTEGRATION WITHOUT DUMMY VARIABLES							
	69;72;74	F	2	U	1	+0.330				-1.000	1.075	-0.231	
						(2.206)	(2.561)	(0.023)		(4.265)	(4.255)	(1.053)	(3.906)
FRA	NONE		2	R	1	-0.101			0.750	-1.000	1.344	-0.754	
						(16.614)	(16.800)	(4.987)	(8.728)	(0.577)	(1.060)	(0.378)	(13.707)
	74	F	1	U	1	+0.226				-1.000	0.844	-0.878	
						(24.518)	(18.210)	(2.430)		(4.743)	(3.294)	(4.190)	(17.238)
GER	NONE		2			STANDARD MODEL IS MISSPECIFIED							
	91	F	2	R	1	+0.308			-0.159	-1.000	0.975	0.038	
						(10.444)	(5.878)	(0.230)	(1.417)	(4.933)	(4.265)	(0.002)	(0.425)
GRE	NONE		2	U	1	+0.408				-1.000	0.912	-0.902	
						(8.360)	(8.474)	(0.398)		(4.035)	(4.073)	(8.317)	(2.392)
	NONE	F	1	R	1	+0.519			-0.110	-1.000	0.904	-0.632	
						(44.627)	(26.338)	(5.242)	(3.805)	(17.113)	(17.391)	(19.835)	(8.912)
ICE	NONE	F	2	R	1	+0.414			-0.203	-1.000	1.053	0.113	
						(2.291)	(0.105)	(0.011)	(7.054)	(4.756)	(3.986)	(0.690)	(0.391)
IRE	NONE		2			NO COINTEGRATION WITHOUT DUMMY VARIABLES							
	73;82		1	R	1	-0.657			-0.040	-1.000	1.019	-0.113	
						(7.546)	(13.641)	(0.506)	(0.151)	(4.944)	(5.065)	(0.501)	(1.239)
	73;82		1	R	2	*+0.680*			*-0.081*	*-1.000*	*1.001*	*-0.373*	
						(31.098)	*(38.150)*	*(0.587)*	*(0.910)*	*(30.907)*	*(31.239)*	*(12.633)*	*(1.246)*

Table 4.3 Continued

ITA	NONE	2			STANDARD MODEL IS MISSPECIFIED							
	93 **F**	2	R	1	+0.027			-0.945	-1.000	0.569	-3.645	
					(10.854)	(12.227)	(0.182)	(0.202)	(0.091)	(0.021)	(1.880)	(0.474)
JAP	NONE	2	R	1	-0.353			0.219	-1.000	1.068	-1.288	
					(8.931)	(4.135)	(17.130)	(1.064)	(4.117)	(4.925)	(4.528)	(4.567)
	NONE	2	R	2	*+0.351*			*-0.596*	*-1.000*	*0.921*	*-1.502*	
					(20.240)	*(7.059)*	*(30.748)*	*(14.686)*	*(10.223)*	*10.408)*	*(15.606)*	*(16.463)*
	74	2	R	1	-0.211			-0.055	-1.000	1.020	-1.688	
					(7.417)	(1.319)	(28.851)	(0.112)	(6.889)	(7.370)	(13.003)	(0.648)
	74 **F**	2	R	2	*+0.305*			*-0.620*	*-1.000*	*0.916*	*-1.348*	
					(16.317)	*(2.159)*	*(41.997)*	*(11.176)*	*(12.686)*	*(12.353)*	*(23.917)*	*(8.831)*
NET	NONE	2			NO COINTEGRATION WHEN L=2							
	NONE **F**	1	R	1	+0.451			-0.564	-1.000	0.880	-0.248	
					41.304	30.853	0.290	16.028	3.044	2.394	0.814	20.608
NOR	NONE	2			STANDARD MODEL IS MISSPECIFIED							
	708081;78; 8586 **F**	1	R	1	+0.420			-0.101	-1.000	0.958	0.138	
					(40.718)	(23.266)	(2.419)	(1.445)	(14.373)	(12.685)	(0.307)	(2.035)
SPA	NONE	2	U	1	-0.200				-1.000	1.075	-0.011	
					(2.285)	(19.330)	(1.215)		(12.201)	(13.467)	(0.006)	(25.367)
	74;77 **F**	1	U	1	+0.769				-1.000	1.037	-0.382	
					(7.028)	(0.355)	(7.841)		(17.270)	(17.145)	(11.342)	(5.959)
SWE	NONE	2			STANDARD MODEL IS MISSPECIFIED							
	NONE **F**	3	R	1	-0.098			0.827	-1.000	1.389	1.012	
					(3.918)	(17.292)	(1.406)	(2.285)	(3.144)	(3.726)	(1.122)	(3.952)
SWZ	NONE	2	R	1	+0.007			-7.811	-1.000	-1.184	11.130	
					(1.751)	(0.025)	(9.127)	(3.684)	(0.042)	(0.071)	(1.094)	(3.903)
	6386;71;79 **F**	2	R	1	-0.032			0.963	-1.000	1.346	-1.858	
					(1.970)	(0.230)	(7.616)	(1.195)	(0.669)	(1.363)	(0.423)	(2.287)
UK	NONE				NO COINTEGRATION WHEN L=2							
	74;75	2	U	1	-0.032				-1.000	1.054	0.122	
					(0.044)	(1.046)	(0.562)		(0.665)	(0.691)	(0.061)	(1.244)
	74;75	2	U	2	*+0.081*				*-1.000*	*1.016*	*-0.544*	
					(0.344)	*(5.844)*	*(10.655)*		*(15.209)*	*(15.378)*	*(13.039)*	*(5.297)*
	NONE **F**	3	U	1	+0.126				-1.000	1.038	-0.250	
					(0.263)	(6.822)	(0.102)		(20.570)	(20.684)	(9.772)	(9.257)
USA	NONE **F**	2	R	1	+0.642			0.221	-1.000	1.059	-0.623	
					(18.180)	(9.611)	(7.028)	(2.329)	(6.831)	(6.497)	(4.604)	(2.516)
5% Critical Values: $\chi^2(r)$				r=1	*3.84*	*3.84*	*3.84*	*3.84*	*3.84*	*3.84*	*3.84*	*3.84*
				r=2	*5.99*	*5.99*	*5.99*	*5.99*	*5.99*	*5.99*	*5.99*	*5.99*

Table 4.3 Notes: Dummy variables are specified as for Tables 4.1 and 4.2 - an **F** denotes a country's favoured cointegrating vector. Int refers to whether the intercept is restricted into the cointegration space (R) or not (U). r= refers to the number of the cointegrating vector, where r=2 means the results refer to the second of two long run

relations (indicated with italic emphasis). The estimated adjustment coefficients for the consumption growth equation is reported in column six. The estimated cointegrating vectors (normalised upon consumption) are reported in columns nine to twelve. Likelihood ratio tests for the statistical significance of adjustment coefficients (in all three equations) and the estimated parameters are reported below their corresponding coefficients (where reported) in brackets. The thirteenth column (headed "Unit") reports the test statistic for the hypothesis that consumption is homogeneous of degree one in income. The test statistics follow a chi-square distribution with r degrees of freedom, critical values are reported at the bottom of the Table. Bold emphasis indicates rejection of the null hypothesis: that a variable is statistically significant or the rejection of the homogeneity postulate, depending upon context. Bold emphasis is also used to indicate that the adjustment coefficient in the consumption growth equation exhibits the *correct* positive sign.

cointegrating vector or not (U) while the fifth column denotes whether the results refer to the first or second cointegrating vector (r=). The next three columns report, in brackets, the LR test statistic for the statistical significance of the adjustment coefficient in the consumption growth (α_1=0), income growth (α_2=0) and change in inflation (α_3=0) equations. The estimated adjustment coefficient is reported above the LR statistic for the consumption equation.[38] Columns nine to twelve report the estimated coefficients of the intercept (β_0=0), consumption (β_1=0), income (β_2=0) and inflation (β_3=0) terms in the cointegrating vector. Corresponding LR test statistics for these coefficients' statistical significance are given in brackets below them. The consumption parameter is normalised to be minus one. The final column reports the LR test statistic for the unit income elasticity hypothesis ($\beta_1+\beta_2$=0). Bold emphasis indicates the rejection of a hypothesis and a *correctly* signed adjustment coefficient in the consumption equation. Italic emphasis indicates that the reported information refers to the second cointegrating vector, where tests are conducted on *both* first and second vectors.

The standard model (r=1, L=2 and no dummy variables) satisfies all of the four criteria for a plausible long run consumption function (specified above) for only two of the twenty countries (Austria and the USA). Three countries (Austria, Canada and Greece) satisfy all these criteria when nonstandard models are employed.[39] The standard and nonstandard models for Austria

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[38] Adjustment coefficients are not reported for the income growth and change in inflation equations because their *values* are of no interest in the current investigation.

[39] In the case of Canada the favoured model features a restricted intercept, L=2 and dummy variables. From Table 4.2 this model appears to feature two cointegrating vectors regardless of whether one employs the Maximum Eigenvalue or Trace statistic or the one or five percent level of significance. However, if one uses the trace statistic adjusted for degrees of

yield similar results, however, we prefer the nonstandard form because it incorporates a statistically significant intercept in the cointegrating vector, which is not present in the standard model. We therefore favour the specified nonstandard models for Austria, Canada and Greece and the standard model for the USA.

For six of the twenty countries (France, Germany, Iceland, the Netherlands, Norway, and the UK) we can identify a specification where only one of the four desirable conditions is not satisfied. These represent the best obtainable specifications for these countries and, therefore, represent their favoured long run consumption functions. We outline the criteria which is *not* satisfied for each country. In the case of France the income term is just statistically insignificant while both consumption and income are insignificant for the Netherlands. For Germany and Norway the unit income elasticity hypothesis cannot be rejected and both the intercept and inflation terms are statistically insignificant, suggesting that the APC is unity in the long run (there is no aggregate saving).[40] The adjustment coefficient in the consumption equation is statistically insignificant for Iceland and the UK (although it is positive for both countries).[41] Although not satisfying all the specified criteria these six countries' cointegrating vectors are plausible in many senses and are presented as reasonable approximations of their countries' long run consumption functions. For only one of these countries (Iceland) is the standard model favoured.

For four of the remaining countries (Australia, Belgium, Finland and Italy) two of these

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freedom, which is 58.40 for the null hypothesis of $r=0$ and 23.92 for the null hypothesis of $r=1$, one cannot reject the inference of a unique cointegrating vector at the 1% level. Although Doornik and Hendry (1995) note that it is not yet clear whether this is the preferred small sample correction (see p. 222) we utilise this result to provide statistical support for our strong prior economic belief of a single cointegrating vector. Further, we find that the overidentification restrictions when applied (assuming $r=2$) are rejected (the test statistic is 6.158).

[40] The coefficients on income for Germany and Norway are both less than one, if not statistically different from unity, so *may* be considered completely plausible.

[41] The three best cointegrating vectors reported for the UK provide very similar inference. However, the specification without dummy variables, with unrestricted intercept and where L=3 is favoured because it is Table 4.2 suggests that it is a unique cointegrating vector. In contrast, the other pair of reported vectors come from a VECM which suggests $r=2$.

plausibility criteria are not met. For only one country (Belgium) is the standard specification favoured. In the case of Australia both consumption and income terms are statistically insignificant and there is no significant and negative intercept or inflation term to compensate for the evidence against the presence of a below unit income elasticity. However, the adjustment coefficient in the consumption equation is positive and statistically significant and the cointegrating vector's estimated parameters are plausible, if not well determined, including a negative, if statistically insignificant, coefficient on the inflation term.

The favoured cointegrating vectors for Belgium and Italy are comprised of statistically insignificant (β) coefficients (the consumption and income terms are insignificant). Further, although the coefficient on income is less than one for both countries, it is not (statistically) significantly less than one, implying a unit long run APC because both intercept and inflation terms are statistically insignificant. However, the adjustment coefficient in the consumption equation is positive and statistically significant and the estimated cointegrating vector is plausible for both countries, if the income elasticity is quite low for Italy (being 0.569).[42]

For Finland the adjustment coefficient in the consumption equation is statistically insignificant, if featuring the *correct* positive sign.[43] Although consumption and income are both statistically significant in the cointegrating vector, there is evidence that the income elasticity is significantly greater than one.[44]

The cointegrating vectors for Australia, Belgium, Finland and Italy are presented as usefully

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[42] Although this low income elasticity is consistent with Italy historically exhibiting a low APC (see Guiso *et al* 1991) it may also be due to this parameter's poor determination.

[43] There is *support* for zero or three cointegrating vectors for Finland's favoured cointegrating vector (see Table 4.2). In this case the statistics seem completely unhelpful regarding the choice of the value of r so we impose our strong economic prior belief of r=1. Pesaran and Pesaran (1997) similarly impose theoretical priors when the Johansen procedure is found to be "hopelessly uninformative on the choice of r" (Pesaran and Pesaran 1997, p. 297) when modelling UK exchange rates.

[44] The estimated income elasticity being greater than unity for Finland may be due to the omission of explanatory factors capturing the effects of financial deregulation.

plausible because they exhibit many desirable features for credible long run consumption functions and their departures from the specified criteria do not seem too severe.

The favoured cointegrating vector for Denmark fails to satisfy three of the desirable features specified above. There is evidence of an above unit income elasticity (with a rather high income elasticity of 1.464), the adjustment coefficient is statistically insignificant, if it exhibits the *correct* positive sign and the coefficient on inflation has the *incorrect* positive sign. However, all the estimated coefficients in the cointegrating vector are statistically significant and the adjustment coefficient in the consumption growth equation exhibits the *correct* positive sign. Therefore, we believe this vector represents an approximate long run consumption function for Denmark, if we have some reservations.

For Sweden and Switzerland, the most plausible cointegrating vector fails to satisfy many of the above specified criteria to be interpretable as a long run consumption function. The adjustment coefficient in the consumption equation exhibits a negative coefficient, which suggests that consumption is continually forced *away* from its equilibrium, for both countries. All of the variables are statistically insignificant for both economies (although the test statistics on consumption and income are greater than three for Sweden, which provides some encouragement). The estimated income elasticities are rather high (1.389 and 1.346 for Sweden and Switzerland, respectively).[45] This income elasticity is significantly greater than one for Sweden but not significantly different from unity for Switzerland. In neither country are there statistically significant and negative intercepts or inflation terms to allow a below unit long run APC. That these countries' cointegrating vectors have large and poorly determined income elasticities and are inconsistent with error correction behaviour suggest that they provide poor approximations to a credible long run consumption function. This is disappointing because there are no better alternative cointegrating vectors for either country.

For the above countries we provided *support* for the existence of a unique cointegrating vector.

[45] These large estimated income elasticities may reflect the income coefficient's poor determination and/or, in the case of Sweden, be due to financial deregulation.

For Ireland, Japan and Spain we could not uncover statistical support for a *plausible* unique cointegrating relation.[46] However, models indicating evidence for two cointegrating vectors were presented for all three countries in Table 4.2. For each country we apply the overidentification restrictions that consumption and income form one cointegrating vector and that inflation provides a second, distinct, stationary vector. The results for this test are presented in Table 4.4. The first five columns detail the VECM specification for each country in the same manner as for Table 4.3. Column six (headed "Over-Identification Restrictions $\beta_{13}=0$; $\beta_{21}=0$; $\beta_{22}=0$") presents the test statistic for the overidentification restrictions, with the probability up to which it is statistically insignificant reported underneath in squared brackets. Bold emphasis denotes rejection of these restrictions at the five percent level. The seventh column reports the estimated adjustment coefficient, associated with both restricted cointegrating vectors, in the consumption growth equation of the VAR. Columns eight to eleven give the estimated coefficients of the two restricted cointegrating vectors for each country; where the first vector is normalised on the log of consumption.

The overidentification restrictions are rejected for Japan and Spain but not Ireland. The first cointegrating vector for Ireland is plausible as a long run consumption function in the sense that the adjustment coefficient is positive and the income elasticity is very close to unity (1.010) with a negative intercept (which allows the long run APC to be less than one).[47] This overidentified consumption vector therefore represents our favoured long run consumption function for Ireland.

Although our cointegration tests (Table 4.2) suggest that r=2, the rejection of the overidentification restrictions (Table 4.4) and the findings of Greenslade *et al* (1998) that the Johansen procedure can often indicate *too much* cointegration leads us to impose our economic prior of a unique cointegrating vector for Spain. The favoured Spanish vector is reported in

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[46] The negative adjustment coefficient revealed for the cointegrating vector obtained from Spain's standard model (with unrestricted intercept) is regarded as providing an implausible long run consumption function.

[47] We do not conduct tests of significance on the adjustment coefficients and the parameters in the cointegrating vector because they involve testing several (overidentification) restrictions jointly, so do not simply refer to the hypothesis of interest.

TABLE 4.4: Over-Identification Restrictions when r=2

VECM Specification					Over-Identification Restrictions	α	β			
		Lags	Int			α_{r1}	β_{r0}	β_{r1}	β_{r2}	β_{r3}
Country	**Dummies**	**(L)**	**U/R**	**r**	$\beta_{13}=0$; $\beta_{21}=0$; $\beta_{22}=0$	**ΔlnC**	**Int**	**lnC**	**lnY**	**ΔlnP**
IRE	73;82 **F**	1	R	1	0.565	+0.002	-0.116	-1.000	1.010	
				2	[0.452]	+0.182	0.216			-0.998
JAP	74	2	R	1	**7.360**	+0.210	-0.288	-1.000	0.968	
				2	**[0.007]**	+0.008	-0.371			-11.420
SPA	74;77	1	U	1	**9.544**	+0.573		-1.000	1.043	
				2	**[0.002]**	+0.229				-2.305

Table 4.4 notes: The country to which the results relate, the dummy variables and lag lengths used along with whether the intercept is restricted (R) into the cointegrating vector or not (U) are given in the first four columns. An **F** in the second column denotes a country's favoured cointegrating vector for consumption. The fifth column, headed "r", specifies the restricted and (over) identified cointegrating vector to which the results relate. The sixth column gives the test statistic for the over-identification restrictions (the critical value is 3.84). The probability up to which this statistic is statistically insignificant is given in squared brackets below this statistic. The seventh column provides the estimated adjustment coefficient, associated with both restricted cointegrating vectors, in the consumption growth equation of the VAR. Columns eight to eleven give the two estimated restricted cointegrating vectors for each country; where the first is normalised on the log of consumption. A bold test statistic/probability value indicates rejection of the over-identification restrictions.

Table 4.3 - denoted with an **F**. It is regarded as providing a reasonable approximation to this country's long run consumption function because only one of the four desirable features of a unique vector, outlined above, is not satisfied. This unsatisfied feature is that there is evidence of an above unit income elasticity.

For Japan the overidentification restriction is rejected so we do not favour the overidentified consumption function. However, we do not assume r=1 because the first cointegrating vectors reported for standard and non-standard specifications (Table 4.3) feature statistically significant and negative adjustment coefficients (which is inconsistent with valid error correction behaviour). In contrast, the second vector in the nonstandard model (which incorporates a dummy variable to account for the sharp slowdown in Japanese growth in 1974) satisfies all four of the desirable criteria for a long run consumption function detailed above.[48] We select this second vector as our favoured long run consumption function for Japan despite not strictly being identified. All of the Japanese specifications we could have selected faced some objection, we choose this model due to its desirable theoretical features.

We have chosen a favoured long run consumption function for each country to be that which is the most plausible according to the four criteria outlined above. For only three countries (Belgium, Iceland and the USA) is the standard model the favoured specification. Indeed, for eleven countries (Austria, Canada, Finland, France, Germany, Ireland, Italy, Japan, Norway, Spain and Switzerland) dummy variables were employed in the favoured models' specifications. These probably capture omitted factors such as German reunification, the substantial slowdown in Japanese growth in the 1970s and unmodelled wealth/deregulation effects. These results confirm the inferences drawn by Carruth *et al* (1996) and Pesaran, Shin and Smith (1997) that consumption functions are heterogeneous across OECD countries and that factors beyond consumption, income and inflation would ideally also be considered.

Overall, the long run consumption functions uncovered here are presented as reasonable first approximations, perhaps excepting Sweden and Switzerland. The only recommendation for these two countries' long run consumption functions is that their coefficients' signs and magnitudes are (almost) within the vicinity of what might reasonably be expected. These reservations will be borne in mind when we construct error-correction models and when we analyse the favoured long run consumption functions' general characteristics across the OECD.

4.5 General Characteristics of OECD Countries' Long Run Consumption Functions

This section outlines general similarities and differences of the favoured long run consumption functions identified for each of the twenty OECD countries. We consider whether consumption is homogeneous of degree one in income, whether inflation enters with a negative and

[48] Although we could justify r=1 using the Trace test statistic, as indicated in Table 4.2, r=2 cannot be rejected at the 5% level using the Maximum Eigenvalue test.

statistically significant coefficient, whether weak exogeneity appears to be rejected and whether error correction behaviour is valid.

4.5.1 Is Consumption Homogeneous of Degree One in Income?

The favoured long run consumption functions of six countries (Austria, Canada, France, Greece, Japan and the Netherlands) exhibit a below unit income elasticity,[49] in the sense that the unit income elasticity hypothesis is rejected and the coefficient on income is less than one.[50] The hypothesis that consumption is homogeneous of degree one in income cannot be rejected for nine countries (Australia, Belgium, Germany, Iceland, Ireland, Italy, Norway, Switzerland and the USA).[51] An above unit income elasticity is inferred for five countries (Denmark, Finland, Spain, Sweden and the UK), in the sense that the unit income elasticity hypothesis is rejected and the coefficient on income exceeds one.[52] These results suggest a general heterogeneity of inference regarding the long run unit income elasticity postulate, which is consistent with Pesaran, Shin and Smith's (1997) findings.

On the basis of these results it may be tempting to argue that consumption is, in general, homogeneous of degree one in income for the OECD economies because the unit income elasticity postulate cannot be rejected for nine countries. However, for four of these economies (Australia, Belgium, Italy and Switzerland) the income elasticity is not statistically different

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[49] Another four countries' favoured cointegrating vectors feature coefficients on income below unity (Belgium, Germany, Italy, and Norway) although the unit income elasticity cannot be rejected.

[50] We note that although the χ^2 distribution conducts a one-tail test it does not distinguish whether rejection of the null hypothesis is due to the coefficient being greater or less than the hypothesised value.

[51] In the case of Ireland we do not carry out a statistical test for this homogeneity postulate because our favoured consumption function is an overidentified vector. However, because the estimated income elasticity (1.010) is so close to unity we believe it is safe to assume that there is a unit income elasticity.

[52] The estimated parameter on income exceeds unity for another five countries (Australia, Iceland, Ireland, Switzerland and the USA) while not exceeding one with statistical significance.

from zero either, suggesting that these countries' income elasticities are poorly determined, perhaps explaining the inability to reject the unit income elasticity hypothesis. Indeed, our evidence rejects this postulate for eleven of the twenty countries we consider, suggesting that one should not automatically assume consumption is homogeneous of degree one in income for any particular OECD country. This would be consistent with Carruth *et al*'s (1996) findings which reject this postulate for fourteen of the fifteen EU countries.

The evidence of an above unit long run income elasticity for five countries may be regarded as theoretically implausible and indicative of omitted variables.[53] Thus, it may be that the estimated income elasticities are picking up effects of omitted variables. For example, financial deregulation caused the APC of Finland, Sweden and the UK to rise dramatically during the 1980s. Thus, the above unit income elasticities may reflect the omission of explanatory factors such as wealth and credit from these countries' long run consumption functions.

4.5.2 Is Inflation A Determinant of Long Run Consumption?

For inflation to have a plausible role in itself or as a proxy for wealth effects, its coefficient should be negative and statistically significant. For seven countries (Canada, France, Greece, Japan, Spain the UK and the USA) inflation enters the long run consumption function with a negative coefficient and statistical significance. This is consistent with the evidence presented by Carruth *et al* (1996) and Pesaran, Shin and Smith (1997), both of whom found that inflation was negative and statistically significant in consumption functions for less than half of the countries that they considered. This could indicate that inflation is neither a fundamental determinant of, nor a proxy for wealth in, many OECD countries' models of consumer behaviour.

[53] We do not restrict dummy variables into the cointegration space, hence they do not constitute long run effects.

4.5.3 The Adjustment Coefficient, Weak Exogeneity and Error Correction

If the coefficient on the error correction term enters the income growth and/or change in inflation equations of the restricted VECM with statistical significance then these variables are not weakly exogenous with respect to consumption. For sixteen of the twenty countries (the *exceptions* are Australia, Finland, Iceland and Ireland) the adjustment coefficient is statistically significant in either the income growth or change in inflation equations.[54] This suggests a general need to allow for the simultaneous determination of consumption, income and inflation when building error correction models using this data - see Chapter five.

The adjustment coefficient is positively signed and statistically significant in the consumption growth equation for thirteen countries and positive, if insignificant, for a further five economies (Denmark, Finland, Iceland, Ireland and the UK).[55] This is consistent with consumption being continually forced towards its equilibrium value and the development of error correction models for these countries. However, for the two countries (Sweden and Switzerland) where this adjustment coefficient is negative there is no coherence with the equilibrium value of consumption and one may, therefore, be unable to develop error correction models based upon their identified long run consumption functions.

4.6 Conclusion

The Johansen procedure has been employed to test whether consumption, disposable income and inflation cointegrate in twenty OECD countries. The use of disposable income rather than

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[54] For eleven countries (Austria, Canada, France, Germany, Greece, Italy, the Netherlands, Norway, Sweden, the UK and the USA) the adjustment coefficient is statistically significant in the income growth equation. For eight countries (Belgium, Canada, Denmark, Greece, Japan, Spain, Switzerland and the USA) the adjustment coefficient is significant in the change in inflation equation. We do not conduct tests for Ireland so draw no inference regarding weak exogeneity for this country.

[55] We have not tested the statistical significance of the Irish adjustment coefficient in the (overidentified) consumption growth equation, however, Table 4.4 shows it to be very small (0.002) so we assume that it is insignificant.

proxies such as national disposable income or GDP represents a development of the present study compared to previous work. Various forms of VECM were considered to remove evident misspecification and to allow for heterogeneity of specification across countries. The allowance for both restricted and unrestricted intercepts provides a degree of flexibility in the long run specification while experimentation with the VECM's lag length facilitates unique specification of the models' short run components. Conducting such a specification search represents an improvement of the work by Carruth *et al* (1996) and Pesaran, Shin and Smith (1997) which should enable us to select more appropriate models for each country.

For seventeen countries statistical evidence is presented to *support* the favoured models being unique cointegrating vectors. For Ireland, Japan and Spain the favoured models are obtained where the evidence suggests two cointegrating relations exist. Overidentification restrictions cannot be rejected for Ireland (hence yielding this country's favoured long run consumption function) but are rejected for the other two economies. We choose one of the two estimated cointegrating vectors as the favoured long run consumption functions for Japan (we select the second vector, with some reservations) and Spain (we choose the first vector). We selected the favoured long run consumption function for each country as that which receives most statistical support and which is the most theoretically plausible. For eighteen countries we uncover plausible long run consumption functions, however, the favoured cointegrating vectors for Sweden and Switzerland are not convincing as equilibrium consumption functions. The favoured models appear to be reasonable approximations, providing useful insights into OECD countries' consumer behaviour. In Chapter five these favoured vectors will be used to build dynamic error-correction models for each country, facilitating further cross-country comparisons. Our results suggest that this may not be possible for Sweden and Switzerland.

The favoured long run consumption functions are heterogeneous across countries regarding estimated income and inflation elasticities. There is evidence of a below unit income elasticity for six countries, a unit income elasticity for nine countries and an above unit income elasticity for five countries. The above unit income elasticity in the long run is difficult to justify and possibly reflects omitted variable bias. The impact of omitted variables, the poor determination of some countries' income elasticities and the evidence of a below unit income elasticity for six countries suggests that one should not automatically assume that consumption is homogenous of degree one in income for any particular OECD country. We analyse the cross-country variation of long run income elasticities in Chapter seven.

Inflation is statistically significant and negative in the long run consumption function for seven countries. This suggests that inflation is not a fundamental explanatory factor of consumption for all countries, though may act as a wealth proxy for some economies. In Chapter seven we investigate whether the cross-country variation in the inflation coefficient is consistent with a wealth proxy interpretation.

For sixteen countries there is evidence that income and inflation are not weakly exogenous. We therefore consider the need to account for simultaneity when building error correction models in Chapter five.

CHAPTER 5

ERROR-CORRECTION MODELS OF CONSUMPTION, INCOME AND INFLATION, WITH AND WITHOUT ASYMMETRIC ADJUSTMENT

5.1 Introduction

This chapter seeks to build single equation structural error correction models of consumer behaviour for twenty OECD countries based upon the favoured long run consumption functions developed in Chapter 4. The variables incorporated in our models are, following our analysis in Chapter 4, confined to the log of per-capita total private consumption (lnC), the log of per-capita private disposable income (lnY) and consumer price inflation (ΔlnP). Given that weak exogeneity was rejected in the majority of countries' VECMs we estimate our error correction models using instrumental variables (IV). We also test for weak and strong exogeneity within our structural modelling framework. Parsimonious structural models are sought using the general-to-specific methodology with model selection guided by misspecification and specification tests as well as economic prior beliefs.

We aim to establish whether reasonably specified dynamic consumption functions can be obtained using consumption, income and inflation for our twenty OECD countries. We are only aware of two previous analyses which have attempted to build dynamic error correction models of consumer behaviour, based upon these three variables, for EU or OECD countries, being Carruth *et al* (1996) and Pesaran, Shin and Smith (1997). Our work represents an advancement of these analyses because we use a more appropriate measure of income and a broader model specification - see Chapter 4. The results of Chapter 4 suggest that the empirical importance of inflation in determining long run consumer behaviour is questionable therefore, we wish to determine the importance of inflation as a short run explanatory factor of consumption. In Chapter 7 we conduct a cross-country comparison of the estimates of short run income and inflation elasticities as well as adjustment coefficients produced in this Chapter.

We develop our initial dynamic error correction models to allow for asymmetric/nonlinear adjustment to equilibrium - we are not aware of any previous analysis which has considered asymmetries for OECD countries' consumer behaviour. We investigate two distinct forms of asymmetry to assess whether either is a more appropriate than the *standard symmetric* error correction formulation. The first form of asymmetry that we consider is the Granger and Lee (1989) *partitioned* specification to investigate, primarily, whether consumers adjust their long run consumption at a different speed when they are above equilibrium relative to when they are below it. The second specification of asymmetric adjustment is the *cubic* nonlinear form employed by, for example, Hendry and Ericsson (1991). In addition to allowing different speeds of adjustment when consumption is above equilibrium relative to when it is below it, this functional form also characterises agents as adjusting more rapidly towards equilibrium the greater is the disequilibrium. Because the signs of the estimated coefficients on the cubic function of error correction terms can no longer unambiguously indicate whether adjustment is towards or away from equilibrium we introduce a simple summary statistic to check for the validity of error correction behaviour. Finally, we consider a *reduced*, parsimonious nonlinear specification by removing statistically insignificant elements of the cubic error correction function. We suggest that, for sensible model reduction, in might be advisable to preserve the sign on the squared error correction term so that parsimonious models can solely incorporate this term whilst maintaining valid error correction behaviour. This modified form of the cubic error correction function will emphasise the dependency of the speed of adjustment upon the degree of equilibrium more than the amount of adjustment being determined by whether agents are above or below equilibrium.

The next section will outline the error-correction model and empirical methods to be employed in its assessment. The third section will present empirical results of the *standard* single equation error-correction models while section 5.4 discusses the results of models which embody *partitioned* asymmetric adjustments towards long run equilibrium. Section 5.5 assesses both the full and parsimonious estimated forms of *cubic* nonlinear adjustment models. The final section draws conclusions.

5.2 *Standard* Error Correction Model Specification and Methodology

The general single equation consumption function to be estimated is:

$$\Delta \ln C_t = \delta_0 + \sum_{i=1}^{L-1}\delta_{1i}\Delta \ln C_{t-i} + \sum_{i=0}^{L-1}\delta_{2i}\Delta \ln Y_{t-i} + \sum_{i=0}^{L-1}\delta_{3i}\Delta\Delta \ln P_{t-i} + \sum_{j=1}^{J-1}\delta_{4j}D_{jt} + \alpha_1 ECM_{t-1} + u_{1t} \quad \textbf{(5.1)}$$

where ECM_{t-1} is the lagged error correction mechanism (ECM) which embodies the favoured long run consumption function identified for each country in Chapter 4. In contrast to Chapter 4 the ECM, defined by **(5.2)**, is normalised upon consumption in the standard manner, that is β_1=1, rather than β_1=-1.[1]

$$ECM_t = \beta_1 \ln C_t - \beta_0 - \beta_2 \ln Y_t - \beta_3 \Delta \ln P_t. \quad \textbf{(5.2)}$$

With this normalisation of the error correction term we expect α_1 in **(5.2)** to be negative and statistically significant for the error correction interpretation to be valid, that is, to ensure that consumption is continually forced towards its long run equilibrium - see, for example, Davidson *et al* 1978 (DHSY hereafter). Since the ECM is only defined over the period 1960-1994, use of its lagged value will cause the estimation period to contract to 1961-1994 (34 observations). Therefore, all the OECD countries' dynamic consumption functions will be estimated over this period.

In Chapter 3 we argued that inflation and the logs of consumption and income were generally integrated of order one across countries. Therefore, they are entered in their I(1) form in **(5.2)** - defining long run behaviour - and in their differenced, stationary, form to represent short-run dynamics in **(5.1)**. That the linear combination of I(1) terms in **(5.2)** cointegrate for each country ensures that each term in equation **(5.1)** is stationary and guards against the problem of spurious, or nonsense, regression which would result in the exaggeration of fit and t-ratios, see Hendry (1980).

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[1] Note that $ECM_t = -Z_t$, where Z_t was used to define the long run consumption function in Chapter 4 in a way which provided directly interpretable coefficients on the explanatory variables.

D_{jt} represents the J-1 dummy variables which may be required to improve model specification and remove misspecification, especially non-normally distributed residuals and structural instability, and are suggested to be capturing unmodelled effects - see Chapter 4 for a discussion of these potential effects.

5.2.1 IV Estimation and Instrument Validity

In Chapter 4 we found evidence that income and inflation are not weakly exogenous with respect to consumption. We therefore use the IV method of estimation to allow for the potential endogeneity of the right hand side variables in the structural equation, **(5.1)**. The treatment of an endogenous variable as if it were exogenous would cause it to be correlated with the model's error term and its OLS parameter estimate to be biased - see, for example, Pindyck and Rubinfeld (1997) ch. 12. One way of overcoming this problem of simultaneous equations bias is to use exogenous variables to instrument the right hand side endogenous variables. Appropriate instruments should be exogenous and highly correlated with the endogenous variable(s) to be instrumented. Lagged values of the endogenous variable(s) to be instrumented, and the lags of variables (causally) related to it, are often considered good instruments and are typically employed. We follow this procedure by using an intercept and the first L lags of consumption growth, income growth and the change in inflation as instruments for current income growth and the change in inflation.[2] Thus, the general unrestricted reduced form system of equations corresponding to the structural equation **(5.1)** is:

$$\Delta\ln C_t = \pi_{10} + \sum_{i=1}^{L}\pi_{11i}\Delta\ln C_{t-i} + \sum_{i=1}^{L}\pi_{12i}\Delta\ln Y_{t-i} + \sum_{i=1}^{L}\pi_{13i}\Delta\Delta\ln P_{t-i} + \sum_{j=1}^{J-1}\pi_{14i}D_{jt} + \pi_{15}ECM_{t-1} + e_{1t} \qquad \textbf{(5.3a)}$$

$$\Delta\ln Y_t = \pi_{20} + \sum_{i=1}^{L}\pi_{21i}\Delta\ln C_{t-i} + \sum_{i=1}^{L}\pi_{22i}\Delta\ln Y_{t-i} + \sum_{i=1}^{L}\pi_{23i}\Delta\Delta\ln P_{t-i} + \sum_{j=1}^{J-1}\pi_{24i}D_{jt} + \pi_{25}ECM_{t-1} + e_{2t} \qquad \textbf{(5.3b)}$$

$$\Delta\Delta\ln P_t = \pi_{30} + \sum_{i=1}^{L}\pi_{31i}\Delta\ln C_{t-i} + \sum_{i=1}^{L}\pi_{32i}\Delta\ln Y_{t-i} + \sum_{i=1}^{L}\pi_{33i}\Delta\Delta\ln P_{t-i} + \sum_{j=1}^{J-1}\pi_{34i}D_{jt} + \pi_{35}ECM_{t-1} + e_{3t} \qquad \textbf{(5.3c)}$$

[2] Any dummy variables incorporated in the structural model are also employed as instruments to help ensure (over) identification. These dummies are assumed to be exogenous.

Equations **(5.3b)** and **(5.3c)** will be used as instruments for income growth and the change in inflation, respectively. The fitted values of these equations, estimated using OLS, are linear combinations of exogenous variables and are, therefore, exogenous themselves and not subject to simultaneity bias. Substituting the fitted values of equations **(5.3b)** and **(5.3c)** into the structural equation, **(5.1)**, yields the restricted reduced form consumption equation, which is not subject to simultaneous equations bias:

$$\Delta \ln C_t = d_0 + \sum_{i=1}^{L-1}\delta_{1i}\Delta \ln C_{t-i} + d_{20}\widehat{\Delta \ln Y_t} + \sum_{i=1}^{L-1} d_{2i}\Delta \ln Y_{t-i} + d_{30}\widehat{\Delta\Delta \ln P_{t-i}} + \sum_{i=0}^{L-1} d_{3i}\Delta\Delta \ln P_{t-i} + \sum_{j=1}^{J-1} d_{4j}D_{jt} \qquad \textbf{(5.4)}$$

$$+ \alpha_1 ECM_{t-1} + u_{2t}$$

where a carat over a variable indicates its fitted value - the instrumented variable.

A necessary (order) condition to be able to identify the unknown parameters of the structural model, **(5.1)**, from the estimable unrestricted reduced form parameters of equations **(5.3a)**, **(5.3b)** and **(5.3c)**, is that the number of instruments used is greater or equal to the number of variables in the restricted reduced form, **(5.4)**. We therefore specify the number of lagged variables (L) in the instrument equations, **(5.3b)** and **(5.3c)**, to be two and the number of lags (L-1) in **(5.4)** to be one. This ensures overidentification with one more instrument than variables in **(5.4)**.[3]

The use of one lagged difference term in **(5.4)** follows the findings of many previous researchers working with annual data and is consistent with the general results of Chapter 4, where only four countries' (Australia, Denmark, Sweden and the UK) favoured VECMs employed a greater augmentation. Given the general likely redundancy of second lagged difference terms, the desire to preserve degrees of freedom and the need to (over) identify equation **(5.4)** we consider one lag as sufficient for the general model. However, in the model reduction process, as variables are omitted from **(5.4)**, second lagged difference terms are considered to see if they improve

[3] Although there are three more lagged variables in the instrument equations relative to **(5.4)** we have to offset this against the two contemporaneous variables which do not feature in the former equations but do in the latter.

model specification.

When the model is overidentified the legitimacy of the instruments can be assessed using Sargan's (1964) test for instrument validity. This test involves regressing the residuals of the restricted reduced form equation, **(5.4)**, against the instruments employed, and testing the statistical significance of the latter. If the instruments are statistically insignificant, they are valid because the assumption that the instruments are independent of equation **(5.4)**'s error is satisfied. It can also be interpreted as indirectly testing whether the restricted reduced form (of the structural model), **(5.4)**, parsimoniously encompasses the unrestricted reduced form, **(5.3a)** - see Doornik and Hendry (1995) p. 320. The test is based upon the auxiliary regression:

$$\hat{u}_{2t} = b_{10} + \sum_{i=1}^{L} b_{11i}\Delta \ln C_{t-i} + \sum_{i=1}^{L} b_{12i}\Delta \ln Y_{t-i} + \sum_{i=1}^{L} b_{13i}\Delta\Delta \ln P_{t-i} + \sum_{j=1}^{J-1} b_{14j}D_{jt} + b_{15}ECM_{t-1} + v_t. \quad \textbf{(5.5)}$$

The lagrange multiplier (LM) test statistic, **(5.6)**, which approximately follows a chi-square distribution under the null of instrument validity with m (which equals the difference between the number of parameters in the unrestricted reduced form, **(5.3a)**, and the restricted reduced form, **(5.4)**) degrees of freedom, is:

$$\text{Sargan}(m) = T.R^2 \sim \chi^2(m) \quad \textbf{(5.6)}$$

where T is the sample size and R^2 is the fit of the auxiliary regression **(5.5)**. The instruments are invalid if the test statistic exceeds the critical value.

5.2.2 (Weak) Exogeneity: Wu-Hausman and Granger Non-Causality Tests

When the variables in a model may be simultaneously determined it is important to consider whether the right hand side variables in a regression should be treated as endogenous or exogenous. That is, whether IV estimation is necessary, or if OLS would produce valid estimates. If there is no evident simultaneity, OLS will be an efficient and consistent parameter estimator while the IV coefficient estimator will be inefficient, if consistent. Conversely, if simultaneity is present, OLS estimates will be biased and inconsistent while IV will be both efficient and consistent - see, for example, Pindyck and Rubinfeld (1997) p. 353.

Engle *et al* (1983) define various degrees of exogeneity (weak, strong and super) to recognise the aim of the econometric analysis and how the parameters of interest are affected. Weak exogeneity of the right hand side variables is necessary to secure valid estimation. This may be illustrated by the useful simplified partition of the data generation process (DGP) into conditional and marginal distributions - see, for example, Charemza and Deadman (1997) for an excellent discussion. In the present context, the conditional distribution may be viewed as our (structural) consumption function (of the *form* of equations **(5.1)** and **(5.4)**) and the marginal distribution as the instrumented (reduced form) equations for income growth and the change in inflation (equations **(5.3b)** and **(5.3c)**, respectively). The parameters of interest are the coefficients to be estimated in the DGP of consumption. Weak exogeneity is secured if the parameters of the DGP can be efficiently estimated (without loss of information) solely from the parameters in the conditional process and if there are no cross restrictions between the parameters in the conditional and marginal processes. Under these conditions, the parameters of interest are efficiently estimated from equation **(5.1)**, by OLS, without reference to the instruments (containing the parameters of the marginal process), equations **(5.3b)** and **(5.3c)**. If weak exogeneity is violated IV estimation is appropriate because the additional information contained in the marginal process is required to estimate the parameters of the DGP.

The Wu-Hausman test for weak exogeneity involves collecting the residuals, $\hat{e}_{2t}$ and $\hat{e}_{3t}$, from the OLS estimates of the reduced form equations of the potentially endogenous variables - in this case equations **(5.3b)** and **(5.3c)**. The structural equation augmented with these residuals, **(5.7)**, is then estimated by OLS.

$$\Delta\ln C_t = c_0 + \sum_{i=1}^{L-1} c_{1i}\Delta\ln C_{t-i} + \sum_{i=0}^{L-1} c_{2i}\Delta\ln Y_{t-i} + \sum_{i=0}^{L-1} c_{3i}\Delta\Delta\ln P_{t-i} + \sum_{j=1}^{J-1} c_{4j}D_{jt} + c_5 ECM_{t-1} + c_5\hat{e}_{2t} + c_6\hat{e}_{3t} + u_{3t} \quad \textbf{(5.7)}$$

One tests the joint hypothesis that c_6=c_7=0, which follows the standard F-distribution under the null. Rejection of the null hypothesis implies that the use of instrumented values for the right hand side variables believed to be endogenous significantly alters the coefficients estimates -

given the error terms are the difference between the actual and instrumented values of these *endogenous* variables. Thus, rejection of the null implies the need to use IV rather than OLS: the right hand side variables cannot be treated as weakly exogenous.

When using error correction models, as we do here, weak exogeneity further requires that there is long run Granger non-causality (LRGNC) in the marginal process.[4] Assessing LRGNC involves testing the statistical significance of the error correction term in the instrument equations, **(5.3b)** and **(5.3c)**. In the income growth equation the hypothesis is $\pi_{25}=0$. Rejection of the null hypothesis implies rejection of LRGNC and, therefore, violates weak exogeneity for income. Similarly one tests the null hypothesis ($\pi_{35}=0$) of weak exogeneity in the change in inflation equation.

LRGNC in the instrumented equations amounts to the error correction mechanism (equation **(5.2)**) appearing in the consumption growth equation but not the income growth and change in inflation equations (marginal processes). If LRGNC is violated, the error correction mechanism enters all the equations simultaneously, making it impossible to obtain inference about each equation separately. That is, cross restrictions between the parameters of the marginal and conditional processes arise (the parameters of the two processes are not variation free). With the parameters of the conditional and marginal processes being interlinked one cannot obtain efficient estimates of the parameters of interest solely from the conditional distribution. However, if the error correction term does not enter the marginal process, there are no cross restrictions between the coefficients and separate estimation of the conditional process is valid - weak exogeneity is secured.

Short run Granger non-causality (SRGNC) is required, *in addition* to weak exogeneity, to secure strong exogeneity. SRGNC, within the context of the exogeneity of income growth and the

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[4] In the current context we use the term long run Granger non causality to mean error correction non causality (see Holland and Scott 1998 p. 1082) because we consider the statistical significance of the impact of the whole error correction term on the dependent variable. The term long run Granger non causality can also be used to consider the statistical significance of a *specific* element of the error correction term upon the dependent variable - see Fraser and Paton (1999) for an implicit example of such an application.

change in inflation, involves testing whether the lags on the growth rate of consumption are statistically significant in the instrumented equations **(5.3b)** and **(5.3c)**. For example, for income (equation **(5.3b)**) one tests the joint hypothesis $\pi_{211}=\pi_{212}=0$. Rejection of the null suggests rejection of SRGNC which, in this case, implies that income growth is not strongly exogenous with respect to consumption growth. Strong exogeneity is of prime interest when the model is to be used for forecasting purposes and implies that the present behaviour of our *exogenous* variables, $\Delta \ln Y_t$ and $\Delta\Delta \ln P_t$, are unaffected by the past behaviour of our endogenous variable, $\Delta \ln C_t$. That is, $\Delta \ln Y_t$ and $\Delta\Delta \ln P_t$ are not better predicted using lagged consumption growth than without it. Although our current aim is to obtain inference from our model rather than use it for forecasting, both weak and strong exogeneity tests will be conducted to provide information on the degree of exogeneity of the variables in our model.[5]

5.2.3 Testing the Robustness of the Estimated Long Run Consumption Functions

Although our application of the Johansen procedure in Chapter 4 estimated each country's long run consumption function whilst simultaneously accounting for short run dynamics, it remains a possibility that different estimates of these equilibriums may arise in the single equation error correction model. For example, Malley and Moutos (1996) cite the finding of Campbell and Perron (1991) that the Johansen estimates are sensitive to misspecification in any of the equations in the VECM.[6] To assess the robustness of the Johansen estimates embodied in the error correction term we test the joint statistical significance of $\ln Y_{t-1}$ and $\Delta \ln P_{t-1}$ when added to the favoured form of **(5.4)** using a standard Wald (F) test with the IV (OLS) estimator. If the null is rejected, these long run effects are statistically significant, suggesting that the Johansen estimates of the long run consumption function may not be robust.

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[5] Strong exogeneity combined with parameter stability yields super exogeneity. This property is required for policy analysis. Some information on super exogeneity may be gleaned from the forecast tests we conduct on our estimated models.

[6] We note that we were careful in Chapter 4 to ensure that the VECM from which the Johansen estimates were obtained were free from evident misspecification.

5.3 Empirical Results for *Standard* Error Correction Models

This section summarises the results of the OECD countries' favoured error correction models with *standard symmetric* adjustment towards long run consumption. The favoured error correction model for each country is reported in Table 5.1. In this Table "Country" denotes the country to which the results relate, "IV/OLS" indicates the estimation method used, "Drop Instr" refers to the variables *excluded* from the instrument set specified by equations **(5.3a)** and **(5.3b)**. "Dummies" gives the dummy variables included in each model, with the specified dates indicating the periods which take on a unit value (all other periods are zero). Estimated coefficients are given in the rows of the variables to which they relate with t-ratios specified directly below them in brackets. "$\Sigma\Delta \ln Y_{t-i}$" gives the sum of the coefficients on the income growth terms and "$\Sigma\Delta\Delta \ln P_{t-i}$" the sum of parameters on the change in inflation terms. "Adjusted R^2" is the coefficient of determination adjusted for degrees of freedom, "s" is the unbiased estimate of the regression's standard error and "DW" is the Durbin-Watson statistic for first order autocorrelation. "Wu-Haus(n)" denotes the Wu-Hausman test for weak exogeneity for a model including n endogenous variables (given in normal brackets) with the probability *up to* which the test statistic is statistically insignificant [given in squared brackets]. "LRGNC($\Delta \ln Y_t$)" denotes the probability for the tests for long run Granger non-causality (weak exogeneity) with respect to the income growth equation. "SRGNC($\Delta \ln Y_t$)" specifies the probability of the test for short run Granger non-causality, with the number of lagged consumption terms tested given in (normal brackets). "Sargan(m)" refers to the probability value of Sargan's test for instrument validity with degrees of freedom (m) given in normal brackets. "$\ln Y_{t-1}, \Delta \ln P_{t-1}$" gives the probability up to which lagged income and inflation exhibit joint statistical insignificance, which is a Wald test with two degrees of freedom - specified in (normal brackets). "SC:$\chi^2(2)$" is a χ^2-distributed LM test for second order residual autocorrelation, "ARCH:F(1)" is an F-distributed test for first order autoregressive conditional heteroscedasticity, "N:$\chi^2(2)$" is the Jarque and Bera test for skewness and excess kurtosis which follows a χ^2 distribution with two degrees of freedom. "H:F(k-1)" is White's (1980) test for heteroscedasticity which is F-distributed with k-1 degrees of freedom. "H/FF:F" is a heteroscedasticity/functional form F-test - this is only reported if there are sufficient degrees of freedom. "FOR:χ^2(H)" is a forecast test for numerical parameter constancy where the sample is split in period T-H/T-H+1: in this case H=1 (1993/94), 2

TABLE 5.1: Standard Error Correction Models

Country	Australia	Austria	Belgium	Canada	Denmark	Finland	France	Germany	Greece	Iceland
IV/OLS	IV	IV	IV	IV	IV	IV	IV	IV	IV	IV
Drop Instr	NONE	NONE	NONE	NONE	NONE	NONE	NONE	$\Delta\Delta\ln P_{t-2}$	NONE	NONE
Dummies	6775	77;78;83	83	82;91	86	6988;71;	7493	91	80	75
						9192				
Intercept	0.002	-0.006	0.003	-0.008	0.058	0.107	-0.059	-0.003	0.008	0.006
	(0.052)	(-1.444)	(0.404)	(-1.163)	(1.304)	(3.700)	(-2.823)	(-0.600)	(2.360)	(1.131)
$\Delta\ln Y_t$	0.523	0.399	0.449	0.770	0.864	0.882	0.391	0.687	0.310	0.588
	(4.536)	(2.747)	(3.667)	(4.446)	(5.147)	(5.936)	(2.236)	(4.555)	(3.181)	(5.571)
$\Delta\ln Y_{t-1}$								0.119		
								(-1.783)		
$\Delta\ln Y_{t-2}$	-0.175	0.198					-0.168			-0.179
	(-4.006)	(2.633)					(-2.465)			(-2.693)
$\Delta\Delta\ln P_{t-1}$			-0.284							
			(-3.014)							
$\Delta\ln C_{t-1}$					0.713					
					(4.240)					
ECM_{t-1} (U/R)	-0.133 (R)	-0.315 (R)	-0.097 (R)	-0.423 (R)	-0.075 (U)	-0.497 (U)	-0.156 (U)	-0.125 (R)	-0.254 (R)	-0.386 (R)
	(-3.052)	(-3.853)	(-2.083)	(-3.908)	(-1.525)	(-3.208)	(-3.237)	(-2.144)	(-3.977)	(-4.798)
$\Sigma\Delta\ln Y_{t-i}$	0.348	0.597	0.449	0.770	0.864	0.882	0.223	0.806	0.310	0.409
$\Sigma\Delta\Delta\ln P_{t-i}$	0.000	0.000	-0.284	0.000	0.000	0.000	0.000	0.000	0.000	0.000
Adjusted R^2	0.806	0.868	0.691	0.824	0.549	0.865	0.832	0.874	0.828	0.858
s	0.0059	0.0065	0.0106	0.0109	0.0197	0.0122	0.0065	0.0087	0.0109	0.0236
DW	2.370	1.780	1.740	1.340	2.140	2.010	2.110	1.900	2.300	2.390
Wu-Haus(n)	[0.636](1)	[0.527](1)	[0.156](1)	[0.121](1)	[0.092](1)	[0.565](1)	[0.892](1)	[0.327](1)	[0.323](1)	[0.510](1)
LRGNC($\Delta\ln Y_t$)	[0.312]	[0.960]	[0.173]	[0.535]	[0.224]	[0.238]	[0.052]	[0.382]	[0.028]	[0.583]
SRGNC($\Delta\ln Y_t$)	[0.224](2)	[0.276](2)	[0.019](2)	[0.908](2)	[0.060](2)	[0.525](2)	[0.521](2)	[0.646](2)	[0.838](2)	[0.545](2)
Sargan(m)	[0.660](4)	[0.862](4)	[0.898](4)	[0.321](5)	[0.105](4)	[0.615](5)	[0.557](4)	[0.795](3)	[0.365](5)	[0.089](4)
$\ln Y_{t-1}\Delta\ln P_{t-1}$	[0.945](2)	[0.874](2)	[0.636](2)	[0.260](2)	[0.468](2)	[0.867](2)	[0.644](2)	[0.468](2)	[0.158](2)	[0.853](2)
SC:$\chi^2(2)$	[0.626]	[0.110]	[0.911]	[0.081]	[0.332]	[0.704]	[0.532]	[0.777]	[0.567]	[0.074]
ARCH:F(1)	[0.094]	[0.623]	[0.958]	[0.956]	[0.389]	[0.294]	[0.772]	[0.289]	[0.355]	[0.372]
N:$\chi^2(2)$	[0.271]	[0.802]	[0.633]	[0.805]	[0.356]	[0.915]	[0.167]	[0.646]	[0.238]	[0.572]
H:F(k-1)	[0.195]	[0.298]	[0.816]	[0.939]	[0.959]	[0.200]	[0.598]	[0.139]	[0.977]	[0.617]
H/FF:F	[0.429]	N/A	[0.852]	[0.818]	[0.961]	[0.402]	[0.855]	[0.129]	[0.992]	[0.623]
FOR:$\chi^2(1)$	[0.451]	[0.536]	[0.879]	[0.191]	[0.913]	[0.079]	[0.288]	[0.798]	[0.469]	[0.943]
FOR:$\chi^2(2)$	[0.682]	[0.533]	[0.083]	[0.373]	[0.606]	[0.082]	N/A	[0.829]	[0.329]	[0.906]
FOR:$\chi^2(5)$	[0.822]	[0.800]	[0.229]	N/A	[0.785]	N/A	N/A	N/A	[0.669]	[0.983]
FOR:$\chi^2(9)$	[0.709]	[0.983]	[0.613]	N/A	N/A	N/A	N/A	N/A	[0.929]	[0.113]
FOR:$\chi^2(15)$	[0.219]	N/A	N/A	N/A	N/A	N/A	N/A	N/A	N/A	[0.123]

Table 5.1 continued

Country	Ireland	Italy	Japan	Netherlnd	Norway	Spain	Sweden	Switz.	UK	USA
IV/OLS	IV	OLS	IV	IV	OLS	IV	IV	IV	IV	IV
Drop Instr	NONE	NONE	NONE	NONE	NONE	NONE	NONE	NONE	NONE	NONE
Dummies	7375;82	7679;93	61	648182;	69;78;	87;93	9092	6272;	868788;	NONE
				7791	8586			666879	9192	
Intercept	0.029	-0.004	-0.003	-0.001	-0.009	0.104	0.004	0.001	0.046	-0.003
	(5.950)	(-0.584)	(-0.934)	(-0.325)	(-2.029)	(5.033)	(0.879)	(0.252)	(3.962)	(-0.750)
$\Delta \ln Y_t$	0.880		0.681	0.776		0.777	0.646	0.997	0.793	0.828
	(4.357)		(8.572)	(7.451)		(12.484)	(2.764)	(8.128)	(6.911)	(3.925)
$\Delta \ln Y_{t-1}$		0.205			0.314			-0.505		
		(2.132)			(2.195)			(-4.414)		
$\Delta \ln Y_{t-2}$										
$\Delta\Delta \ln P_t$			-0.417						-0.226	
			(-6.126)						(-2.968)	
$\Delta\Delta \ln P_{t-1}$								-0.141		-0.237
								(-2.075)		(-1.935)
$\Delta\Delta \ln P_{t-2}$		-0.345		-0.168						
		(-4.364)		(-2.273)						
ECM_{t-1} (U/R)	-0.521 (R)	-0.035 (R)	-0.289 (R)	-0.137 (R)	-0.373 (R)	-0.433 (U)	0.042 (R)	0.009 (R)	-0.267 (U)	-0.238 (R)
	(-2.916)	(-4.095)	(-5.256)	(-2.580)	(-5.115)	(-4.819)	(0.641)	(0.642)	(-2.643)	(-1.949)
$\Sigma \Delta \ln Y_{t-i}$	0.880	0.205	0.681	0.776	0.314	0.777	0.646	0.492	0.793	0.828
$\Sigma \Delta\Delta \ln P_{t-i}$	0.000	-0.345	-0.417	-0.168	0.000	0.000	0.000	-0.141	-0.226	-0.237
Adjusted R^2	0.787	0.776	0.942	0.915	0.748	0.925	0.512	0.887	0.872	0.728
s	0.0143	0.0113	0.0069	0.0064	0.0135	0.0085	0.0154	0.0060	0.0083	0.0089
DW	1.730	2.260	2.390	1.670	2.020	1.650	2.260	2.090	2.390	1.960
Wu-Haus(m)	[0.853](1)	N/A	[0.902](2)	[0.204](1)	N/A	[0.651](1)	[0.738](1)	[0.057](1)	[0.189](2)	[0.447](1)
LRGNC($\Delta \ln Y_t$)	[0.001]	N/A	[0.410]	[0.062]	N/A	[0.131]	[0.002]	[0.213]	[0.027]	[0.224]
SRGNC($\Delta \ln Y_t$)	[0.366](2)	N/A	[0.606](2)	[0.046](2)	N/A	[0.096](2)	[0.034](2)	[0.371](2)	[0.804](2)	[0.084](2)
Sargan(m)	[0.422](5)	N/A	[0.612](4)	[0.950](4)	N/A	[0.547](5)	[0.636](5)	[0.336](3)	[0.902](4)	[0.529](4)
$\ln Y_{t-1} \Delta \ln P_{t-1}$	[0.053](2)	[0.881](2)	[0.768](2)	[0.711](2)	[0.547](2)	[0.367](2)	[0.584](2)	[0.536](2)	[0.397](2)	[0.388](2)
SC:$\chi^2(2)$	[0.277]	[0.075](F)	[0.524]	[0.732]	[0.639](F)	[0.057]	[0.630]	[0.788]	[0.362]	[0.609]
ARCH:F(1)	[0.732]	[0.500]	[0.579]	[0.683]	[0.691]	[0.361]	[0.590]	[0.334]	[0.214]	[0.327]
N:$\chi^2(2)$	[0.387]	[0.751]	[0.808]	[0.725]	[0.863]	[0.635]	[0.918]	[0.865]	[0.548]	[0.339]
H:F(k-1)	[0.384]	[0.969]	[0.489]	[0.378]	[0.603]	[0.663]	[0.462]	[0.544]	[0.987]	[0.380]
H/FF:F	[0.509]	[0.905]	[0.740]	[0.772]	[0.694]	[0.777]	[0.657]	N/A	[0.998]	[0.606]
FOR:$\chi^2(1)$	[0.307]	[0.467]	[0.943]	[0.317]	[0.069]	[0.809]	[0.140]	[0.521]	[0.400]	[0.460]
FOR:$\chi^2(2)$	[0.389]	N/A	[0.609]	[0.063]	[0.185]	N/A	[0.113]	[0.738]	[0.659]	[0.571]
FOR:$\chi^2(5)$	[0.727]	N/A	[0.444]	N/A	[0.068]	N/A	N/A	[0.655]	N/A	[0.925]
FOR:$\chi^2(9)$	[0.060]	N/A	[0.386]	N/A	N/A	N/A	N/A	[0.649]	N/A	[0.912]
FOR:$\chi^2(15)$	N/A	N/A	[0.532]	N/A	N/A	N/A	N/A	[0.134]	N/A	[0.970]

Table 5.1 notes. See main text for details of the reported statistics.

(1992/92), 5 (1989/90), 9 (1985/86) and 15 (1979/80).[7] Probabilities rather than test statistics are reported for all of these misspecification tests. All statistics/probabilities were produced using PcGive 8.0 except the adjusted R^2 and Wald/F tests for SRGNC and omitted long run effects, which were obtained with Microfit 3.22.

For no country is there evidence of model misspecification (serial correlation, ARCH, non-normality, heteroscedasticity, non-linear functional form or non-constancy of parameters) at the 5% level - all the reported probabilities exceed 0.05. The Wu-Hausman test statistic is statistically insignificant for all eighteen countries where contemporaneous variables enter the favoured model, however, LRGNC, with respect to income, is rejected for four countries (Greece, Ireland, Sweden and the UK) suggesting violation of weak exogeneity in these cases.[8] We note that our application of this test may be biased towards the non-rejection of weak exogeneity because we do not pursue model reduction (the general IV approach) in the instrumented equations (the reaction functions). Thus some variables may appear to be statistically insignificant when they would be significant in a parsimonious reaction function which, in the case of the error correction terms, would lead to the erroneous non-rejection of weak exogeneity. Indeed, the results of Chapter 4 suggested that weak exogeneity of income and/or inflation is violated for sixteen of the countries' VECMs. This suggests a far more widespread violation of weak exogeneity compared to the LRGNC tests conducted here.[9] We therefore use IV rather than OLS when contemporaneous variables enter our models.[10] OLS is only employed for the two countries (Italy and Norway) where no current dated variables are

[7] All forecast tests cannot be conducted for every model due to the presence of dummy variables, however, as many as possible are produced.

[8] LRGNC with respect to inflation cannot be rejected for the two countries, Japan and the UK, where inflation enters the models contemporaneously, with probabilities of 0.511 and 0.968, respectively.

[9] Our inclusion of error correction terms in all of the instrumented equations allows for feedbacks of long run information into all of the contemporaneous variables' equations. This is incompatible with weak exogeneity and LRGNC - see Steel (1987) p. 1444.

[10] We prefer to suffer inefficient estimation by using IV than risk bias and inconsistent estimation from using OLS based upon false *acceptance* of weak exogeneity.

maintained in the favoured model. According to the Sargan test the instruments used in IV estimation are valid for all countries.[11] SRGNC, with respect to income, is rejected for Belgium, the Netherlands and Sweden, indicating violation of strong exogeneity for these three countries suggesting that these models would not provide good forecasts.[12] Finally, the long run effects of $\ln Y_{t-1}$ and $\Delta\ln P_{t-1}$ cannot be entered with joint statistical significance (at the 5% level) for any country which provides some evidence that the long run consumption functions embodied in the error correction terms are robustly estimated. We therefore present the error correction models reported in Table 5.1 as providing valid inference.

For sixteen of the twenty countries the adjustment coefficients on error correction terms are negative and statistically significant at the 5% level suggesting coherence with the long run relations estimated in Chapter 4. For two of the four countries where this condition is not met, Denmark and the USA, the error correction terms are negatively signed and are almost statistically significant, with t-ratios equal to -1.525 and -1.949, respectively. We argue that the USA's error correction model is consistent with valid error correction behaviour because the adjustment coefficient is so close to being statistically significant and that with the use of a 5% level of significance we would expect one incorrect inference from the twenty countries' results. For Denmark there is some doubt over the model's coherence with equilibrium consumption, which is consistent with the reservations expressed in Chapter 4 regarding the plausibility of this country's estimated long run consumption function.[13] However, for Sweden and Switzerland the estimated coefficient on the error correction term is positive and statistically insignificant suggesting no adjustment towards long run equilibrium. This is consistent with the severe

[11] $\Delta\Delta\ln P_{t-2}$ was excluded from the instrument set for Germany to secure instrument validity.

[12] SRGNC with respect to inflation is not rejected for Japan and the UK - the associated probabilities are 0.191 and 0.709, respectively.

[13] We note that if we were to exclude the intercept from Denmark's model, the t-ratio on the adjustment coefficient of the error correction term becomes -1.982, which is very close to being statistically significant, and more favourable towards finding valid error correction behaviour. However, we resist removing the positive intercept, despite it being statistically insignificant (its t-ratio is 1.304), to obtain these more desirable results, because removing such an intercept causes parameter bias.

reservations expressed in Chapter 4 regarding these countries' cointegrating vectors. For both countries we interpret their models as only usefully characterising short run (and not long run) consumer behaviour.[14]

Excepting the intercepts, which are always included to avoid parameter bias, and two minor anomalies regarding lagged income growth in the German equation (t-ratio is -1.783) and the lagged change in inflation in the USA's equation (t-ratio is -1.935), all other coefficients are statistically significant. Since these two terms are so nearly significant and remove misspecification and/or improve specification, we argue that these minor anomalies may be ignored. For all twenty countries the overall impacts of income and inflation (when present) are positive and negative, respectively, which is consistent with theoretical prior beliefs. Thus, the models are presented as economically sensible, statistically sound, parsimonious consumption

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[14] Exclusion of the error correction term for the Swedish and Swiss models yields the estimated short run consumption functions reported below (all statistics are as defined for Table 5.1). Both equations are estimated by IV with the error correction term excluded from the instruments:

Sweden

$$\widehat{\Delta \ln C_t} = \underset{\mathbf{(1.639)}}{0.006} + \underset{(4.202)}{0.760\Delta \ln Y_t}$$

$\bar{R}^2 = 0.509$, $s = 0.0155$, DW = 2.15, Sargan(m) = [0.589](5); SC:$\chi^2(2)$ = [0.901], ARCH:F(1) = [0.571], N:$\chi^2(2)$ = [0.840], H:F(k-1) = [0.546], H/FF:F = [0.719]; FOR:$\chi^2(1)$ = [0.528], FOR:$\chi^2(2)$ = [0.103].

Switzerland

$$\widehat{\Delta \ln C_t} = \underset{\mathbf{(1.724)}}{0.003} + \underset{(9.671)}{1.011\Delta \ln Y_t} - \underset{(-4.851)}{0.513\Delta \ln Y_{t-1}} - \underset{\mathbf{(-1.985)}}{0.133\Delta\Delta \ln P_{t-1}}$$

$\bar{R}^2 = 0.884$, $s = 0.0061$, DW = 2.00, Sargan(m) = [0.463](3), $\Sigma\Delta \ln Y_t = 0.498$; SC:$\chi^2(2)$ = [0.947], ARCH:F(1) = [0.423], N:$\chi^2(2)$ = [0.621], H:F(k-1) = [0.783], H/FF:F = [0.936]; FOR:$\chi^2(1)$ = [0.412], FOR:$\chi^2(2)$ = [0.584], FOR:$\chi^2(5)$ = [0.772], FOR:$\chi^2(9)$ = [0.608], FOR:$\chi^2(15)$ = [0.056].

The inferences from the short run models given above are essentially the same as those reported in Table 5.1. However, the equations given in Table 5.1 exhibit slightly better fit so represent the favoured equations from which inference is discussed in the text. Dummies are not reported.

functions - with some reservations regarding long run coherence for Sweden, Switzerland and, to a lesser extent, Denmark.

The fit of the equations, according to the adjusted R^2, range from 51.2% for Sweden to 94.2% for Japan relative to an average of 80.4%, which indicates that there is very high explanatory power for the majority of countries' equations. Dummy variables are employed in some countries' equations and we argue that they help capture omitted factors such as wealth, deregulation, demographic and income uncertainty effects. For example, in the UK equation the dummy variable for 1986, 1987 and 1988 captures the exceptional consumer boom of the mid to late 1980s while the dummy variable for 1991 and 1992 captures the subsequent deep slump of the early 1990s.[15] Other events also require interventions, for example, a dummy for 1991 is used to capture German reunification.

The speed of adjustment toward equilibrium, indicated by the coefficient on the error correction term, ranges from a very slow -0.035 (3.5% per annum) for Italy to a moderately fast -0.521 (52.1%% per annum) for Ireland. The average for the eighteen countries where there is coherence with the long run equilibrium is -0.264 (26.4% per annum). The overall short run response of consumption to income (the sum of the income coefficients denoted by $\Sigma\Delta \ln Y_{t-i}$ in Table 5.1) also varies widely across countries. The overall impact is always positive, which is consistent with economic prior beliefs, is always less than the long run income elasticity (see Chapter 4) and less than unity. The sum of the coefficients on income varies from 0.205 for Italy to 0.882 for Finland compared to an average value of 0.603 for all twenty countries. The overall impact of (the change in) inflation upon consumption is given by the sum of coefficients on the inflation terms in each country's equation (denoted by $\Sigma\Delta\Delta \ln P_{t-i}$ in Table 5.1). For only seven (Belgium, Italy, Japan, the Netherlands, Switzerland, the UK and the USA) of the twenty countries are there short run inflation effects. In all seven cases the overall impact of (the change in) inflation on consumption is negative, which conforms with our prior belief. The overall impacts range from -0.141 for Switzerland to -0.417 for Japan compared to an average value (for

[15] Similarly, dummies for the 1980s and 1990s, which also likely capture wealth/deregulation effects, are employed for Denmark, Finland, Norway and Sweden.

the seven countries where inflation effects are evident) of -0.267. That some countries' have zero and others non-zero inflation effects suggests that these elasticities are quite different from country to country. Overall, there is a clear heterogeneity of the speed of adjustment and short run income and inflation elasticities across countries. We attempt to explain this cross country variation in Chapter 7.

We note that short run inflation effects feature in only seven countries' error correction models, whereas short run income effects are apparent in all twenty models. In Chapter 4 we found that inflation was negative and statistically significant in eight of the twenty countries' long run consumption functions. Indeed, we find that there are statistically significant short and/or long run inflation effects for only twelve of the twenty countries - inflation enters with statistical significance in both short and long run components of the model for Japan, the UK and the USA. That inflation does not enter in any part of eight of the twenty countries' models suggests that it may not be a fundamental determinant of consumer behaviour for many countries. Further, inflation being statistically significant in the short and long run aspects of the consumption function for both the UK and the USA is interesting because it is these two countries which have been the subject of the majority of analysis on consumer behaviour and thus provided much of the evidence supporting the role of inflation. However, the international evidence presented here indicates that the importance of inflation for these two countries does not necessarily generalise to other OECD countries. One might argue that inflation would have less impact upon consumption than we report if asset variables were incorporated in our consumption functions. For example, when well defined wealth effects are incorporated in consumption functions used in the major UK macroeconomic models, inflation becomes redundant for the majority of specifications (see Church *et al* 1994). Lattimore (1994) demonstrates a similar redundancy of inflation effects for an Australian consumption function with well defined asset variables.

5.4 *Partitioned* Asymmetric Adjustment to Equilibrium

The error correction model considered so far assumes the standard symmetric adjustment towards the long run consumption function. In this section we consider the *partitioned* form of non-symmetric adjustment suggested by Granger and Lee (1989) where the ECM_{t-1} term is separated into positive, $ECM+_{t-1}$, and negative, $ECM-_{t-1}$, components. When consumption is above (below) its equilibrium level the error correction term solely contains positive (negative) values. This partition facilitates the characterisation of agents as adjusting towards their long run level at a different speed when consumption is above equilibrium compared to when it is below equilibrium. Previous applications of such *partitioned* asymmetric error correction models include Granger and Lee's (1989) analysis of production/inventories in the USA, Arden *et al*'s (1997) investigation of manufacturing prices in the UK and Milas's (1999) study of wage rigidities in Greece. Both Arden *et al* (1997) and Milas (1999) find evidence of asymmetric adjustment. We are not aware of any previous examination of *partitioned* asymmetric error correction behaviour for OECD countries' consumer behaviour, although Holly and Stannett (1995) find some evidence for asymmetries in UK consumption using time-series methods.[16] They suggest further work employing structural models of consumer behaviour would be desirable. We do exactly this by considering asymmetric adjustments towards equilibrium using structural consumption functions for twenty OECD countries.

Holly and Stannett (1995) "do not offer any particular explanations for asymmetric behaviour in UK consumption except to note that there have been particular difficulties in explaining the pattern of expenditure over the last six years with a failure to predict the strength of the boom and the depth of the subsequent recession. Many commentators have drawn attention to the way in which financial liberalization has affected the ability of UK household to go into debt and this, coupled with a close relationship between the housing market and consumption and savings, may create the opportunity for households to adjust differently depending on whether the housing prices are rising or falling." (Holly and Stannett 1995, p. 767). Carruth and Dickerson (1997) suggests that asymmetries may arise if consumers behave differently in the manner in which they release and spend equity in housing market booms compared to housing market recessions. "When times are good, lenders may behave differently than when times are bad and this

[16] Carruth and Dickerson (1997) find evidence of asymetric adjustment to equilibrium in UK consumption by adding a dummy variable, which is unity when consumption is above equilibrium and zero otherwise, multiplied by the error correction term to an analogue of the DHSY model. They find evidence of faster adjustment to equilibrium when consumption is above its long run value (in good times).

generates an asymmetry in the housing market and, potentially, in consumer spending." (Carruth and Dickerson 1997, p. 4). Thus, one might expect that consumers are more able to adjust towards their desired level of consumption when above equilibrium than when below it because of, for example, the tightness of liquidity constraints. Alternatively, one might argue that agents are less inclined to moderate consumption when above their long run level than to adjust it upwards when below equilibrium because of rigidities in lowering living standards. This could be justified according to the Relative Income Hypothesis of Duesenberry (1949) and Brown (1952).[17] Under the former hypothesis adjustment is faster when consumers are above equilibrium whilst the alternative hypothesis suggests that agents adjust more quickly when below their long run target. We explore whether such asymmetries prevail for OECD countries and, if they do, whether adjustment is faster when above or below equilibrium.

To detect the presence and direction of such asymmetric behaviour we apply two modifications to the error correction terms employed so far. The first modification is to subtract the mean from the Johansen error correction term.[18] The need for this alteration arises because the Johansen method does not ensure that the error correction term is of zero mean, even when the intercept is restricted into the cointegration space.[19] Indeed, the Johansen procedure may yield an error correction series which is completely positive or completely negative. Under the assumption that consumption is being continually forced to equilibrium and is, on average, in equilibrium, the error correction term should be centred around zero (since being in equilibrium implies no

[17] Bowman, Minehart and Rabin (1999) argue that in the face of sufficient income uncertainty, consumers resist lowering consumption in response to bad news about future income more than they resist increasing consumption in response to good news. This provides another potential rationalisation for a faster adjustment to desired consumption when below equilibrium than when above it.

[18] The mean of each country's error correction term is calculated over the period 1960-1993 which, given it is lagged, defines when it affects the regression, which is estimated over the period 1961-1994.

[19] The role of the intercept in the Johansen VECM, whether restricted or unrestricted, is to ensure the equalisation of means of both sides of each of the VAR's equations. It provides no mechanism for ensuring that the error correction mechanism is of zero mean. This contrasts with the use of the Engle and Granger (1987) method where the inclusion of the intercept secures a zero mean error correction term.

error). Peel and Davidson (1998) suggest defining the intercept in the cointegrating vector as the difference between the means of the two sides of the equilibrium relation to yield a zero mean error correction term.[20] The second modification is required to partition the error correction term into positive and negative terms. The positive (negative) error correction term is constructed to equal the actual value of the (symmetric) error correction term when positive (negative) and to be zero otherwise. Thus, $ECM_{t-1} = ECM+_{t-1} + ECM-_{t-1}$. The *partitioned* asymmetric error correction model takes the general form:[21]

$$\Delta lnC_t = f_0 + f_{11}\Delta lnC_{t-1} + f_{20}\Delta ln\hat{Y}_t + \Sigma f_{21}\Delta lnY_{t-i} + f_{30}\Delta\Delta ln\hat{P}_t + \Sigma f_{3i}\Delta\Delta lnP_{t-1} + \Sigma f_{4j}D_{jt} \qquad \textbf{(5.8)}$$

$$+ a_{1+}ECM+_{t-1} + a_{1-}ECM-_{t-1} + u_{3t}.$$

The null hypothesis of symmetric adjustment is:

$$H_0: a_{1+} = a_{1-}, \qquad \textbf{(5.9)}$$

which we assess using a standard Wald test. Rejection of the null implies asymmetric adjustment. Examination of the relative magnitudes of the estimated coefficients will indicate whether adjustment is quicker when agents are above or below equilibrium. However, Cook, Holly and Turner (1999) demonstrate that this Wald test possesses low power in rejecting the null of symmetric adjustment, especially when using small samples. Because of the Wald test's low probability of uncovering asymmetries when they exist alternative means of assessing asymmetries are also employed.

Some evidence for asymmetries can be gleaned by comparing the magnitudes of the estimated coefficients on the *positive* and *negative* error correction terms. *Substantial* differences in these magnitudes would suggest asymmetry, where an adjustment coefficient which is twice the

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[20] Such a constant adjustment appears to be automatically applied when calculating (unrestricted intercept) Johansen cointegrating vectors in the computer package Eviews 2.0.

[21] Both positive and negative error correction terms are used as instruments in the estimation of this model.

magnitude of the other will certainly be regarded as substantial, although other sizeable differences, if smaller, are not discounted. Further, if the *partitioned* asymmetric adjustment model's fit improves relative to that of the standard symmetric adjustment model, this will also be regarded as tentative evidence in favour of (partitioned) asymmetries.

The results of the favoured *partitioned* asymmetric error correction models are reported in Table 5.2. All reported statistics are as for Table 5.1 except "$ECM+_{t-1}$" and "$ECM-_{t-1}$" denote the variables for the positive and negative error correction terms. "Symmetry(1)" is the F-version of a Wald test for the single symmetry restriction (produced by PcGive 8.0). We find some evidence of non-constant parameters, at the 5% level, for Belgium, Iceland and Switzerland. Serial correlation is evident for Canada at the 5% level but not the 1% level. We present the results as providing generally valid inference if with some reservations regarding Belgium, Iceland and Switzerland.

The symmetry restriction cannot be rejected at the 5%, 10% and, indeed, 15% levels of significance for any country, suggesting no evidence of asymmetric adjustment to the long run equilibrium. However, because this test has low power we compare the coefficients of the positive and negative error correction terms. The coefficient on one of the error correction terms is twice the magnitude of the other for five countries. The parameters on positive and negative error correction terms where this is the case are, -0.101 and -0.608 for Canada, -0.169 and -0.006 for Denmark, -0.241 and -1.00 for France, -0.044 and -0.219 for the Netherlands and -0.434 and -0.211 for Norway. *Large* differences in the adjustment coefficients for the positive and negative error correction terms are also observed for the following six countries: -0.221 and -0.382 for Austria, -0.611 and -0.364 for Finland, -0.441 and -0.325 for Iceland, -0.332 and -0.629 for Ireland, -0.536 and -0.332 for Spain and -0.154 and -0.308 for the USA. These differences in estimated parameters suggest that asymmetries may exist. However, because the error correction terms are generally poorly determined, *both* positive and negative terms only enter with statistical significance in the equations for Iceland and Japan, simple comparison of the estimated parameters may be misleading. The overall improvement of fit of asymmetric models (Table 5.2) relative to symmetric formulations (Table 5.1) may well provide a more discerning criteria. The fit of the equations for Austria, France and the Netherlands improve with the removal of the

TABLE 5.2: *Partitioned* Asymmetric Error Correction Models

Country	Australia	Austria	Belgium	Canada	Denmark	Finland	France	Germany	Greece	Iceland
IV/OLS	IV	IV	IV	IV	IV	IV	IV	IV	IV	IV
Drop Instr	NONE	NONE	NONE	NONE	NONE	NONE	NONE	$\Delta\Delta \ln P_{t-2}$	NONE	NONE
Dummies	6775	77;78;83	83	82;91	86	6988;71;	7493	91	80	75
						9192				
Intercept	0.013	0.007	0.014	0.001	-0.009	0.008	0.025	0.006	0.022	0.022
	(4.601)	(1.035)	(3.077)	(0.080)	(-1.392)	(1.212)	(4.735)	(1.506)	(3.965)	(2.595)
$\Delta \ln Y_t$	0.531	0.409	0.451	0.808	0.895	0.864	0.396	0.686	0.305	0.588
	(4.064)	(2.820)	(3.601)	(4.215)	(4.467)	(5.553)	(2.302)	(4.531)	(3.025)	(5.363)
$\Delta \ln Y_{t-1}$								0.118		
								(1.724)		
$\Delta \ln Y_{t-2}$	-0.176	0.217					-0.186			-0.184
	(-3.915)	(2.780)					(-3.129)			(-2.654)
$\Delta\Delta \ln P_t$										
$\Delta\Delta \ln P_{t-1}$			-0.283							
			(-2.953)							
$\Delta \ln C_{t-1}$					0.753					
					(3.923)					
$ECM+_{t-1}$	-0.147 (R)	-0.221 (R)	-0.106 (R)	-0.101 (R)	-0.169 (U)	-0.611 (U)	-0.241 (U)	-0.120 (R)	-0.225 (R)	-0.441 (R)
	(-1.621)	(-1.627)	(-1.217)	(-0.334)	(-1.331)	(-2.648)	(-3.142)	(-1.205)	(-1.974)	(-3.098)
$ECM-_{t-1}$	-0.118 (R)	-0.382 (R)	-0.087 (R)	-0.608 (R)	-0.006 (U)	-0.364 (U)	-0.100 (U)	-0.132 (R)	-0.287 (R)	-0.325 (R)
	(-1.183)	(-3.380)	(-0.940)	(-3.019)	(-0.067)	(-1.225)	(-1.574)	(-0.900)	(-2.325)	(-2.291)
$\Sigma\Delta \ln Y_{t-i}$	0.355	0.626	0.451	0.808	0.895	0.864	0.210	0.804	0.305	0.404
$\Sigma\Delta\Delta \ln P_{t-i}$	0.000	0.000	-0.283	0.000	0.000	0.000	0.000	0.000	0.000	0.000
Adjusted R^2	0.796	0.869	0.681	0.814	0.530	0.864	0.837	0.795	0.821	0.854
s	0.0060	0.0065	0.0108	0.0112	0.0201	0.0122	0.0064	0.0089	0.0110	0.0239
DW	2.35	1.86	1.75	1.37	2.20	2.00	2.16	1.90	2.26	2.41
Sargan(m)	[0.648](4)	[0.802](4)	[0.900](4)	[0.335](5)	[0.105](4)	[0.631](5)	[0.784](4)	[0.795](3)	[0.382](5)	[0.072](4)
Symmetry(1)	[0.868]	[0.397]	[0.901]	[0.266]	[0.393]	[0.571]	[0.176]	[0.955]	[0.756]	[0.621]
SC:χ^2(2)	[0.686]	[0.122]	[0.936]	[0.038]	[0.386]	[0.696]	[0.587]	[0.772]	[0.632]	[0.110]
ARCH:F(1)	[0.101]	[0.558]	[0.964]	[0.951]	[0.462]	[0.336]	[0.102]	[0.310]	[0.399]	[0.353]
N:χ^2(2)	[0.239]	[0.888]	[0.622]	[0.552]	[0.617]	[0.999]	[0.335]	[0.649]	[0.232]	[0.423]
H:F(k-1)	[0.266]	[0.416]	[0.907]	[0.858]	[0.973]	[0.285]	[0.816]	[0.248]	[0.973]	[0.740]
H/FF:F	[0.711]	N/A	[0.779]	[0.956]	[0.962]	N/A	[0.865]	[0.418]	[0.994]	[0.830]
FOR:χ^2(1)	[0.452]	[0.522]	[0.890]	[0.069]	[0.964]	[0.079]	[0.235]	[0.788]	[0.470]	[0.933]
FOR:χ^2(2)	[0.678]	[0.534]	[0.092]	[0.301]	[0.620]	[0.091]	N/A	[0.812]	[0.334]	[0.812]
FOR:χ^2(5)	[0.847]	[0.834]	[0.253]	N/A	[0.769]	N/A	N/A	N/A	[0.656]	[0.945]
FOR:χ^2(9)	[0.773]	[0.974]	[0.009]	N/A	N/A	N/A	N/A	N/A	[0.924]	[0.000]
FOR:χ^2(15)	[0.083]	N/A	N/A	N/A	N/A	N/A	N/A	N/A	N/A	[0.000]

Table 5.2 continued

Country	Ireland	Italy	Japan	Nethrlnd	Norway	Spain	Sweden	Switz.	UK	USA
IV/OLS	IV	OLS	IV	IV	OLS	IV	IV	IV	IV	IV
Drop Instr	NONE	NONE	NONE	NONE	NONE	NONE	NONE	NONE	NONE	NONE
Dummies	7375;82	7679;93	61	648182;	69;78;	87;93	9092	6272;	868788;	NONE
				7791	8586			666879	9192	
Intercept	0.005	0.028	0.014	0.005	0.015	0.01	0.006	0.004	0.004	0.003
	(0.764)	(4.569)	(3.957)	(1.466)	(3.001)	(2.639)	(1.026)	(1.415)	(1.267)	(0.494)
ΔlnY_t	0.881		0.686	0.746		0.765	0.650	0.995	0.794	0.842
	(4.360)		(7.678)	(8.155)		(11.883)	(2.741)	(7.949)	(7.015)	(3.465)
ΔlnY_{t-1}		0.205			0.325			-0.503		
		(2.092)			(2.267)			(-4.311)		
ΔlnY_{t-2}										
$\Delta\Delta lnP_t$			-0.412						-0.227	
			(-5.508)						(-2.939)	
$\Delta\Delta lnP_{t-1}$								-0.136		-0.248
								(-1.947)		(-1.956)
$\Delta\Delta lnP_{t-2}$		-0.347		-0.162						
		(-4.152)		(-2.289)						
$ECM+_{t-1}$	-0.332 (R)	-0.034 (R)	-0.283 (R)	-0.044 (R)	-0.434 (R)	-0.536 (U)	*0.063 (R)*	*0.001 (R)*	-0.299 (U)	-0.154 (R)
	(-1.261)	(-1.517)	(-4.125)	(-0.418)	(-4.552)	(-3.081)	(0.612)	(0.022)	(-1.680)	(-0.503)
$ECM-_{t-1}$	-0.629 (R)	-0.035 (R)	-0.288 (R)	-0.219 (R)	-0.211 (R)	-0.332 (U)	*0.017 (R)*	*0.017 (R)*	-0.233 (U)	-0.308 (R)
	(-2.932)	(-2.087)	(-2.761)	(-3.320)	(-1.179)	(-1.947)	(0.155)	(0.651)	(-1.272)	(-1.369)
$\Sigma\Delta lnY_{t-i}$	0.881	0.205	0.686	0.746	0.325	0.765	0.650	0.492	0.794	0.842
$\Sigma\Delta\Delta lnP_{t-i}$	0.000	-0.347	-0.412	-0.162	0.000	0.000	0.000	-0.136	-0.227	-0.248
Adjusted R^2	0.787	0.767	0.940	0.922	0.746	0.924	0.497	0.884	0.867	0.716
s	0.0143	0.0115	0.0071	0.0061	0.0135	0.0085	0.0157	0.0061	0.0085	0.0091
DW	1.81	2.26	2.41	1.88	2.10	1.60	2.20	2.09	2.40	1.91
Sargan(m)	[0.503](5)	N/A	[0.579](4)	[0.883](4)	N/A	[0.600](5)	[0.648](5)	[0.318](3)	[0.907](4)	[0.233](4)
Symmetry(1)	[0.350]	[0.972]	[0.970]	[0.226]	[0.332]	[0.494]	[0.789]	[0.721]	[0.826]	[0.743]
SC:χ^2(2)	[0.220]	[0.079](F)	[0.537]	[0.963]	[0.579](F)	[0.051]	[0.753]	[0.808]	[0.355]	[0.666]
ARCH:F(1)	[0.557]	[0.497]	[0.608]	[0.782]	[0.736]	[0.329]	[0.590]	[0.318]	[0.222]	[0.321]
N:χ^2(2)	[0.359]	[0.747]	[0.771]	[0.807]	[0.823]	[0.846]	[0.870]	[0.842]	[0.599]	[0.447]
H:F(k-1)	[0.685]	[0.828]	[0.070]	[0.166]	[0.409]	[0.524]	[0.367]	[0.753]	[0.998]	[0.579]
H/FF:F	[0.884]	N/A	[0.083]	N/A	N/A	[0.742]	[0.646]	N/A	N/A	[0.864]
FOR:χ^2(1)	[0.487]	[0.468]	[0.936]	[0.484]	[0.095]	[0.914]	[0.650]	[0.573]	[0.279]	[0.477]
FOR:χ^2(2)	[0.443]	N/A	[0.611]	[0.127]	[0.224]	N/A	[0.110]	[0.789]	[0.615]	[0.538]
FOR:χ^2(5)	[0.545]	N/A	[0.496]	N/A	[0.089]	N/A	N/A	[0.414]	N/A	[0.918]
FOR:χ^2(9)	[0.102]	N/A	[0.475]	N/A	N/A	N/A	N/A	[0.021]	N/A	[0.924]
FOR:χ^2(15)	N/A	N/A	[0.603]	N/A	N/A	N/A	N/A	[0.000]	N/A	[0.922]

Table 5.2. All reported statistics are as for Table 5.1 except "$ECM+_{t-1}$" and "$ECM-_{t-1}$" denote the variables for the positive and negative error correction terms. "Symmetry" is the F-version of a Wald test for the single symmetry restriction (produced by PcGive 8.0). Bold emphasis indicates rejection of the symmetry restriction.

symmetry restriction.[22] This improvement in fit combined with the clearly different adjustment coefficients on the error correction terms, is taken as evidence of non-symmetric adjustment to equilibrium for these three countries.[23] It is interesting to note that the coefficient is larger on the negative error correction term for Austria and the Netherlands, possibly reflecting a greater reluctance to adjust living standards down rather than up in these two countries. In contrast, the positive error correction term's coefficient is larger for France. This might be due to consumers greater ability to adjust towards equilibrium in good times (when above equilibrium) than in bad times.

5.5 Cubic Nonlinear Adjustment to Equilibrium

Another interesting hypothesis is that agents will adjust more quickly the further their actual consumption is from equilibrium - the urgency and pressure to move toward the long run level increases as the degree of disequilibrium rises. This might be especially true of consumption which incorporates a durable component, as in the present application. For example, Caballero (1994) argues that fixed adjustment costs may cause agents to tolerate small departures from an ever changing optimal durable stock. However, once the deviation from equilibrium becomes large, abrupt buying or selling may occur to secure tolerable disequilibrium - adjustment is faster the further an agent is from equilibrium. Such nonlinear adjustment may be portrayed by specifying error correction terms raised to powers greater than one, such as the cubic formulation employed by Hendry and Ericsson (1991).[24] The general form of the *cubic* error correction model is:

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[22] For Ireland the fit of the symmetric model (s=0.0142936) is slightly superior to that of the asymmetric specification (s=0.0143192) suggesting no evident asymmetry according to our criteria.

[23] Although the fit of the Swedish equation improves with the relaxation of the symmetry restriction the coefficients are not consistent with error correction behaviour because they are positive. Therefore, we do not suggest there is evident asymmetry for this country.

[24] Previous applications include Hendry and Ericsson (1991), who find evidence of cubic long run adjustment for UK money demand and Sarantis and Hadjimatheou (1998) who find evidence of quadratic adjustment for UK manufacturing employment.

$$\Delta \ln C_t = f_0 + f_{11}\Delta \ln C_{t-1} + f_{20}\Delta \ln \hat{Y}_t + \Sigma f_{2i}\Delta \ln Y_{t-i} + f_{30}\Delta\Delta \ln \hat{P}_t + \Sigma f_{3i}\Delta\Delta \ln P_{t-i} + \Sigma f_{4j}D_{jt} \qquad \textbf{(5.10)}$$

$$+ a_{11}ECM_{t-1} + a_{12}ECM^2_{t-1} + a_{13}ECM^3_{t-1} + u_{3t}.$$

Once again we use the error correction terms based upon our favoured long run consumption functions estimated in Chapter 4 and then subtract their mean. This ensures that a zero error implies equilibrium and the magnitude of error reflects the degree of disequilibrium. Further, squaring and cubing these demeaned error correction terms will appropriately depict amplified distances from the long run value of consumption, hence portraying faster adjustment the further agents are from equilibrium.[25]

This formulation may also yield a different degree of adjustment when consumption is above equilibrium relative to when it is below its long run value. The quadratic term is particularly responsible for this because the process of squaring makes all of the errors positive so that adjustment from this term is only in one direction. Therefore, the sign of the coefficient on the quadratic error correction term will, in general, indicate the direction of greatest adjustment. However, because this term only corrects errors in one direction it cannot yield valid error correction behaviour in the absence of other error correction terms. For this reason we report models incorporating all three components of the cubic error correction function for all countries.

However, this full cubic specification may involve statistically insignificant error correction terms, suggesting that the model is overparameterised. Thus, we additionally consider model reduction of the cubic specification. In this formulation we propose the preservation of the sign

[25] The coefficients on the squared and cubed error correction terms need not fall between zero and one in magnitude. Indeed, one would not expect this to be the case because squaring and/or cubing the error correction series decreases (increases) the size of numbers smaller (greater) in magnitude than unity and so would raise (lower) their coefficients. In the present application the demeaned error correction terms take on values of less than one in absolute value and so their estimated parameters may turn out to be much larger than unity. Thus, the magnitude of coefficients does not provide a check on whether the error correction is appropriate.

of the elements in the quadratic error correction term - negative errors are squared and then multiplied by minus one while positive errors are simply squared. This facilitates model reduction such that, where appropriate, the squared error correction term can be entered on its own whilst characterising valid error correction behaviour.[26] One further consequence of preserving the sign of the elements of the squared term is that it will substantially reduce the sensitivity of the degree of adjustment to whether consumption is above or below its equilibrium value. Instead, this modified nonlinear formulation will primarily characterise the extent of the adjustment to equilibrium as being sensitive to the degree of disequilibrium. As suggested earlier, this form of adjustment may be appropriate for consumer behaviour where there is a component of durability in the expenditure series. We are not aware of previous studies which preserve the sign of the elements in the squared error correction term in such a cubic nonlinear specification.

For both forms of (cubic) nonlinear specification the inclusion of more than one error correction term can render determination of the validity of error correction behaviour extremely difficult. That is, whether positive (negative) errors last period lead to reduced (increased) consumption this period. This is the case when the error correction terms' adjustment coefficients exhibit opposite signs. We check for valid error correction behaviour in both forms of cubic nonlinear specification by directly calculating the *overall* error correction *function*'s average contribution to consumption for negative errors (when consumption is below equilibrium) and positive errors (when consumption is above its long run value).

We propose the following procedure, which we have not seen previously applied, to assess the validity of error correction behaviour in a cubic nonlinear model. The steps are:

1) Partition the error correction term into positive and negative components, which yield the same terms constructed for the *partitioned* asymmetric models, $ECM+_{t-1}$ and $ECM-_{t-1}$, respectively. This indicates when consumption is above or below equilibrium.

2) Calculate the squared and cubed values of these partitioned terms: $ECM\text{-}^2_{t-1} = [ECM+_{t-1}]^2$;

[26] If the sign of the elements in the squared error correction term were not preserved and it was the sole error correction term specified in the model it would not provide coherence with the long run equilibrium because adjustment would only ever occur in one direction.

$ECM-^2_{t-1} = [ECM-_{t-1}]^2$ or $-[ECM-_{t-1}]^2$ when the sign is preserved; $ECM+^3_{t-1} = [ECM+_{t-1}]^3$; and $ECM-^3_{t-1} = [ECM-_{t-1}]^3$. This evaluates the amplification of error correction when consumption is above or below equilibrium.

3) Define the overall error correction for positive and negative errors, respectively, as: $fECM+_{t-1} = a_{11}ECM+_{t-1} + a_{12}ECM+^2_{t-1} + a_{13}ECM+^3_{t-1}$ and, $fECM-_{t-1} = a_{11}ECM-_{t-1} + a_{12}ECM-^2_{t-1} + a_{13}ECM-^3_{t-1}$. Where a_{1i} are the estimated coefficients on the error correction terms from equation **(5.10)**.

4) Calculate the mean of the overall error corrections for the period 1960-1993. That is, $AvgECM+ = \Sigma(fECM+_{t-1})/T$ and $AvgECM- = \Sigma(fECM-_{t-1})/T$, where T is the number of observations (34 in the current application). Since the error correction terms are lagged by one period 1960-1993 defines when they impact upon the model (which is estimated over the sample 1961-1994).

5) For appropriate error correction Avg(ECM+) should be negative and Avg(ECM-) positive: when agents are above (below) equilibrium their error is positive (negative) so consumption should be reduced (increased).

In addition to checking the consistency of the model with plausible error correction behaviour one can also assess whether the nonlinear error correction terms are statistically significant by testing the following joint hypothesis with a Wald (IV) or F (OLS) test:[27]

$$H_0: a_{12} = a_{13} = 0. \qquad \textbf{(5.11)}$$

Cook, Holly and Turner (1998) cite Monte-Carlo evidence which suggests that such tests of nonlinear error correction terms' statistical significance have low power. Therefore, we also compare the fit of these nonlinear models with the standard symmetric specifications (reported in Table 5.1) to determine whether nonlinearities improve model specification.

Table 5.3 presents the estimation results of the cubic nonlinear error correction models which include all three error correction terms (referred to as the cubic model hereafter). All reported

[27] Only the nonlinear error correction terms are tested because our interest is whether they have significant explanatory power beyond or instead of the *normal* error correction term. When either the squared or cubic term is absent the test involves a single restriction.

TABLE 5.3: Cubic Nonlinear Asymmetric Error Correction Models

Country	Australia	Austria	Belgium	Canada	Denmark	Finland	France	Germany	Greece	Iceland
IV/OLS	IV	IV	IV	IV	IV	IV	IV	IV	IV	IV
Drop Instr	NONE	NONE	NONE	NONE	NONE	NONE	NONE	$\Delta\Delta \ln P_{t-2}$	NONE	NONE
Dummies	6775	77;78;83	83	82;91	86	6988;71;	7493	91	80	75
						9192				
Intercept	0.013	0.006	0.013	0.001	-0.013	0.009	0.025	0.006	0.002	0.021
	(4.638)	(1.200)	(3.347)	(0.152)	(-1.873)	(1.499)	(4.601)	(1.593)	(4.855)	(2.737)
$\Delta \ln Y_t$	0.539	0.399	0.446	0.834	0.901	0.787	0.377	0.666	0.302	0.593
	(4.054)	(2.966)	(3.736)	(4.065)	(4.795)	(4.872)	(1.972)	(4.388)	(3.020)	(5.392)
$\Delta \ln Y_{t-1}$								0.106		
								(1.396)		
$\Delta \ln Y_{t-2}$	-0.182	0.224					-0.210			-0.181
	(-3.928)	(3.049)					(-2.808)			(-2.579)
$\Delta\Delta \ln P_{t-1}$			-0.257							
			(-2.788)							
$\Delta \ln C_{t-1}$					0.738					
					(3.918)					
ECM_{t-1}	-0.070 (R)	-0.503 (R)	-0.267 (R)	-0.291 (R)	-0.031 (U)	-0.703 (U)	-0.211 (U)	-0.043 (R)	-0.365(R)	-0.346(R)
	(-0.725)	(-3.076)	(-2.886)	(-1.281)	(-0.366)	(-3.471)	(-1.912)	(-0.262)	(-3.828)	(-1.997)
ECM^2_{t-1}	-0.565 (R)	3.800 (R)	0.427 (R)	6.898 (R)	-0.299 (U)	-1.686 (U)	-0.648 (U)	1.625 (R)	0.551 (R)	-0.480(R)
	(-0.305)	(1.607)	(0.500)	(0.670)	(-0.537)	(-0.322)	(-0.719)	(0.635)	(0.468)	(-0.428)
ECM^3_{t-1}	-46.956(R)	167.07(R)	38.378(R)	-28.436(R)	-2.989(U)	265.94(U)	6.111 (U)	-58.990(R)	32.046(R)	-3.587(R)
	(-0.743)	(1.632)	(2.079)	(-0.096)	(-0.729)	(1.795)	(0.345)	(-0.647)	(1.501)	(-0.256)
$\Sigma\Delta \ln Y_{t-i}$	0.357	0.623	0.446	0.834	0.901	0.787	0.167	0.772	0.302	0.412
$\Sigma\Delta\Delta \ln P_{t-i}$	-0.070	-0.503	-0.781	-0.291	0.000	-0.703	-0.211	-0.043	-0.365	-0.346
AvgECM+	-0.00132	-0.00221	-0.00177	-0.00123	-0.00258	-0.00449	-0.00520	-0.00109	-0.00470	-0.00907
AvgECM-	+0.00098	+0.0037	+0.00274	+0.00364	+0.00123	+0.00371	+0.00328	+0.00192	+0.00578	+0.00833
Adj R^2	0.790	0.878	0.716	0.799	0.502	0.879	0.834	0.873	0.827	0.849
s	0.0061	0.0062	0.0102	0.0117	0.0207	0.0116	0.0064	0.0088	0.0109	0.0243
DW	2.45	1.78	2.09	1.40	2.16	1.78	2.12	1.86	2.56	2.41
Sargan(m)	[0.742](4)	[0.853](4)	[0.765](4)	[0.405](5)	[0.132](4)	[0.442](5)	[0.803](4)	[0.653](3)	[0.506](5)	[0.076](4)
Nonlinear	[0.745](2)	[0.191](2)	[0.115](2)	[0.496](2)	[0.742](2)	[0.152](2)	[0.323](2)	[0.783](2)	[0.317](2)	[0.913](2)
SC:$\chi^2(2)$	[0.495]	[0.209]	[0.676]	[0.031]	[0.537]	[0.637]	[0.695]	[0.754]	[0.157]	[0.094]
ARCH:F(1)	[0.167]	[0.268]	[0.745]	[0.782]	[0.427]	[0.433]	[0.135]	[0.514]	[0.689]	[0.374]
N:$\chi^2(2)$	[0.237]	[0.489]	[0.927]	[0.665]	[0.491]	[0.903]	[0.525]	[0.890]	[0.394]	[0.409]
H:F(k-1)	[0.499]	[0.795]	[0.930]	[0.873]	[0.966]	[0.794]	[0.825]	[0.469]	[0.947]	[0.830]
H/FF:F	N/A	N/A	N/A	N/A	N/A	N/A	N/A	N/A	[0.991]	N/A
FOR:$\chi^2(1)$	[0.483]	[0.205]	[0.578]	[0.093]	[0.896]	[0.199]	[0.238]	[0.850]	[0.606]	[0.874]
FOR:$\chi^2(2)$	[0.643]	[0.464]	[0.168]	[0.332]	[0.659]	[0.442]	N/A	[0.862]	[0.489]	[0.804]
FOR:$\chi^2(5)$	[0.907]	[0.704]	[0.467]	N/A	[0.826]	N/A	N/A	N/A	[0.783]	[0.657]
FOR:$\chi^2(9)$	[0.762]	[0.681]	[0.000]	N/A	N/A	N/A	N/A	N/A	[0.421]	[0.000]
FOR:$\chi^2(15)$	[0.000]	N/A	N/A	N/A	N/A	N/A	N/A	N/A	N/A	[0.000]

Table 5.3 continued

Country	Ireland	Italy	Japan	Netherlnd	Norway	Spain	Sweden	Switzerlnd	UK	USA
IV/OLS	IV	OLS	IV	IV	OLS	IV	IV	IV	IV	IV
Drop Instr	NONE	NONE	NONE	NONE	NONE	NONE	NONE	NONE	NONE	NONE
Dummies	7375;82	7679;93	61	648182;	69;78;8586	87;93	9092	6272;	868788;	NONE
				7791				666879	9192	
Intercept	0.006	0.028	0.016	0.006	0.015	0.007	0.004	0.002	0.004	0.005
	(0.380)	(6.043)	(3.507)	(1.726)	(3.344)	(2.186)	(0.791)	(0.560)	(1.380)	(0.789)
$\Delta \ln Y_t$	0.895		0.650	0.742		0.793	0.712	1.063	0.793	0.828
	(4.001)		(6.145)	(7.880)		(11.693)	(3.031)	(6.982)	(6.871)	(3.501)
$\Delta \ln Y_{t-1}$		0.203			0.307			-0.540		
		(2.037)			(2.183)			(-4.061)		
$\Delta\Delta \ln P_t$			-0.395						-0.222	
			(-4.421)						(-2.922)	
$\Delta\Delta \ln P_{t-1}$								-0.140		-0.243
								(-1.835)		(-1.928)
$\Delta\Delta \ln P_{t-2}$		-0.345		-0.169						
		(-3.824)		(-2.436)						
ECM_{t-1}	-0.506 (R)	-0.033 (R)	-0.368 (R)	-0.192 (R)	-0.459 (R)	-0.287 (U)	0.165 (R)	0.051 (R)	-0.282 (U)	-0.008 (R)
	(-2.410)	(-2.164)	(-3.493)	(-2.187)	(-3.314)	(-1.666)	(1.622)	(1.398)	(-1.763)	(-0.030)
ECM^2_{t-1}	1.811 (R)	0.003 (R)	-0.287 (R)	2.539 (R)	-4.612 (R)	0.824 (U)	0.893 (R)	0.005 (R)	-0.882 (U)	0.366 (U)
	(0.488)	(0.079)	(-0.329)	(1.828)	(-1.754)	(0.198)	(1.056)	(0.033)	(-0.245)	(0.045)
ECM^3_{t-1}	6.617 (R)	-0.009 (R)	14.469 (R)	35.242 (R)	51.501 (R)	-156.65(U)	-18.875(R)	-2.638 (R)	17.681(U)	-508.71(R)
	(0.082)	(-0.110)	(0.974)	(1.190)	(1.538)	(-0.994)	(-1.572)	(-1.266)	(0.138)	(-1.087)
$\Sigma\Delta \ln Y_{t-i}$	0.895	0.203	0.255	0.742	0.307	0.793	0.712	0.523	0.793	0.828
$\Sigma\Delta\Delta \ln P_{t-i}$	0.000	-0.345	-0.395	-0.169	0.000	0.000	0.000	-0.140	-0.222	-0.243
AvgECM+	-0.00447	-0.00450	-0.00422	-0.00086	-0.00636	-0.00246	*+0.00274*	*+0.00086*	-0.00195	-0.00120
AvgECM-	+0.00562	+0.00503	+0.00430	+0.00317	+0.00350	+0.00252	-0.00006	*-0.00021*	+0.00171	+0.00137
Adj R^2	0.774	0.759	0.940	0.925	0.757	0.921	0.514	0.861	0.863	0.722
s	0.0147	0.0117	0.0071	0.0060	0.0133	0.0087	0.0154	0.0067	0.0086	0.0090
DW	1.78	2.26	2.23	1.87	2.04	1.74	2.16	2.17	2.37	2.08
Sargan(m)	[0.460](5)	N/A	[0.480](4)	[0.535](4)	N/A	[0.714](5)	[0.842](5)	[0.457](3)	[0.893](4)	[0.290](4)
Nonlinear	[0.375](2)	[0.982](F,2)	[0.616](2)	[0.184](2)	[0.233](F,2)	[0.583](2)	[0.285](2)	[0.405](2)	[0.968](2)	[0.541](2)
SC:$\chi^2(2)$	[0.250]	[0.074](F)	[0.672]	[0.912]	[0.606](F)	[0.142]	[0.648]	[0.681]	[0.406]	[0.582]
ARCH:F(1)	[0.630]	[0.459]	[0.912]	[0.975]	[0.850]	[0.593]	[0.622]	[0.532]	[0.232]	[0.363]
N:$\chi^2(2)$	[0.425]	[0.776]	[0.475]	[0.686]	[0.321]	[0.525]	[0.305]	[0.800]	[0.558]	[0.657]
H:F(k-1)	[0.680]	[0.901]	[0.552]	[0.202]	[0.699]	[0.782]	[0.723]	[0.511]	[0.998]	[0.593]
H/FF:F	N/A	N/A	N/A	N/A	N/A	N/A	[0.940]	N/A	N/A	[0.917]
FOR:$\chi^2(1)$	[0.519]	[0.457]	[0.851]	[0.812]	[0.104]	[0.938]	[0.392]	[0.346]	[0.254]	[0.597]
FOR:$\chi^2(2)$	[0.426]	N/A	[0.774]	[0.704]	[0.229]	N/A	[0.183]	[0.257]	[0.086]	[0.678]
FOR:$\chi^2(5)$	[0.343]	N/A	[0.098]	N/A	**[0.002]**	N/A	N/A	[0.564]	N/A	[0.960]
FOR:$\chi^2(9)$	[0.113]	N/A	[0.165]	N/A	N/A	N/A	N/A	[0.477]	N/A	[0.951]
FOR:$\chi^2(15)$	N/A	N/A	[0.699]	N/A	N/A	N/A	N/A	[0.000]	N/A	[0.913]

Table 5.3. All reported statistics are as for Table 5.1 except "ECM^2_{t-1}" and "ECM^3_{t-1}" denote the squared and cubed

error correction terms, respecitively. All three error correction terms are included in each equation. "AvgECM+" and "AvgECM-" are the average adjustments towards equilibrium when consumption is above and below its long run value, respectively. Bold emphasis indicates whether the greater adjustment is downwards or upwards. "Nonlinearity" is a Wald χ^2 (IV) / F (OLS) test for the joint statistical significance of the *higher power* error correction terms, produced by Microfit 3.22. The degrees of freedom equal the number of error correction terms and are given in (normal brackets). Bold emphasis indicates statistically significant nonlinearity.

statistics are the same as for Table 5.1 except "ECM^2_{t-1}" and "ECM^3_{t-1}" denote the squared and cubed error correction terms, respectively. "AvgECM+" and "AvgECM-" indicate the average adjustments towards equilibrium when consumption is above and below its long run value, respectively. "Nonlinearity" is a Wald χ^2 (IV) / F (OLS) test for the joint statistical significance of the *higher power* error correction terms, produced by Microfit 3.22. The degrees of freedom equal the number of nonlinear error correction terms being tested and are given in (normal brackets).

Only for Canada is there evidence of statistically significant autocorrelation at the 5% level. Since autocorrelation is not evident for any country, including Canada, at the 1% level we argue that serial correlation is not adversely affecting inference. The only other form of evident misspecification is the numerical non constancy of parameters for Australia, Belgium, Iceland, Norway and Switzerland. We will be cautious when drawing inference for these countries' equations.

The average adjustment to equilibrium is downward (negative) when consumption is above equilibrium (AvgECM+) and upward (positive) when consumption is below equilibrium in all cases, except Sweden and Switzerland. Thus these cubic error correction models are consistent with valid error correction behaviour except for Sweden and Switzerland. This is consistent with our cointegration results from Chapter 4 and the results from the error correction models reported in Table 5.1.[28] For all countries the joint test for the statistical significance of the nonlinear terms is rejected. This apparent evidence against nonlinearities may be due to the low power of the test and/or overparameterisation. To the extent that low power is responsible for this evidence against nonlinearity, we use fit measures to assess whether the addition of nonlinear error correction terms improve model specification relative to the *standard symmetric* models reported in Table 5.1. For Austria, Belgium, Finland, France, the Netherlands and Norway, the cubic nonlinear error correction model provides superior fit relative to the standard *symmetric* specification.[29] We interpret this as evidence of nonlinear adjustment towards equilibrium for these countries, although we recognise that the models for Belgium and Norway exhibit some evidence of parameter instability. We also note that for Austria, Belgium and the Netherlands, the magnitude of average downward adjustment is greater than upward adjustment. While upward adjustment is larger than downward adjustment for Finland, France, and Norway.

Table 5.4 presents the favoured estimated nonlinear error correction formulations where the sign of the quadratic error correction term is preserved to facilitate sensible model reduction. There is some evidence of misspecification at the 5% level for Australia (non-constant parameters), Canada (serial correlation) and Iceland (serial correlation and non-constant parameters) but not

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[28] Interestingly, greater upward (downward) adjustment occurs for all countries whose adjustment coefficient on the squared error correction term is positive (negative), except for Japan. In Japan's case the adjustment coefficient on the squared error correction term is extremely small (less than the coefficient on the untransformed error correction term) suggesting it has little impact in the overall error correction function and therefore explains this anomaly. Indeed, the difference in error correction when consumption is above equilibrium (the average is -0.00422) and when it is below its long run value (the average is +0.00430) is very small for this country. As expected, it is the squared error correction term's coefficient which determines the direction of greatest adjustment.

[29] The fit of the cubic nonlinear model is greater compared to that of the standard symmetric formulation for Sweden. However, we do not interpret this as evidence favouring nonlinear error correction behaviour because there is no valid error correction behaviour for this country.

TABLE 5.4: Reduced Nonlinear Asymmetric Error Correction Models

Country	Australia	Austria	Belgium	Canada	Denmark	Finland	France	Germany	Greece	Iceland
IV/OLS	IV	IV	IV	IV	IV	IV	IV	IV	IV	IV
Drop Instr	NONE	NONE	NONE	NONE	NONE	NONE	NONE	$\Delta\Delta \ln P_{t-2}$	NONE	NONE
Dummies	6775	77;78;83	83	82;91	86	6988;71;	7493	91	80	75
						9192				
Intercept	0.014	0.009	0.013	0.003	-0.013	0.008	0.023	0.007	0.022	0.02
	(5.665)	(1.576)	(3.625)	(0.457)	(-1.891)	(2.025)	(4.514)	(1.931)	(5.412)	(3.635)
$\Delta \ln Y_t$	0.477	0.398	0.449	0.825	0.859	0.791	0.427	0.662	0.340	0.639
	(4.573)	(2.883)	(3.900)	(4.558)	(5.316)	(5.654)	(2.633)	(4.264)	(3.841)	(6.194)
$\Delta \ln Y_{t-1}$								0.131		
								(1.845)		
$\Delta \ln Y_{t-2}$	-0.19	0.202					-0.223			-0.17
	(-4.515)	(2.949)					(-2.867)			(-2.605)
$\Delta\Delta \ln P_{t-1}$			-0.262							
			(-2.942)							
$\Delta \ln C_{t-1}$					0.701					
					(4.341)					
ECM_{t-1}	-0.690 (R)	-0.503 (R)	-0.375 (R)							
	(-2.203)	(-2.386)	(-2.785)							
ECM^2_{t-1}	38.870 (R)	5.623 (R)	4.428 (R)	-14.206(R)	-0.601(U)	-53.644(U)	-7.035 (U)		-13.042(R)	-9.049 (R)
	(2.011)	(1.076)	(2.167)	(-3.845)	(-1.681)	(-4.593)	(-2.683)		(-4.342)	(-2.782)
ECM^3_{t-1}	-604.58(R)					1135.6 (U)	65.275 (U)	-58.271(R)	141.95 (R)	47.339 (R)
	(-2.116)					(3.883)	(2.197)	(-2.206)	(3.434)	(1.809)
$\Sigma\Delta \ln Y_{t-i}$	0.287	0.600	0.449	0.825	0.859	0.791	0.204	0.793	0.340	0.471
$\Sigma\Delta\Delta \ln P_{t-i}$	0.000	0.000	-0.262	0.000	0.000	0.000	0.000	0.000	0.000	0.000
AvgECM+	-0.001544	-0.002842	-0.002116	-0.001625	-0.001575	-0.003909	-0.004097	-0.001346	-0.004689	-0.008706
AvgECM-	+0.001565	+0.002953	+0.002569	+0.002850	+0.00187	+0.003721	+0.003552	+0.000754	+0.004726	+0.007606
Adj R^2	0.829	0.870	0.731	0.812	0.561	0.893	0.831	0.879	0.853	0.864
s	0.0055	0.0064	0.0099	0.0113	0.0194	0.0108	0.0065	0.0086	0.0100	0.0231
DW	2.72	1.70	2.17	1.43	2.13	1.75	2.13	1.86	2.69	2.44
Sargan(m)	[0.546](4)	[0.925](4)	[0.716](4)	[0.561](5)	[0.236(4)	[0.528](5)	[0.646](4)	[0.689](3)	[0.386](5)	[0.058](4)
Nonlinear	[0.101](2)	[0.282](1)	[0.030](1)	[0.000](1)	[0.093](1)	[0.000](2)	[0.003](2)	[0.027](1)	[0.000](2)	[0.000](2)
SC:$\chi^2(2)$	[0.098]	[0.148]	[0.620]	[0.038]	[0.455]	[0.624]	[0.548]	[0.897]	[0.077]	[0.043]
ARCH:F(1)	[0.923]	[0.478]	[0.832]	[0.734]	[0.409]	[0.564]	[0.345]	[0.348]	[0.750]	[0.357]
N:$\chi^2(2)$	[0.286]	[0.735]	[0.979]	[0.817]	[0.355]	[0.919]	[0.695]	[0.892]	[0.264]	[0.398]
H:F(k-1)	[0.518]	[0.661]	[0.930]	[0.818]	[0.939]	[0.615]	[0.498]	[0.171]	[0.976]	[0.892]
H/FF:F	N/A	N/A	[0.641]	[0.846]	[0.941]	N/A	[0.881]	[0.282]	[0.989]	[0.599]
FOR:$\chi^2(1)$	[0.557]	[0.345]	[0.528]	[0.127]	[0.568]	[0.252]	[0.265]	[0.930]	[0.596]	[0.889]
FOR:$\chi^2(2)$	[0.800]	[0.572]	[0.156]	[0.400]	[0.657]	[0.081]	N/A	[0.677]	[0.473]	[0.662]
FOR:$\chi^2(5)$	[0.866]	[0.836]	[0.432]	N/A	[0.804]	N/A	N/A	N/A	[0.772]	[0.764]
FOR:$\chi^2(9)$	[0.403]	[0.986]	[0.862]	N/A	N/A	N/A	N/A	N/A	[0.967]	[0.044]
FOR:$\chi^2(15)$	[0.019]	N/A	N/A	N/A	N/A	N/A	N/A	N/A	N/A	[0.028]

Table 5.4 continued

Country	Ireland	Italy	Japan	Netherlnds	Norway	Spain	Sweden	Switzerlnd	UK	USA
IV/OLS	IV	OLS	IV	IV	OLS	IV	IV	IV	IV	IV
Drop Instr	NONE	NONE	NONE	NONE	NONE	NONE	NONE	NONE	NONE	NONE
Dummies	7375;82	7679;93	61	648182;	69;78;8586	87;93	9092	6272;	868788;	NONE
				7791				666879	9192	
Intercept	0.003	0.028	0.017	0.008	0.013	0.007	0.006	0.002	0.003	0.005
	(0.432)	(7.710)	(3.460)	(2.670)	(2.918)	(2.625)	(1.467)	(1.043)	(1.213)	(1.238)
ΔlnY_t	0.983		0.620	0.750		0.814	0.760	1.042	0.818	0.805
	(3.775)		(5.403)	(7.773)		(13.449)	(3.147)	(7.209)	(6.993)	(4.031)
ΔlnY_{t-1}		0.210			0.324			-0.526		
		(2.209)			(2.161)			(-4.138)		
$\Delta\Delta lnP_t$			-0.387						-0.191	
			(-4.380)						(-2.506)	
$\Delta\Delta lnP_{t-1}$								-0.132		-0.246
								(-1.826)		(-2.063)
$\Delta\Delta lnP_{t-2}$		-0.325		-0.176						
		(-4.004)		(-2.438)						
ECM_{t-1}			-0.439 (R)	-0.461 (R)						
			(-2.756)	(-1.961)						
ECM^2_{t-1}	-10.986(R)	-0.171 (R)	1.959 (R)	13.885 (R)	-11.398 (R)	-14.344(U)		0.775 (R)	-8.315 (U)	
	(-2.537)	(-2.438)	(1.027)	(1.330)	(-2.459)	(-4.908)		(1.319)	(-2.538)	
ECM^3_{t-1}		0.191 (R)		-137.30(R)	75.383 (R)		-1.297	-5.391 (R)		-534.6(R)
		(1.355)		(-1.260)	(1.569)		(-0.178)	(-1.280)		(-2.461)
$\Sigma\Delta lnY_{t-i}$	0.982	0.210	0.233	0.750	0.324	0.814	0.760	0.516	0.818	0.805
$\Sigma\Delta\Delta lnP_{t-i}$	0.000	-0.325	-0.387	-0.176	0.000	0.000	0.000	-0.132	-0.191	-0.246
AvgECM+	-0.002778	-0.004103	-0.004358	-0.002675	-0.004932	-0.002091	-0.000222	*+0.029770*	-0.001279	-0.001248
AvgECM-	+0.00491	+0.004943	+0.004758	+0.002277	+0.0034595	+0.001965	+0.00018	*-0.025126*	+0.00114	+0.00134
Adj R^2	0.754	0.774	0.940	0.920	0.725	0.924	0.494	0.872	0.870	0.745
s	0.0154	0.0113	0.0071	0.0062	0.0141	0.0085	0.0157	0.0064	0.0084	0.0086
DW	2.01	2.28	2.22	1.67	1.94	1.86	2.12	2.17	2.64	2.09
Sargan(m)	[0.412](5)	N/A	[0.595](4)	[0.725](4)	N/A	[0.931](5)	[0.567](5)	[0.381](3)	[0.909](4)	[0.450](4)
Nonlinear	[0.011](1)	[0.001](F,2)	[0.305](1)	[0.409](2)	[0.000](F,2)	[0.000](1)	[0.858](1)	[0.419](2)	[0.011](1)	[0.014](1)
SC:$\chi^2(2)$	[0.415]	[0.086](F)	[0.691]	[0.702]	[0.816](F)	[0.421]	[0.887]	[0.672]	[0.115]	[0.513]
ARCH:F(1)	[0.987]	[0.455]	[0.637]	[0.134]	[0.894]	[0.795]	[0.580]	[0.376]	[0.501]	[0.298]
N:$\chi^2(2)$	[0.854]	[0.825]	[0.392]	[0.461]	[0.946]	[0.483]	[0.844]	[0.785]	[0.833]	[0.660]
H:F(k-1)	[0.447]	[0.971]	[0.606]	[0.703]	[0.502]	[0.719]	[0.719]	[0.604]	[0.963]	[0.755]
H/FF:F	[0.549]	[0.430]	[0.463]	N/A	[0.768]	[0.824]	[0.859]	N/A	[0.992]	[0.833]
FOR:$\chi^2(1)$	[0.314]	[0.396]	[0.859]	[0.426]	[0.080]	[0.884]	[0.540]	[0.397]	[0.369]	[0.619]
FOR:$\chi^2(2)$	[0.324]	N/A	[0.761]	[0.125]	[0.210]	N/A	[0.096]	[0.487]	[0.541]	[0.727]
FOR:$\chi^2(5)$	[0.600]	N/A	[0.149]	N/A	[0.124]	N/A	[0.320]	[0.581]	N/A	[0.953]
FOR:$\chi^2(9)$	[0.281]	N/A	[0.136]	N/A	N/A	N/A	N/A	[0.659]	N/A	[0.946]
FOR:$\chi^2(15)$	N/A	N/A	[0.667]	N/A	N/A	N/A	N/A	[0.295]	N/A	[0.924]

Table 5.4 notes. All reported statistics are as for Table 5.3, excepting the preserved sign for the squared ECM term.

at the 1% level. Inference from these models is, therefore, presented as legitimate.

The average adjustment to equilibrium is downward (negative) when consumption is above equilibrium (AvgECM+) and upward (positive) when consumption is below equilibrium for all countries, except Switzerland.[30] Although the Swedish equation provides adjustment in the correct direction the retained (cubed) error correction term is statistically insignificant (its t-ratio is -0.178). This suggests that Sweden's model does not provide genuine coherence with the long run equilibrium. Thus, the cubic error correction models are consistent with valid error correction behaviour for all countries, except Sweden and Switzerland.

The nonlinear error correction terms are found to feature joint statistical significance for thirteen of the twenty countries (Belgium, Canada, Finland, France, Germany, Greece, Iceland, Ireland, Italy, Norway, Spain, the UK and the USA). However, this does not necessarily imply that nonlinear adjustment is preferred to linear adjustment because the reduced models can exclude linear error correction terms. The models with reduced cubic nonlinear adjustment exhibit improved fit relative to the standard *symmetric* adjustment specification (Table 5.1) for nine countries: Australia, Austria, Belgium, Denmark, Finland, Germany, Greece, Iceland, and the USA.[31] We interpret this as evidence favouring nonlinear adjustment for many OECD countries.

5.6 Conclusion

We have utilised the long run consumption functions (based upon consumption, income and inflation) estimated by the Johansen procedure in Chapter 4 to develop error correction models

[30] The average value of downward adjustment is similar to upward adjustment for all countries except Canada (average downward adjustment is -0.001625 while upward adjustment is +0.002850), Germany (-0.001346 and +0.000754), Ireland (-0.002778 and +0.004910) and Norway (-0.004932 and +0.003460). Thus, it is only for these countries where nonlinearity causes differences in the speed of adjustment when consumption is above equilibrium compared to when it is below it.

[31] In the case of Denmark the quadratic term is the sole means of error correction. This model could only have been selected, with valid error correction, if the elements of this term have their signs preserved.

for twenty OECD countries. Models which allow the standard symmetric linear adjustment towards equilibrium as well as formulations with asymmetric nonlinear adjustment are produced for all countries. Standard error correction models which are coherent with the specified long run equilibrium are secured for all countries except Sweden and Switzerland. This is consistent with the results from Chapter 4 where reasonably specified cointegrating vectors were obtained for all but these two countries. There are also some minor reservations regarding the validity of error correction behaviour in the model for Denmark - this is also consistent with the cointegration results of Chapter 4. These models provide satisfactory explanatory power with the adjusted R^2 being, on average, 80.4%, and never being below 50%. The estimated adjustment coefficients and overall short run income and inflation elasticities are clearly heterogeneous across countries. Explaining the cross country variation of these estimates is the subject of Chapter 7. Interestingly only seven countries' models incorporate short run inflation effects. Further, only twelve countries models embody statistically significant inflation effects in either the short or long run part of their equation. This suggests that inflation may not be a fundamental determinant of consumer behaviour for many countries.

These standard error correction models were then extended to allow asymmetric nonlinear adjustment towards long run equilibrium. Three asymmetric adjustment formulations were considered: the partitioned asymmetric form, the full cubic nonlinear form and the reduced cubic form; the latter specification preserves the sign of the elements in the quadratic error correction term. The models are consistent with valid error correction behaviour for all countries (except Sweden and Switzerland) for all three forms of asymmetric/nonlinear adjustment. The Wald test suggests no evidence of asymmetry/nonlinearity for all countries for the first two specifications but indicates asymmetry/nonlinearity for thirteen countries in the reduced cubic model. However, the low power of this test and the potential removal of linear error correction terms from the reduced cubic model suggests that these Wald test results may not accurately reflect the prevalence of asymmetric/nonlinear adjustment. We therefore compare the fit of the asymmetric adjustment models with the standard symmetric specification. The partitioned asymmetric adjustment model exhibits greater fit than the standard model for three countries. The full cubic nonlinear specification exhibits improved fit for six countries while the reduced cubic formulation features superior fit to the standard model for nine economies. Comparing the fit of all four specifications for the eighteen countries where there is valid error correction behaviour we find that the standard form is preferred for six countries (Canada, Ireland, Italy, Japan, Spain and the UK), the partitioned specification for one country (France), the full cubic model for three countries (Austria, the Netherlands and Norway) and the reduced cubic form for eight countries (Australia, Belgium, Denmark, Finland, Germany, Greece, Iceland and the USA). We interpret this as suggesting asymmetric/nonlinear adjustment towards equilibrium for the majority of OECD countries. Further, when asymmetric adjustment exists, the cubic nonlinear form is generally preferred. This suggests that the speed of adjustment is generally a function of the degree of disequilibrium. This might be expected of models explaining consumption incorporating lumpy durable expenditures where there is threshold which defines when tolerable disequilibrium becomes intolerable.

As Carruth and Dickerson (1997) point out, the prevalence of such asymmetries in consumer behaviour represents an omitted parameterisation in standard consumption functions and offers an alternative explanation to, for example, omitted variables for the poor performance of consumption functions in the post-financial liberalisation era.

CHAPTER 6

A MODIFIED RATIONAL EXPECTATIONS PERMANENT INCOME / LIFE CYCLE HYPOTHESES MODEL

6.1 Introduction

This Chapter aims to expand Hall's (1978) rational expectations permanent income/life cycle hypotheses (REPIH/RELCH) representative agent model to allow for current income consumers and explicitly account for the durable component of total expenditures. An extension to allow for intertemporal substitution is also given. The model is developed in logarithmic form. This logarithmic form helps guard against potential heteroscedasticity which may arise from *levels* estimation, facilitates the extension allowing for intertemporal substitution (see Hall 1988), and allows comparison with logarithmic error correction models (estimated in Chapter 5). Further, Campbell and Mankiw (1991) - CM hereafter - point out, "the process driving aggregate consumption and income seem to be log-linear rather than linear." (p. 729).[1] Directly producing the model in logarithms avoids the approximation involved when deriving models in *levels* and then estimating them in logarithms (see, Caballero 1994 for a recent example).[2] Further, the logarithmic REPIH/RELCH model removes two assumptions made in the levels derivation. First, there is no need to assume that the utility function's form is quadratic and additive over time, yielding marginal utility which is linear in consumption, which seems unrealistic. Secondly, the interest rate does not need to equal the discount rate to maintain the first difference form.

The derivation also seeks to ensure both the left and right hand sides of the model are stationary

[1] In general, empirical applications appear to prefer log-linear to level specifications. The former transformation generally yields I(1) aggregates so satisfying stationarity conditions in differenced formulations, which is not the case for level variables. "The log-linear model has an offsetting disadvantage, however, which is that the interpretation of our coefficient π in terms of the fraction of current income consumers can no longer be exact." (CM, p. 729, my italics).

[2] We have also conducted the simpler levels derivation, though do not detail it here.

and, more specifically, that each individual variable in the model is stationary.[3] If each variable is stationary this prevents the parameter estimates being influenced by nonstationary right hand side regressors forming stationary linear combinations, which normally provides an equilibrium interpretation.[4]

This derived model will be compared to those of Hall (1988), CM, Caballero (1994) and Jin (1994). Each of these authors have derived various components of the specification detailed here but none have produced a logarithmic formulation incorporating current income consumers, explicitly allowing for durability and extended for variable interest rates.

Two techniques for estimating this model are discussed: Generalised Method of Moments (GMM) with Newey and West (1987) adjusted standard errors and instrumental variables with moving average error terms (IV/MA). We also examine the instrument set to be used. Estimates of the proportion of current income consumers and assessment of the degree of intertemporal substitution and the impact of durability are provided for twenty OECD countries.

The Chapter is organised as follows. Section 6.2 derives a representative agent rational expectations model with a proportion of current income consumers whilst explicitly allowing for durability in logarithmic form. An extended version accommodating variable interest rates is also given. Previous researchers' findings for REPIH/RELCH style models are reviewed in Section 6.3. Section 6.4 discusses issues of estimation and analyses the empirical results. Summary and conclusions are given in section 6.5.

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[3] Stationarity ensures that the estimated parameters of interest are not subject to the problems of nonsense regression (associations between integrated but mutually independent variables) or spurious regression (unrelated series depending on other common factors), so causing the appearance of high correlations between unconnected series - see Doornik and Hendry 1995.

[4] In the empirical analysis, we may employ instruments which are nonstationary, in the belief they will form stationary linear combinations. This should not affect the interpretation of the structural parameters of the model but will help secure instrumented regressions of statistical significance.

6.2 Deriving A Logarithmic Form of Hall's (1978) Model Modified for Durability and Current Income Consumers

This section derives logarithmic REPIH/RELCH models for non-durable and durable consumption and, therefore, total expenditures. This is then modified to accommodate a proportion of current income consumers and an extension allowing for intertemporal substitution is given.

6.2.1 Deriving the Logarithmic Form of Hall's (1978) Model

Following Muellbauer (1983) we start from an additive homogeneous lifetime utility function which consumers are assumed to maximise from the current period, 0, to the end of their life, period T.

$$U = \sum_{i=0}^{T} [1/(1+\delta)^i](E_tC_{t+i}{}^{1-\beta}-1)/(1-\beta) \qquad \text{where, } \delta \geq 1,\ \beta > 0. \qquad \textbf{(6.2.1)}$$

This maximisation is subject to the following lifetime budget constraint on which income is earned over the working life; period M to N:

$$(1+r)A_{t-1} + \sum_{i=M}^{N} E_tY_{t+i}/(1+r)^i = \sum_{i=0}^{T} E_tC_{t+i}/(1+r)^i \qquad \textbf{(6.2.2)}$$

where E_tC_{t+i} is the consumption *expected* in period t by the representative agent, for period t+i, E_t being the expectation operator. The expectation may be based upon either point or stochastic income expectations - see Muellbauer (1994).[5] In empirical work the consumption measure

[5] Muellbauer (1983) calls this *planned* rather than *expected* consumption. This plan is based upon income expectations which Muellbauer (1983) argues will not yield a tractable solution if they are probabilistic because there would be a corresponding range of consumption plans, with corresponding probabilities, and not a single consumption plan. However, with point income expectations we have a single expected value of income in any particular future period and, therefore, a single planned level of consumption. It may be argued that even with stochastic income expectations there is still only a single planned level of consumption corresponding to the *expected* level of income and so a tractable solution could still be obtained. Having said this, we shall follow Muellbauer's (1983) assumption of point expectations in the following derivation. We use the *expectation* notation following CM.

typically used is per-capita non-durable expenditures. Use of non-durables ensures synchronisation between time of purchase and when utility is derived. Per-capita measures relate the model to a single consumer but does not overcome all problems associated with aggregation. Y_t is non-property income (labour income plus transfers) earned in period t.[6] r is the real rate of return on end of period assets, A_t, and is assumed constant. δ is a subjective discount factor which is at least zero in value to ensure current consumption bears at least as much utility as future consumption. β is the coefficient of relative risk aversion.

We set up the constrained optimisation problem to maximise **(6.2.1)** subject to **(6.2.2)**, using the following Lagrangian function, L:

$$L = \sum_{i=0}^{T} [1/(1+\delta)^i](E_tC_{t+i}{}^{1-\beta}-1)/(1-\beta) - \lambda_i\{[\sum_{i=0}^{T} E_tC_{t+i}/(1+r)^i] - [(1+r)A_{t-1} + \sum_{i=M}^{N} E_tY_{t+i}/(1+r)^i]\}. \quad \textbf{(6.2.3)}$$

Differentiating **(6.2.3)** with respect to C_t and E_tC_{t+1} and setting equal to zero yields:

$$\partial L/\partial C_t = [(1-\beta)C_t^{-\beta}]/(1-\beta) - \lambda_0 = C_t^{-\beta} - \lambda_0 = 0 \quad \textbf{(6.2.4a)}$$

$$\partial L/\partial E_tC_{t+1} = [1/(1+\delta)]E_tC_{t+1}{}^{-\beta} - [\lambda_1/(1+r)] = 0. \quad \textbf{(6.2.4b)}$$

Normalising **(6.2.4a)** and **(6.2.4b)** on λ_0 and λ_1 (the lagrange multipliers), respectively, gives:

$$\lambda_0 = C_t^{-\beta} \quad \textbf{(6.2.5a)}$$

$$\lambda_1 = [(1+r)/(1+\delta)]E_tC_{t+1}{}^{-\beta}. \quad \textbf{(6.2.5b)}$$

Setting **(6.2.5a)** equal to **(6.2.5b)** yields the *Euler equation* which equates the marginal utilities of the expected (planned) levels of consumption in periods t and t+1 ($\lambda_0=\lambda_1$). This is the equilibrium condition and defines the optimal allocation of consumption between adjacent

[6] Interest income is assumed to have already been paid on last period's assets, A_{t-1}, in the first part of the budget constraint **(6.2.2)**, as $(1+r)A_{t-1}$.

periods given the budget constraint. That is:

$$C_t^{-\beta} = [(1+r)/(1+\delta)]E_tC_{t+1}^{-\beta}. \tag{6.2.6}$$

Taking the natural logarithm of both sides of **(6.2.6)** gives:[†]

$$-\beta \ln C_t = \ln[(1+r)/(1+\delta)] - \beta \ln E_tC_{t+1} \tag{6.2.7}$$

which after re-arrangement yields:

$$\ln E_tC_{t+1} - \ln C_t = \{\ln[(1+r)/(1+\delta)]\}/\beta. \tag{6.2.8}$$

Assuming the real interest rate is constant and setting $\mu=\{\ln[(1+r)/(1+\delta)]\}/\beta$, **(6.2.8)** becomes:

$$\ln E_tC_{t+1} - \ln C_t = \mu. \tag{6.2.9}$$

The expected (planned) level of consumption in period t+1 only differs from its actual level if there is a *surprise* in expected income arising *after* the expectation was formed. Assuming this income innovation, ϵ_{t+1}, is stochastic means we can define actual consumption in period t+1 as:

$$\ln C_{t+1} = \ln E_tC_{t+1} + \epsilon_{t+1} \tag{6.2.10}$$

$$\text{or, } \ln E_tC_{t+1} = \ln C_{t+1} - \epsilon_{t+1} \tag{6.2.11}$$

Substituting **(6.2.11)** into **(6.2.9)** yields the logarithmic version of Hall's (1978) REPIH/RELCH equation:[7]

[7] Deriving the model in levels form, following Hall (1978), yields: $\Delta C_{t+1} = \epsilon_{t+1}$. This implies a model where the change in consumption is zero, which seems to be counterfactual because many industrial economies grow through time. Although an intercept can be secured by letting $r \neq \delta$ this would cause the coefficient on lagged consumption to deviate from unity, undermining the first difference form. This appears to be why an intercept is often included or consumption (growth) is detrended (demeaned).

[†] **(6.2.7)** is only valid if consumption is log-normally distributed, in which case **(6.2.7)** should have an extra constant term. μ in **(6.2.9)** would reflect this.

$$\Delta \ln C_{t+1} = \mu + \epsilon_{t+1}. \qquad \textbf{(6.2.12)}$$

(6.2.12) facilitates the test of the REPIH/RELCH. It suggests that the best prediction of next period's log of consumption is this period's log of consumption adjusted by some constant amount μ.[8] Model **(6.2.12)** rests on the following assumptions: (1) no credit restrictions or other non-linearities in the budget constraint; (2) no habits or adjustment costs (nondurable goods); (3) there are no measurement errors or transitory shocks to consumption; (4) the frequency of consumer decisions coincides with the data's periodicity; (5) the real interest rate is constant; (6) expectations are formulated "rationally" rather than using, say, contingent plans; (7) consumers plan between two adjacent periods which are not just the present and the future;[9] (8) there are no significant problems in applying a model appropriate for the individual representative agent to aggregate data.

If one did not assume constant real interest rates, relaxing assumption **(5)**, one would, following Hall (1988), obtain:[10]

$$\Delta \ln C_{t+1} = \mu^* + \sigma r_{t+1} + \epsilon_{t+1}^* \qquad \textbf{(6.2.12')}$$

Where $\sigma=1/\beta$, which may be interpreted as the intertemporal elasticity of substitution and

[8] Sometimes it is assumed that $\mu=0$, implying $\delta=r$. However, most economies grow through time so one would expect this constant to be positive, which implies, in this model, that $r>\delta$. This might be regarded unrealistic given the widely held belief that income uncertainty may create a risk premium causing $\delta>r$. Since we have not derived the model within this framework we do not necessarily draw such conclusions.

[9] Pemberton (1993) suggests the LCH implies a detailed plan of consumption for each future period which is argued to be unrealistic.

[10] Relaxing the constancy of the real interest rate in **(6.2.8)** yields: $\ln E_tC_{t+1} - \ln C_t = \{\ln[1/(1+\delta)]\}/\beta + (1/\beta)\ln(1+E_tr_{t+1})$. Substituting **(6.2.11)** into this expression gives Hall's (1988) intertemporal substitution model: $\Delta \ln C_{t+1} = \{\ln[1/(1+\delta)]\}/\beta + (1/\beta)\ln(1+E_tr_{t+1})$. Assuming $r_{t+1} = E_tr_{t+1}$ one obtains the form of model estimated by CM (without current income consumers). In this model a positive intercept naturally arises because $E_tr_{t+1} \neq \delta$. The real interest rate variable is measured as: $\ln(1+r_t) = \ln[1+(I_t/100)]-\Delta \ln P_t$, where I_t is the nominal interest rate.

$\mu^*=(1/\beta)\ln[1/(1+\delta)]$.[11] Strictly, the model is: $\Delta \ln C_{t+1} = \mu^* + (1/\beta)\ln(1+E_t r_{t+1}) + \epsilon_{t+1}^*$, but may be approximated by **(6.2.12')**.[12]

The implication is that if *both* the rational expectations and life cycle/permanent income hypotheses are true then all available information at time t (t-1) should be incorporated in C_t (C_{t-1}) and so information dated in period t (t-1) and earlier should have no explanatory power for consumption in period t+1 (t). There appears to be general agreement that lagged regressors when added to the right-hand-side of **(6.2.12)** are statistically significant, called orthogonality or exclusion tests, suggesting rejection of the REPIH/RELCH (see, for examples, Caballero 1994 and Gausden and Myers 1997). Thus, removal of the underlying assumptions is necessary. In the next sub-section the model will be modified to apply to durable expenditures - removing assumption **(2)**.

6.2.2 Applying the Logarithmic Form of Hall's (1978) Model to Durables

Caballero (1994) suggests that the implication of the REPIH/RELCH is more emphatically rejected for durable expenditures than nondurables. This suggests that models of total expenditures will need to account for durability. Adjustment costs are suggested to cause durable purchases to be "sporadic and lumpy rather than continuous and smooth" (Caballero 1994, p. 108).[13] Mankiw (1982) applies Hall's (1978) principal insight to the services from durables yielding a model where durable expenditures feature an MA(1) error process. Caballero (1994) extends Mankiw's (1982) specification to allow the spreading of replacement durable

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[11] Under the intertemporal substitution hypothesis, the coefficient on the real interest rate, σ, is expected to be positive. If, in period t, one expects that interest rates will rise (fall) over the period t to t+1 (relative to the period t-1 to t) the consumer will actively defer (raise) current consumption to the future period t+1: a higher real interest rate is associated with a greater change in consumption between periods t and t+1.

[12] The approximation used, that $x \approx \ln(1+x)$, is valid provided x is small, which may reasonably be assumed for real interest rates and the discount factor.

[13] Adjustment costs may cause microeconomic agents to tolerate *small* departures from an ever changing *optimal* level of the durable stock. Once departures are no longer considered *small* the consumer abruptly buys or sells to make the disequilibrium tolerable once more.

expenditures over several periods rather than one period. Thus, we follow Caballero (1994) by applying **(6.2.12)**, lagged by one period, to the aggregate stock of durables, K_t, and allowing the stock of durables to adjust slowly to *surprises*, thus:[14]

$$\Delta \ln K_t = \mu^D + \alpha\epsilon_t^D + (1-\alpha)\epsilon_{t-1}^D. \qquad \textbf{(6.2.13)}$$

Again following Caballero (1994), we assume the stock of durables depreciates geometrically at the rate, γ, and can be combined in the following expression with the flow of durables CD_t:[15]

$$K_t = (1-\gamma)K_{t-1} + CD_t. \qquad \textbf{(6.2.14)}$$

Normalising on CD_t gives:

$$CD_t = \Delta K_t + \gamma K_{t-1}. \qquad \textbf{(6.2.15)}$$

Following Bean (1981) and Young (1992) we assume that the stock of durables grows at the constant rate, g, implying:

$$K_t = (1+g)K_{t-1} \qquad \textbf{(6.2.16)}$$

or, $\Delta K_t = gK_{t-1}$. **(6.2.17)**

Substituting **(6.2.17)** into **(6.2.15)** gives:

$$CD_t = (\gamma+g)K_{t-1}. \qquad \textbf{(6.2.18)}$$

Taking the natural logarithm of both sides of **(6.2.18)** yields:

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[14] For a theoretical justification see Chapter 2 pp. 28-30.

[15] We use geometrically declining depreciation as an approximation following Caballero (1994).

$\ln CD_t = \ln(\gamma+g) + \ln K_{t-1}$. **(6.2.19)**

First differencing both sides gives:

$\Delta \ln CD_t = \Delta \ln K_{t-1}$. **(6.2.20)**

Substituting **(6.2.13)** into **(6.2.20)** gives the REPIH/RELCH model, in logarithmic form, for durables:

$\Delta \ln CD_t = \mu^D + \alpha\epsilon_{t-1}{}^D + (1-\alpha)\epsilon_{t-2}{}^D$ **(6.2.21a)**

or, for $v_t = \alpha\epsilon^D{}_t$,

$\Delta \ln CD_t = \mu^D + v_{t-1} + [(1-\alpha)/\alpha]v_{t-2}$. **(6.2.21b)**

Equation **(6.2.21b)** demonstrates that a moving average error process arises when the rational expectations hypothesis is applied to durable expenditures. This is the essential result of Mankiw (1982) and Caballero's (1994) extension. In contrast to Caballero's (1994) model, we find that the sum of the coefficients on the lagged error terms, $(1/\alpha)$, will likely be positive because $0<\alpha\leq 1$ - see **(6.2.13)**. However, following Caballero (1994), we allow a negative sum MA process as well.[16]

[16] The model derived by Caballero (1994) is in levels rather than logs and yields a specification including a current innovation: $\Delta CD_t = v_t + \{(1+\alpha\gamma-2\alpha)/\alpha\}v_{t-1} - [\{(1-\alpha)(1-\gamma)\}/\alpha]v_{t-2}$. To obtain a tractable solution in logarithms we depart from Caballero's (1994) derivation at equation **(6.2.16)**. It is the steps that follow which cause the omission of the current error in **(6.2.21b)**. The sum of the moving average terms in Caballero's (1994) model is $-\{1-(\gamma/\alpha)\}$. Caballero (1994) argues that provided adjustment costs are not excessively large, $\alpha > \gamma$, this sum will be negative. Caballero (1994) estimates his model with fifteen moving average error terms in logarithmic form and with an intercept, using quarterly data for three durable expenditure categories. The sums of the coefficients, in all three cases, are significantly negative, which is consistent with the slow adjustment interpretation ($\alpha > \gamma$). Of course, the derivation of his model is in levels and not logarithms, and we have found that, using the latter, this sum appears to be positive. It should also be noted that if $\alpha = \gamma$, the moving average error terms in Caballero's (1994) specification will sum to zero and if $\alpha < \gamma$, their sum will be positive. A larger value of

6.2.3 Developing the Logarithmic Form of Hall's (1978) Model for Total Expenditures

Total consumption may be expressed as:

$$TC_t = C_t^{\eta 1} CD_t^{\eta 2}. \qquad \textbf{(6.2.22)}$$

Although this is not the normal definition, provided the proportion of durable to non-durable expenditures is relatively constant through time the above may yield a reasonable approximation which facilitates a tractable solution.[17]

Indeed, **(6.2.22)** is analogous to the splitting of consumers into constrained and unconstrained, following, for examples, CM and Bayoumi and MacDonald (1995), amongst others, using a logarithmic form (see **(6.2.27a)** below). However, it is a less stringent constraint because it does not require that $\eta_1 + \eta_2 = 1$, whereas the total proportion of current income consumers must equal unity. Taking logs of both sides of **(6.2.22)** and differencing yields:

$$\Delta \ln TC_t = \eta_1 \Delta \ln C_t + \eta_2 \Delta \ln CD_t. \qquad \textbf{(6.2.23)}$$

Substituting **(6.2.12)**, lagged one period, and **(6.2.21b)** into **(6.2.23)** gives:

$$\Delta \ln TC_t = [\eta_1 \mu + \eta_2 \mu^D] + \eta_1 \epsilon_t + \eta_2 v_{t-1} + \eta_2 [(1-\alpha)/\alpha] v_{t-2}. \qquad \textbf{(6.2.24)}$$

Defining, $w_t = \eta_1 \epsilon_t$ and assuming the *surprises* to durable and non-durable expenditures are proportional, which implies $v_t = \rho w_t$, **(6.2.24)** becomes:[18]

γ could arise in lower frequency data because durables depreciate over a longer period, yielding the possibility of a non-negative sum MA process.

[17] Although the proportion of durable expenditures probably grows through time, if this growth is not too large over the estimation period the assumption of a constant proportion may provide a reasonable approximation. This approximation is avoided when one combines durable and non-durable expenditures in *levels*. An alternative means of considering total expenditure is to specify the utility function, **(6.2.1)**, as $U = \Sigma[(1+\delta)^{-i}/(1-\beta)][E_t(C_{t+i}^{1-\beta} + bK_{t+i}^{1-\beta})]$, where bK denotes the services from durables – see Favero (1993) p. 458, equation (4).

[18] We invoke this assumption to yield a tractable approximation.

$$\Delta \ln TC_t = [\eta_1\mu + \eta_2\mu^D] + w_t + \eta_2\rho w_{t-1} + \eta_2[(1-\alpha)/\alpha]\rho w_{t-2}. \qquad \textbf{(6.2.25)}$$

Defining $\kappa=[\eta_1\mu + \eta_2\mu^D]$ and extending the possibility of frictions due to durability, following Caballero (1994), we can re-write **(6.2.25)** as:[19]

$$\Delta \ln TC_t = \kappa + w_t + \sum_{i=1}^{L}\theta_i w_{t-i}, \qquad \textbf{(6.2.26)}$$

where, for L=2; $\theta_1 = \eta_2\rho$, $\theta_2 = \eta_2[(1-\alpha)/\alpha]\rho$, and $\Sigma\theta_i = \eta_2\rho/\alpha$.

The sign of $\Sigma\theta_i$ will depend upon the sign of ρ given that both η_2 (the durable *part* of total expenditures, see equation **(6.2.22)**) and α (the proportion of the surprise to durable expenditures occurring in the current period, see equation **(6.2.13)**) should be positive. Since we may expect shocks to durable and non-durable expenditures to be positively correlated ($\rho>0$) our prior belief is that $\Sigma\theta_i$ will be positive. A negative value of $\Sigma\theta_i$ is also possible.

When real interest rates are variable the model becomes,

$$\Delta \ln TC_t = \kappa^* + \sigma^* r_t + w_t^* + \Sigma\theta_i w_{t-i}^*, \qquad \textbf{(6.2.26')}$$

where $\sigma^* = (\eta_1+\eta_2)\sigma$ and $\kappa^* = -(\eta_1+\eta_2)\sigma\ln(1+\delta)$.

6.2.4 Modifying the Model to Allow for Current Income Consumers

Arguably one of the most popular explanations for the failure of Hall's Euler equation applied to non-durables is imperfections in capital markets. Hall (1978) suggested this could be accommodated by assuming that only unconstrained consumers followed the REPIH/RELCH while constrained consumers did not. A simple approximation for the latter's behaviour was that

[19] This is achieved by defining **(6.2.13)** as an MA process of order L-1. We therefore allow the data to indicate the number of periods it takes for the stock of durables to adjust to surprises.

they consumed all their income each period. Although this is objectionable because, for example, it precludes constrained consumers from saving (which is not implied by credit constraints), it appears to be a reasonable approximation which has been widely applied in the literature (see, for examples, Jappelli and Pagano 1989; CM; Jin 1994 and Bayoumi and MacDonald 1995). Another objection, which we are not aware has been raised previously, is that it assumes constrained consumers spend *all* their income on *non-durables*. We would be surprised if they did not spend a small proportion on durables such as televisions etc.. In this derivation, we remove this potential objection by assuming that some proportion of consumers, π, spend *all* their disposable income on *total expenditures*, while the unconstrained consumers follow the REPIH/RELCH as postulated by **(6.2.26)** above. This suggests that this proportion may be larger when applied to total consumption relative to non-durables. Following the literature we specify total consumption as the weighted sum of unconstrained (U) and constrained (C) consumers' expenditures, thus:

$$\ln TC_t = (1-\pi)\ln TC_t^U + \pi \ln TC_t^C, \qquad \textbf{(6.2.27a)}$$

implying $[TC_t = (TC_t^U)^{(1-\pi)}.(TC_t^C)^{\pi}]$. Although not strictly a definition it is a way of yielding a tractable solution and is widely employed (see, for examples, CM and Bayoumi and MacDonald 1995). However, as Jin (1994), footnote 5 p. 4, points out, the interpretation of π is the proportion of *expenditure consumed* by current income consumers rather than the proportion of disposable income accruing to them (which we denote as π_Y). **(6.2.27a)** may also be written in difference form as:

$$\Delta\ln TC_t = (1-\pi)\Delta\ln TC_t^U + \pi\Delta\ln TC_t^C. \qquad \textbf{(6.2.27b)}$$

Assuming constrained agents consume all their income:

$$\ln TC_t^C = \ln Y_t^C \qquad \textbf{(6.2.28)}$$

where their share of total income may expressed as:

$$\ln Y_t^C = \ln(\pi_Y Y_t) = \ln\pi_Y + \ln Y_t \qquad \textbf{(6.2.29)}$$

or, in difference form, as:

$$\Delta\ln Y_t^C = \Delta\ln Y_t. \qquad \textbf{(6.2.30)}$$

It is because differencing removes π_Y from **(6.2.30)** that the proportion of income accruing to current income consumers cannot be identified. Substituting **(6.2.30)** into the differenced form of **(6.2.28)** gives:

$$\Delta\ln TC_t^C = \Delta\ln Y_t. \qquad \textbf{(6.2.31)}$$

Even recognising that only the proportion, $(1-\pi_Y)$, of total income accrues to unconstrained consumers, it becomes clear that the growth of unconstrained consumers' total consumption equals the growth of aggregate total consumption - again π_Y cannot be identified. That is:

$$\Delta\ln TC_t^U = \ln(1-\pi_Y) - \ln(1-\pi_Y) + \Delta\ln TC_t = \Delta\ln TC_t, \qquad \textbf{(6.2.32)}$$

or, after substitution of **(6.2.26)** into **(6.2.32)**:

$$\Delta\ln TC_t^U = \kappa + w_t + \Sigma\theta_i w_{t-i} \qquad \textbf{(6.2.33)}$$

Substituting **(6.2.31)** and **(6.2.33)** into **(6.27b)**, we obtain:

$$\Delta\ln TC_t = \kappa' + \pi\Delta\ln Y_t + \tau_t + \sum_{i=1}^{L}\theta_i\tau_{t-i} \qquad (6.2.34)$$

where $\kappa' = (1-\pi)(\eta_1+\eta_2)\sigma\ln[(1+r)/(1+\delta)]$ and $\tau_t = (1-\pi)w_t$.

For variable real interest rates, **(6.2.34)** becomes:

$$\Delta \ln TC_t = \kappa'^* + \pi \Delta \ln Y_t + \sigma'^* r_t + \tau_t'^* + \Sigma \theta_i \tau_{t-i}'^* \qquad \textbf{(6.2.34')}$$

where, $\sigma'^* = [(1-\pi)(\eta_1+\eta_2)\sigma]$ and $\kappa'^* = -(1-\pi)(\eta_1+\eta_2)\sigma \ln(1+\delta)$.

The parameter π represents that *part* of *expenditure* commanded by current income consumers. As CM recognise, the logarithmic derivation means this no longer has the interpretation of the *proportion* of current income consumers. However, this parameter is often presented with this interpretation. CM, footnote 8 p. 731, suggest that the results of such models in *level* and *logarithmic* forms yield similar estimates of this coefficient for USA data, so this interpretation is approximately valid. Jin (1994) derives a *near* logarithmic form of this model in which the proportion of *income* accruing to current income consumers can be identified. As suggested above, **(6.2.34)** and CM-style models estimate that part of *expenditure* consumed by current income consumers.

The additional consideration of variable interest rates in equation **(6.2.34')** extends the model to account for intertemporal substitution. We expect the coefficient on interest rates to be positive.

The models **(6.2.34)** and **(6.2.34')** achieve the aims set out at the beginning: being in complete logarithmic form, explicitly accounting for durability and current income consumers (with an extension to allow for intertemporal substitution). Being applied to total rather than non-durable expenditures ensures that any durable expenditures made by current income consumers are not ignored. This may raise the estimate of π relative to applications excluding the durable component of expenditures.

6.3 Previous Researchers' REPIH/RELCH Findings

In this section we review a selection of previous researchers' findings on the REPIH/RELCH and, therefore, motivate various specifications nested within equation **(6.2.34')**. The main models' results to be reviewed are those of Hall (1978), Hall (1988), Campbell and Mankiw (1991), Caballero (1994) and Jin (1994).

6.3.1 Hall's (1978) Model

Hall (1978) originally derived the random walk implication of the joint hypotheses of the REPIH/RELCH for nondurables (and services from durables). The model was in levels, included an intercept and allowed the coefficient on lagged income to deviate from unity. These models provide a test for the REPIH/RELCH: if information available to consumers at the time of forming their consumption decision (t-1 or earlier) can be added with statistical significance to the random walk model then the joint hypothesis is rejected. A criticism aimed at the early applications of Hall's (1978) test equation was the use of nonstationary regressors (see, for example, Flavin 1981). Consumption and income have generally been found to be difference rather than trend stationary, which is confirmed for the data set under study in Chapter three. Thus, recent work tends to focus upon REPIH/RELCH models in difference form, such as equation **(6.2.12)** above.[20]

The empirical literature reviewed in Chapter two suggests a general rejection for Hall's (1978) REPIH/RELCH model because current consumption is excessively sensitive to lagged, and therefore predictable, information. Anticipating the results of this Chapter we find evidence that Hall's (1978) implication of the REPIH/RELCH is rejected for all twenty OECD countries because instrumented income growth (embodying lagged information) is a statistically significant determinant of total consumption expenditures. This is consistent with Caballero's (1994) assessment that "Researchers now seem to agree that *Hall's (1978)* implication of the PIH does not hold in the data, regardless of the country and sample used." (p. 107, my italics). Subsequent research has focused upon the modification Hall's (1978) random walk model by removing one or more of the assumptions upon which it is based.[21]

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[20] In the context of testing for excess sensitivity Stock and West (1988) argue that there is no bias if all nonstationary regressors are cointegrated.

[21] The main modifications considered in the literature (and here) are the role of durables, a proportion of current income consumers and variable interest rates. An interesting avenue not considered here is the role of transitory consumption, which is highlighted by, for example, Falk and Lee (1998). They suggest that the finding of excess sensitivity may be due to the failure to account for transitory consumption.

6.3.2 Caballero's (1994) Model

Caballero (1994) argues that according to Hall's (1978) insight the growth rate of the stock of durables should be unpredictable but notes that "The rejection of the basic theory for durable goods is an order of magnitude larger than for non-durables." (Caballero 1994, p. 107). This motivates an analysis of this *larger* rejection. Mankiw (1982) is cited as demonstrating that the combination of Hall's (1978) random walk specification with the perpetual inventory expression for the accumulation of the new durable stock yields a model where the change in the *level* of consumption follows a first order moving average (MA) process. Using post-war quarterly US data, Mankiw (1982) finds that the MA coefficient is not significantly different from zero. Since this modified REPIH/RELCH model implies an MA process for durables, the hypothesis is rejected. Caballero (1994) confirms this result for US quarterly data on furniture, automobile and total durable expenditures. The moving average (MA) error is found to be insignificant for all three categories of durables when the *growth rate* of durables is regressed upon an intercept and moving average error.[22]

However, when he applies this test to his extended model, which assumes that the aggregate stock of consumption adjusts slowly to innovations yielding an expression where durable expenditures follow an MA(q) process, the REPIH/RELCH cannot be rejected. That is, Caballero (1994) finds for all three data sets, using an MA(15), that the sum of these MA terms is negative and that they are statistically significant.[23] It is argued that these results are consistent with a slowness of adjustment REPIH/RELCH specification and that models which only consider an MA(1) process may erroneously reject the REPIH/RELCH.

[22] The model derived by Caballero (1994) is in *levels* and excludes an intercept.

[23] According to the Box and Jenkins method of ARIMA model identification the number of statistically significant autocorrelation coefficients indicates the order of MA process. Further, if more than the first four consecutive autocorrelation coefficients are statistically significant this is typically considered indicative of a nonstationary process. *If*, therefore, 15 MA error terms implies that the first 15 autocorrelation coefficients are statistically significant, this *might* suggest the process is nonstationary. Thus, one *may* wish to view Caballero's (1994) results with caution.

6.3.3 Hall's (1988) Model

Hall (1988) argues that if real interest rates are not assumed constant then the random walk implication of the REPIH/RELCH generalises to include this variable as a regressor. This assumes the joint lognormality and homoscedasticity of consumption and real interest rates and that expectations are formed rationally. The essential form of Hall's (1988) model may be characterised by **(6.2.12')**, above. Hall (1988) argues that by estimating the parameters of the representative agent's utility function rather than the coefficients of a consumption/saving function one is more likely to avoid the problems associated with the Lucas (1976) critique. This is because individuals always seek to maximise the same utility function whereas consumption/saving functions may be unstable in the face of policy regime changes. Hall (1988) further argues that accounting for problems of time aggregation will ensure that the estimate of the elasticity of intertemporal substitution, σ, is robust.[24]

Hahm (1998) suggests that although Mankiw (1981), Hansen and Singleton (1983) and Mankiw (1985) all find that σ is positive, only Hansen and Singleton's (1983) study shows it is significant at the 5% level. However, the underlying model is rejected in all the studies. Further, none of these studies account for the problem of time aggregation bias because they instrument interest rates with variables dated in period t-1, which will lead to biased estimates. Hall (1988) shows that when appropriate instruments, dated in period t-2 or earlier, are used, the estimates of σ become negative. It is concluded that intertemporal substitution is not supported when σ is correctly estimated.

Hahm (1998) argues that this evidence against intertemporal substitution may be due to using the incorrect measure of consumption and excluding instruments. Regarding the former, the aggregation of housing services with other expenditures is highlighted as potentially causing misleading results: one needs to be careful with the treatment of (services from) durables. With reference to the latter, Hahm (1998) recommends extending the instrument set to include

[24] It is recognised that the elasticity of substitution, σ, may be interpreted as the reciprocal of the coefficient of relative risk aversion, $1/\beta$, although Hall (1988) argues for the former interpretation.

variables lagged two to four periods. It is suggested that Hall's (1988) results against intertemporal substitution may be due to only using the second lags of variables as instruments. Hahm (1998) finds evidence favouring the presence of both current income consumers and intertemporal substitution for non-durable US consumption. When these models were estimated using non-durables and services as the measure of consumption, no significant relationship between consumption growth and the interest rate is revealed. This suggests the need to accommodate consumer expenditure series including durable components, especially housing services.

6.3.4 Campbell and Mankiw's (1991) Model

Following Hall's (1978) suggestion, CM remove the assumption of perfect capital markets by considering two types of consumer: those *unconstrained* consumers who follow the REPIH/RELCH and current income consumers who consume all their income each period. They focus on the log-linear form for non-durable expenditures (CM's equation (19)) which is reproduced below:

$$\Delta \ln C_t = \mu + \pi \Delta \ln Y_t + \epsilon_t. \qquad \textbf{(6.3.1)}$$

CM estimate **(6.3.1)**, with and without interest rates added, though reject a role for intertemporal substitution. They find that the estimated proportion of current income consumers is significant and between zero and one for Canada, France, Sweden, the UK and the US. They further note that countries with larger values feature less well developed consumer credit markets which is consistent with these estimates representing the proportion of liquidity constrained consumers. Jappelli and Pagano (1989) provide similar support for the liquidity constraint model for Greece, Italy, Japan, Spain, Sweden, the UK and the US. CM find no evidence that the proportion of current income consumers varies through time, except for the UK. Bacchetta and Gerlach (1997) find evidence for significant time-varying parameters for a CM-type model for the UK and the USA. However, no evidence of time variation is found for Canada, France or Japan.

Overall, there is strong evidence supporting the presence of a proportion of current income consumers and that this proportion, in general, varies across countries with the tightness of liquidity constraints. Whether the proportion of current income consumers varies through time for any particular country is an unresolved issue.

6.3.5 Jin's (1994) Model

Using time-series and pooled estimates for nineteen OECD countries, Jin (1994) finds evidence supporting a significant proportion of current income consumers. This proportion is found to systematically vary with proxies for liquidity constraints across countries. However, Jin's (1994) model, which features a semi-logarithmic nonlinear functional form (see Chapter 2, p. 39), can be criticised for a number of reasons. Firstly, it only implicitly accounts for durability through the use of Newey and West (1987) standard errors which are consistent in the face of, at most, a first-order MA process. Second, there are only twenty-five observations available for the time-series estimates. Thirdly, the measure of income used for inference for all nineteen countries is only an approximation to disposable income.[25] Fourthly, the functional form of his model is semi-logarithmic and may be subject to heteroscedasticity, although the Newey and West (1987) standard errors are robust to a non constant error variance. Fifthly, the regressors in these specifications are individually nonstationary because they are functions of the reciprocal of the APC. Results reported in Chapter three and Sarantis and Stewart (1998a) suggest that the APC in our twenty countries, which include eighteen of those considered by Jin (1994), is nonstationary.[26] Our model and method overcomes all these potential criticisms.

We also note that Jin (1994) estimates these models with GMM, rather than nonlinear IV, arguing this will likely yield more efficient estimates. However, Hamilton (1994) p. 409

[25] Jin's (1994) approximate measure of private disposable income, used for all nineteen countries, is national disposable income minus government consumption expenditure. This fails to account for the general government component of government income.

[26] The nonstationary regressors may form a stationary linear combination ensuring the right hand side is stationary. However, if the parameters are chosen to ensure the right hand side of the equation is stationary, one would question whether the estimated coefficients maintain their intended interpretation and, therefore, the validity of the inferences drawn. Preferably individually stationary variables should be included in the structural part of the model.

suggests that "The key advantage of GMM is that it requires specification only of certain moment conditions rather than the full density. This can also be a drawback, in that GMM does not always make efficient use of all the information in the sample." We use both GMM and IV to estimate our linear specifications.

6.4 Empirical Issues and Econometric Results

We seek to obtain reliable insights into the role of current income consumers for twenty OECD countries. We propose estimating a REPIH/RELCH model in stationary logarithmic form, with income growth and a moving average error process to evaluate the importance of current income consumers and durability in total consumer expenditure.

No previous study considers a model incorporating all of these factors. For example, Caballero's (1994) model is derived in levels, does not account for current income consumers and is strictly appropriate for completely durable expenditures while CM's specification does not account for durability. Although Jin's (1994) model enables the proportion of income accruing to current income consumers to be directly estimated, its functional form is semi-logarithmic, includes nonstationary regressors and only implicitly allows for durability in an *ad hoc* manner by employing Newey and West (1987) coefficient standard errors. Further, Jin (1994) uses an approximate measure of private disposable income and the sample size employed for the time series regressions is only twenty five observations.

Following Hall (1988), the interest rate is added to **(6.2.34)**, giving **(6.2.34')**, to consider intertemporal substitution. We view the investigation of current income consumers and durability (and variable interest rates) in a reliable specification as first steps in analysing the rejection of the REPIH/RELCH for a broad range of OECD countries.

6.4.1 Does the Presence of Durability Explain the Rejection of the REPIH/RELCH?

In this section we estimate equation **(6.2.26)** to help determine whether durability can explain the rejection of the REPIH/RELCH. We allow zero to five moving average (MA) error terms in the model, which is in line with Caballero (1994) who specified an MA process to allow for approximately four years of frictions. Specification of more than five MA terms implies, following the Box and Jenkins (1970) methodology, that more than the first five autocorrelation coefficients in the autocorrelation function (ACF) are significant, suggesting the model is inconsistent with a stationary process. Setting an MA(5) model as a maximum specification also helps prevent overfitting the sample. Given such moving average models are nested within the ARIMA modelling framework we utilise the identification and diagnostic checking procedures outlined by Box and Jenkins (1970) - without attempting to employ an autoregressive process.

The Box and Jenkins procedure for identifying an ARIMA model involves setting the (maximum) order of moving average process equal to number of consecutive significant autocorrelation coefficients in the ACF. We bear in mind that the small sample size may distort accurate identification and that Caballero's (1994) insight of slow adjustment potentially causes the first MA term to be insignificant with higher order terms significant. Therefore, we do not apply this identification rule mechanically. Nevertheless, the series should exhibit significant autocorrelation according to the individual autocorrelation coefficients and the Ljung-Box Q-statistic (LB).[27] We also examine the estimated models and consider whether the moving average terms are individually and/or jointly significant according to t-ratios and an F-test [denoted $F(R^2)$], respectively. Schwartz's Bayesian Information Criterion (SBIC), which penalises overparameterised models, is reported to help guard against overfitting and represents our main model selection criteria. In addition, we conduct diagnostic checks for residual autocorrelation and invertibility.

Table 6.1 reports the first five autocorrelation coefficients [denoted r(i), i=1,...,5] of the ACF and the LB statistic for two and six lags [LB(2) and LB(6)] for each country's growth rate of

[27] In the present application the individual autocorrelation coefficients should exceed (approximately) 0.345 in magnitude to be significant. The Ljung-Box statistic is distributed as a $\chi^2(k\text{-}q)$, where k is the number of autocorrelation coefficients employed in the test and q is the order of moving average process.

TABLE 6.1: Box-Jenkins Identification of MA Model with ACF

Country→	Aul	Aut	Bel	Can	Den	Fin	Fra	Ger	Gre	Ice
r(1)	0.147	0.168	**0.374**	**0.459**	0.173	**0.512**	**0.592**	**0.551**	**0.555**	0.252
r(2)	-0.296	0.170	0.181	0.258	-0.096	0.162	**0.403**	0.052	**0.427**	-0.235
r(3)	-0.058	0.230	0.288	0.025	0.145	0.146	**0.437**	-0.099	**0.531**	-0.196
r(4)	0.194	0.194	-0.048	-0.074	-0.205	-0.075	0.283	-0.025	0.339	-0.064
r(5)	0.162	-0.102	-0.028	-0.127	-0.333	-0.152	0.179	-0.019	0.311	0.003
LB(2)	4.269	2.201	**6.622**	**10.663**	1.503	**10.989**	**19.719**	**11.689**	**18.858**	4.598
LB(6)	7.332	8.645	10.133	**14.236**	8.910	**13.427**	**36.352**	12.137	**45.672**	6.824
BJ(q=)	0	0	1	1	0	1	3	1	3	0

Country→	Ire	Ita	Jap	Net	Nor	Spa	Swe	Swz	UK	USA
r(1)	0.268	**0.421**	**0.551**	**0.680**	0.213	**0.388**	**0.452**	**0.605**	**0.383**	**0.356**
r(2)	-0.088	-0.012	**0.459**	**0.430**	-0.031	0.332	0.189	0.209	-0.086	-0.091
r(3)	-0.286	0.152	**0.494**	0.243	-0.173	0.137	0.071	0.098	-0.239	-0.158
r(4)	-0.088	-0.019	**0.377**	0.181	-0.143	0.196	-0.153	-0.037	-0.323	-0.106
r(5)	0.016	0.162	0.226	0.241	-0.015	0.018	-0.234	-0.136	-0.109	-0.120
LB(2)	3.041	**6.742**	**19.806**	**24.873**	1.766	**10.060**	**9.175**	**15.656**	5.876	5.144
LB(6)	7.037	9.718	**46.654**	**32.811**	4.094	12.499	**14.980**	**17.067**	**13.323**	7.584
BJ(q=)	0	1	4	2	0	1	1	1	1	1

Table 6.1 notes. The autocorrelation function (ACF) for the actual data is given in the rows r(1),..., r(5) below the heading ACF, along with the Ljung-Box statistics for second [LB(2)] and sixth [LB(6)] order autocorrelation. The 5% critical value for the autocorrelation coefficients, r(i), is ±0.345 and for LB(2) and LB(6) are 5.99 and 12.59, respectively. BJ(q=) denotes the order of moving average (MA) process which would be identified using the Box-Jenkins procedure. The reported ACF for Germany is for the residual series from regressing consumption growth on an intercept and a dummy variable taking a unit value in 1991 and zero otherwise. Bold emphasis indicates a significant test statistic.

consumption.[28] BJ(q=) denotes the order of MA process identified according to the Box-Jenkins procedure. For six of the twenty countries (Australia, Austria, Denmark, Iceland, Ireland and Norway) none of these autocorrelation coefficients are statistically significant suggesting no moving average terms. An MA(1) is indicated for ten of the countries (Belgium, Canada, France, Germany, Italy, Spain, Sweden, Switzerland, the UK and the USA). For the Netherlands an

[28] The ACF for Germany is for the residuals of consumption growth after being regressed upon a constant and a dummy variable taking the value of unity in 1991 and zero otherwise.

MA(2) is suggested, while Finland and Greece appear to be MA(3) processes and Japan an MA(4). For none of the countries is an MA(5) indicated according to the Box and Jenkins procedure.

The alternative *objective method* of identification involves selecting the preferred MA model as that which minimises the SBIC. We also require it to satisfy the necessary condition for invertibility and to be free from residual autocorrelation. When no model satisfies both diagnostic checks for any particular country we favour the model with the lowest SBIC which is invertible. This is because our primary concern is with the coefficients on the moving average terms, which we seek to ensure take on plausible values.

We report the estimated models using this objective method in Table 6.2.[29] This table reports the coefficients of the MA terms [denoted MA(i), i=1,...,5] with associated t-ratios in brackets below. Also reported are the sum of the MA terms [ΣMA], the adjusted R^2 [Adj R^2], joint significance of the regression [$F(R^2)$], the SBIC, residual autocorrelation coefficients and LB test for residual autocorrelation. No moving average terms are present in the favoured models for Austria, Ireland and Norway. An MA(1) is selected for thirteen of the twenty countries (Belgium, Denmark, Finland, France, Germany, Greece, Iceland, Japan, the Netherlands, Spain, Sweden, Switzerland and the USA). Neither an MA(2) nor MA(3) model is preferred for any country. An MA(4) is chosen for Australia and the UK and an MA(5) for Canada and Italy. It is interesting to note that for fifteen of the seventeen countries where a significant MA process exists, the sum of these terms is positive. Further, all models (not reported) identified using the Box-Jenkins method yielded positive sum moving average parameters. This appears to confirm our prior belief that shocks to non-durable and durable expenditures are positively correlated ($\rho>0$). However, there is evident autocorrelation at the five percent (but not one percent) level for Finland, Greece and Spain, suggesting these countries' results should be treated with caution.[30] Serial correlation is significant at both the one and five percent levels for Japan

[29] The nonlinear iterative estimation procedure available in Eviews 2.0 is employed for the estimation of all MA models.

[30] The second residual autocorrelation coefficient is also significant for the Netherlands.

TABLE 6.2: Estimated (MA) Modified REPIH/RELCH Models, Equation (6.2.26)

Country→	Aul	Aut	Bel	Can	Den	Fin	Fra	Ger*	Gre	Ice
MA(1)	0.068		0.603	0.356	0.177	0.777	0.550	0.859	0.678	0.412
	(0.553)		(4.437)	(2.158)	(0.991)	(7.169)	(3.768)	(10.878)	(5.382)	(2.598)
MA(2)	-0.502				0.042					
	(-5.194)				(0.278)					
MA(3)	-0.090				0.707					
	(-0.707)				(8.678)					
MA(4)	0.746				-0.401					
	(6.756)				(-2.286)					
MA(5)					-0.276					
					(-1.537)					
ΣMA	0.223	0.000	0.603	0.356	0.249	0.777	0.550	0.859	0.678	0.412
Adj R^2	0.323	0.000	0.150	0.134	0.351	0.339	0.276	0.668	0.268	0.093
$F(R^2)$	5.062	0.000	6.983	6.266	4.674	18.400	13.940	35.184	13.470	4.484
SBIC	-8.604	-7.949	-7.943	-7.326	-7.092	-7.062	-8.503	-8.247	-7.478	-5.495
Residuals↓										
$r_w(1)$	0.031	0.168	-0.143	0.090	0.807	-0.048	0.108	0.078	-0.041	-0.049
$r_w(2)$	0.145	0.170	0.090	0.252	0.835	0.102	0.283	0.099	0.279	-0.156
$LB_w(6-q)$	3.371(2)	8.645(6)	5.833(5)	5.930(5)	2.559(1)	2.115(5)	**11.492(5)**	5.396(5)	**11.391(5)**	2.367(5)

Country→	Ire	Ita	Jap	Net	Nor	Spa	Swe	Swz	UK	USA
MA(1)		0.513	0.466	0.646		0.511	0.456	0.678	0.346	0.439
		(7.328)	(3.103)	(4.762)		(4.272)	(2.956)	(5.309)	(2.101)	(2.785)
MA(2)		-0.025							-0.442	
		(-0.314)							(-2.893)	
MA(3)		0.218							-0.353	
		(3.191)							(-2.352)	
MA(4)		-0.321							-0.513	
		(-4.683)							(-3.109)	
MA(5)		-0.841								
		(-12.33)								
ΣMA	0.000	-0.455	0.466	0.646	0.000	0.511	0.456	0.678	-0.962	0.439
Adj R^2	0.000	0.558	0.214	0.445	0.000	0.149	0.166	0.383	0.381	0.141
$F(R^2)$	0.000	9.585	10.284	20.644	0.000	6.953	7.786	22.075	6.237	6.564
SBIC	-6.874	-7.882	-7.154	-7.945	-7.151	-6.788	-13.078	-8.401	-7.678	-8.162
Residuals↓										
$r_w(1)$	0.268	0.222	0.076	0.150	0.213	-0.178	0.029	0.162	0.096	0.005
$r_w(2)$	-0.088	-0.020	0.354	0.368	-0.031	0.471	0.152	0.168	0.030	-0.045
$LB_w(6-q)$	7.037(6)	3.541(1)	**18.94(5)**	9.021(5)	4.094(6)	**13.91(5)**	3.572(5)	4.491(5)	1.889(2)	1.388(5)

Table 6.2 notes. The German model includes a dummy variable taking a unit value in 1991 and zero otherwise to account for the impact of reunification. This may be thought of as an ARIMAX model. The coefficients of estimated moving average terms are given in the rows denoted MA(i) with estimated t-ratios specified in brackets below. ΣMA denotes the sum of estimated moving average coefficients, Adj R^2 is the adjusted coefficient of determination, $F(R^2)$ is the F-statistic for the significance of the regression [5% critical values for q=1,..., 5 are 4.14, 3.30, 2.91, 2.69 and 2.54, respectively] and SBIC is Schwartz's Bayesian Information Criterion. The first [$r_W(1)$] and second [$r_W(2)$] autocorrelation coefficients along with the Ljung-Box [$LB_W(6\text{-}q)$] statistic for the residuals of the model are also given. The 5% critical values for the autocorrelation coefficients is ±0.345. Critical values for the Ljung-Box statistics vary with q. At the 5% (1%) level the critical values are 11.07 (15.09), 9.49 (13.28), 7.81 (11.34), 5.99 (9.21) and 3.84 (6.63), for q=1,...,5, respectively. Bold emphasis indicates a significant autocorrelation coefficient and Ljung-Box statistic and an insignificant parameter and regression (depending upon context). All models include an intercept. Estimates of intercepts are not reported to save space.

indicating that this model is misspecified.

In summary, only four countries feature a moving average process of order greater than one, regardless of the model selection criteria used. Thus, durability does not, in general, seem to induce more than an MA(1) error process in annual total expenditure data. However, the REPIH/RELCH applied to consumption including a significant component of durables is rejected when no moving average process is present. This hypothesis is rejected for six countries according to the Box-Jenkins procedure and for three countries using the objective method. Evidence of autocorrelation for an additional four countries suggests that the REPIH/RELCH is rejected for *up to* ten countries (Australia, Austria, Denmark, Finland, Greece, Iceland, Ireland, Japan, Norway and Spain).

6.4.2 Generalised Method of Moments (GMM) Estimation Results

In this sub-section we estimate the REPIH/RELCH models which allow for durability and variable interest rates, **(6.2.26')**, durability and current income consumers, **(6.2.34)**, and durability, variable interest rates and current income consumers, **(6.2.34')**, using the GMM estimator.

The GMM estimator seeks to ensure consistent parameter estimation by estimating coefficients that make the correlations between each of the instruments and the regression's error zero (with exact identification) or as near to zero as possible (with overidentification). Our regressions are all overidentified because there are more instruments (zero-correlation conditions) than parameters to be estimated. We choose the weighting matrix, required with overidentification, following Newey and West (1987), which provides heteroscedasticity and autocorrelation consistent (HAC) estimates. That is, both parameter and standard error estimates are robust to the presence of heteroscedasticity and autocorrelation. Thus, we *implicitly* account for the durable component of total consumption expenditures because we have characterised this as an MA error process. Strictly, the HAC standard errors we use are only consistent in the face of an MA(1) process. However, since the results of model **(6.2.26)** presented in Table 6.2 suggest that, in general, no more than an MA(1) process is indicated, our inference using GMM should be valid.[31] Further, as Pindyck and Rubinfeld (1997) point out, the GMM estimator "does not require the assumption of normality" (Pindyck and Rubinfeld 1997, p. 293).

To implement this method we need to decide which instruments to use. The primary criteria for the validity of instruments are that they should be uncorrelated with the (structural) equation's disturbance term *and* be highly correlated with the variable to be instrumented. In REPIH/RELCH models instruments cannot be dated earlier than t-2 to satisfy the first condition - see Hall (1988) and Campbell and Mankiw (1991).[32] However, it can become difficult to obtain

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[31] To estimate the linear model, $\mathbf{y} = \mathbf{X}\boldsymbol{\beta} + \mathbf{u}$, using GMM one aims to ensure the orthogonality (zero-correlation) conditions, $\mathbf{W}'(\mathbf{y}-\mathbf{X}\boldsymbol{\beta})=0$, hold, where $\mathbf{W}$ is a matrix of valid instruments. With overidentification $\mathbf{W}'(\mathbf{y}-\mathbf{X}\boldsymbol{\beta})\neq 0$, so we seek to minimise the squared deviations of these orthogonality conditions, subject to a weighting matrix, $\mathbf{A}$, where, $\mathbf{A}$, is the inverse of the variance/covariance matrix of $\mathbf{W}'(\mathbf{y}-\mathbf{X}\boldsymbol{\beta})$. That is, we solve: $\partial\{(\mathbf{y}-\mathbf{X}\boldsymbol{\beta})'\mathbf{WAW}'(\mathbf{y}-\mathbf{X}\boldsymbol{\beta})\}/\partial\boldsymbol{\beta} = 0$; which yields the GMM estimator: $\boldsymbol{\beta} = (\mathbf{X}'\mathbf{WAW}'\mathbf{X})^{-1}\mathbf{X}'\mathbf{WAW}'\mathbf{y}$. It should be noted that the value of the weighting matrix, $\mathbf{A}$, depends upon the parameters, $\boldsymbol{\beta}$, which in turn depend upon, $\mathbf{A}$. Thus, the estimation procedure iterates from the initially specified $\mathbf{A}$. As suggested in the text, we use the Newey and West (1987) weighting matrix (fixed bandwidth [=3]) with Bartlett kernel. For further details at various levels of technicality see, Hamilton (1994), Newey and West (1987), Pindyck and Rubinfeld (1997) and the Eviews 2.0 User's guide.

[32] Instruments should be dated in period t-2 or earlier, for the following reasons. Firstly, consumption and income are specified in particular time periods in the model but the data used is time averaged which can produce spurious first, but not higher, order autocorrelation (and cross correlations). Secondly, if there is white noise measurement error in the levels of consumption and income or if taste shocks create white noise transitory consumption, this will cause a first order moving average error process. Thirdly, goods measured as nondurables can incorporate some degree of durability, especially semidurables like clothing, which will yield a moving average error. This result will be exacerbated when durables are included as well.

instruments which satisfy the second condition if variables dated in period t-1 cannot be used, particularly for income growth. We focus upon the choice of instruments which yield a significant instrumented regression, especially for income growth.

We need to decide the variables to be used as instruments and their lag length. Although Hahm (1998) warns against the perils of using too few lags, this is with reference to quarterly data. We found that variables dated in period t-2 were generally sufficient for the annual data employed here.[33]

The variables chosen as instruments, for both income growth and interest rates, are guided by previous studies and data availability. Following CM we start with lags of consumption and income growth as well the consumption-income ratio. CM and subsequent researchers have found the consumption-income ratio to be a particularly important instrument for income growth, perhaps reflecting the relevance of long run information. However, its significance may appear surprising because we have found it to be nonstationary. We therefore consider the separate inclusion of the logs of consumption and income, allowing for the possibility that they form a stationary linear combination, being a more appropriate explanation of the stationary income growth variable. We also consider the level and changes of inflation ($\Delta \ln P_t$), unemployment, (U_t), and the real interest rate.

After testing various combinations of instruments we find that, in general, the separate inclusion of the logs of consumption and income substantially enhance the fit of, especially, the instrumented income growth equation. We are able to obtain significant instrumented equations, at the five percent level, for all twenty countries, for both income growth and interest rates.

Fourthly, publication delays of data on aggregate consumption and income statistics may mean consumers know their own consumption and income last period but not their aggregate counterparts, making these aggregates invalid instruments.

[33] Lags dated in periods t-3 and t-4 are superfluous when incorporated *with* those dated in period t-2. However, lags in period t-3 and t-4 are often significant if those in period t-2 are omitted.

In addition, we need to check for instrument validity. Within the context of GMM this is generally referred to as a test of the overidentification restrictions arising because the number of zero-correlation conditions (instruments) exceed the number of parameters to be estimated, and so cannot all be exactly zero. Following previous researchers we use Hansen's (1982) J-statistic to test the validity of the overidentifying restrictions - that the zero-correlation conditions are not jointly significantly different from zero.[34] We are able to select instruments which do not violate these overidentifying restrictions in addition to securing significant instrumented equations for all countries.

The following instrument set, $\Delta\ln C_{t-2}$, $\Delta\ln Y_{t-2}$, $\ln C_{t-2}$, $\ln Y_{t-2}$, is used for Austria, Canada, Greece, Italy, Japan, the Netherlands, Spain, Sweden, Switzerland and the UK.[35] For the other countries the instrument sets (specified in brackets) are: Australia ($\Delta\ln Y_{t-4}$, $\ln C_{t-2}$, $\ln Y_{t-2}$); Belgium ($\Delta\ln Y_{t-2}$, $\ln C_{t-2}$, $\ln Y_{t-2}$); Denmark ($\Delta\ln C_{t-2}$, $\Delta\ln Y_{t-2}$, $\ln C_{t-2}$, $\ln Y_{t-2}$, $\Delta\Delta\ln P_{t-2}$, U_{t-2}); Finland ($\Delta\ln C_{t-2}$, $\Delta\ln Y_{t-2}$, $\ln C_{t-2}$, $\ln Y_{t-2}$, ΔU_{t-2}); France ($\Delta\ln C_{t-2}$, $\Delta\ln Y_{t-2}$, $\ln C_{t-2}$, $\ln Y_{t-2}$, $\Delta\Delta\ln P_{t-2}$, U_{t-2}); Germany ($\Delta\ln C_{t-2}$, $\Delta\ln Y_{t-2}$, $\ln C_{t-2}$, $\ln Y_{t-2}$, $\Delta\Delta\ln P_{t-2}$); Iceland ($\Delta\ln C_{t-2}$, $\Delta\ln Y_{t-2}$, $\ln C_{t-2}$, $\ln Y_{t-2}$, $\Delta\ln P_{t-2}$, U_{t-2}); Ireland ($\Delta\ln C_{t-4}$, $\Delta\ln P_{t-2}$, U_{t-3}, U_{t-4}); Norway ($\Delta\ln C_{t-2}$, $\Delta\ln Y_{t-2}$, $\ln C_{t-2}$, $\ln Y_{t-2}$, ΔU_{t-2}); the USA ($\Delta\ln C_{t-4}$, $\Delta\ln Y_{t-2}$, $\ln Y_{t-2}$, Δr_{t-2}). When the real interest rate enters the structural model, either on its own or with income, we add r_{t-2} to the above instrument set for both the income growth and interest rate equations. An intercept is included in all instrumented equations.

Table 6.3 presents the GMM results for each country of the three analogues of the REPIH/RELCH models, allowing for intertemporal substitution (headed **(6.2.26')**), a proportion

[34] The J-statistic is given as $\{T.(\mathbf{y}-\mathbf{X}\boldsymbol{\beta})'\mathbf{WAW}'(\mathbf{y}-\mathbf{X}\boldsymbol{\beta})\}$, where T is the number of observations used in the estimation period and the function $(\mathbf{y}-\mathbf{X}\boldsymbol{\beta})'\mathbf{WAW}'(\mathbf{y}-\mathbf{X}\boldsymbol{\beta})$ is evaluated at its minimum value - which corresponds to the estimate of $\boldsymbol{\beta}$. It follows a χ^2 distribution with degrees of freedom equal to the number of instruments less the number of parameters to be estimated. The overidentifying restrictions cannot be rejected if the J-statistic is below its associated critical value. That is, the smaller is this statistic, the closer are the orthogonality conditions, $\mathbf{W}'(\mathbf{y}-\mathbf{X}\boldsymbol{\beta})$, to zero, the less correlated are the instruments with the regression's error term. See Hamilton (1994) pp. 414-415 for more details.

[35] Hamilton (1994, p. 426) cites Monte Carlo evidence which recommends parsimony in the selection of instruments. Hence, we have a preference for instrument sets incorporating the fewest variables.

TABLE 6.3: GMM Estimates of Modified REPIH/RELCH Models Equations, (6.2.26'), (6.2.34) and (6.2.34')

Country	AUSTRALIA			AUSTRIA			BELGIUM			CANADA		
Model	(6.2.26')	(6.2.34)	(6.2.34')	(6.2.26')	(6.2.34)	(6.2.34')	(6.2.26')	(6.2.34)	(6.2.34')	(6.2.26')	(6.2.34)	(6.2.34')
Instrumented Equation												
Adj R^2	0.533	0.132	0.118	0.357	0.209	0.200	0.469	0.189	0.172	0.473	0.209	0.238
$PrF(R)^2$	[0.000]	[0.061]	[0.101]	[0.003]	[0.025]	[0.043]	[0.000]	[0.023]	[0.045]	[0.000]	[0.025]	[0.023]
REPIH/RELCH Model												
Intercept	0.016	0.009	0.009	0.034	0.005	0.001	0.031	0.010	0.021	0.037	-0.000	-0.004
	(9.602)	(2.566)	(2.915)	(13.87)	(1.566)	(0.164)	(7.538)	(1.845)	(2.599)	(7.908)	(-0.103)	(-0.402)
ΔlnY		0.632	0.611		0.665	0.741		0.574	0.313		0.855	0.916
		(2.838)	(2.886)		(6.792)	(4.843)		(4.020)	(1.433)		(8.853)	(4.647)
r	0.026		-0.018	-0.364		0.083	-0.284		-0.262	-0.703		0.087
	(0.692)		(-0.620)	(-2.429)		(0.473)	(-1.759)		(-2.139)	(-2.977)		(0.430)
Adj R^2	--0.120	0.392	0.390	0.070	0.439	0.391	-0.116	0.576	0.365	-0.077	0.645	0.625
SBIC	-8.309	-8.290	-8.846	-7.950	-8.455	-8.303	-7.671	-8.638	-8.164	-7.107	-8.217	-8.092
J(DOF)	5.25(3)	2.59(2)	2.56(2)	5.74(4)	4.69(3)	4.13(3)	1.96(3)	3.47(2)	1.79(2)	2.45(4)	3.15(3)	3.50(3)
Residuals												
r(1)	0.171	-0.274	-0.257	-0.005	-0.180	-0.151	0.183	-0.192	0.035	0.142	0.099	0.056
r(2)	-0.267	-0.188	0.182	0.060	-0.175	-0.184	-0.011	0.107	-0.065	0.059	-0.023	-0.077
LB(6)	7.569	8.018	7.324	7.266	4.731	4.713	7.996	10.335	9.207	6.315	7.677	8.509

Country	DENMARK			FINLAND			FRANCE			GERMANY		
Model	(6.2.26')	(6.2.34)	(6.2.34')	(6.2.26')	(6.2.34)	(6.2.34')	(6.2.26')	(6.2.34)	(6.2.34')	(6.2.26')	(6.2.34)	(6.2.34')
Instrumented Equation												
Adj R^2	0.677	0.275	0.253	0.254	0.321	0.298	0.682	0.574	0.561	0.245	0.304	0.280
$PrF(R)^2$	[0.000]	[0.017]	[0.032]	[0.024]	[0.005]	[0.012]	[0.000]	[0.000]	[0.000]	[0.027]	[0.007]	[0.016]
REPIH/RELCH Model												
Intercept	0.037	0.015	0.017	0.016	0.014	0.005	0.045	0.010	0.003	0.041	0.004	0.012
	(7.347)	(6.869)	(6.357)	(2.345)	(4.015)	(1.518)	(25.554)	(4.462)	(0.848)	(9.097)	(2.781)	(2.435)
ΔlnY		0.601	0.654		0.642	0.676		0.632	0.841		0.944	0.823
		(10.863)	(11.639)		(7.758)	(8.208)		(7.462)	(8.318)		(18.229)	(8.214)
r	-0.109		-0.265	0.336		0.218	-0.539		-0.095	-0.506		-0.266
	(-0.865)		(-2.244)	(2.233)		(2.748)	(-7.805)		(-2.021)	(-2.678)		(-1.809)
Adj R^2	-0.280	0.179	0.185	-0.073	0.594	0.556	-0.588	0.646	0.486	0.059	0.747	0.815
SBIC	-6.690	-7.134	-7.070	-6.579	-7.551	-7.390	-7.718	-9.218	-8.777	-7.277	-8.588	-8.832
J(DOF)	6.27(6)	5.36(5)	4.83(5)	6.20(5)	2.84(4)	2.66(4)	4.73(6)	3.99(5)	5.01(5)	5.60(5)	1.79(4)	1.93(4)
Residuals												
r(1)	0.162	0.385	0.368	0.601	0.139	0.206	0.471	0.030	0.038	0.204	-0.145	-0.199
r(2)	-0.131	0.024	-0.109	0.279	-0.048	0.016	0.268	-0.025	0.060	0.035	0.171	0.010
LB(6)	9.548	18.908	24.269	18.812	3.632	2.486	21.153	5.101	5.444	3.704	9.099	9.755

TABLE 6.3 continued

Country	GREECE			ICELAND			IRELAND			ITALY		
Model	(6.2.26')	(6.2.34)	(6.2.34')	(6.2.26')	(6.2.34)	(6.2.34')	(6.2.26')	(6.2.34)	(6.2.34')	(6.2.26')	(6.2.34)	(6.2.34')
Instrumented Equation												
Adj R^2	0.458	0.254	0.289	0.777	0.315	0.295	0.450	0.226	0.283	0.816	0.394	0.394
$PrF(R)^2$	[0.000]	[0.011]	[0.009]	[0.000]	[0.009]	[0.017]	[0.000]	[0.019]	[0.011]	[0.000]	[0.001]	[0.001]
REPIH/RELCH Model												
Intercept	0.033	0.006	0.008	0.035	0.009	0.008	0.031	0.019	0.014	0.037	0.010	0.024
	(7.708)	(2.544)	(2.900)	(2.041)	(2.635)	(1.195)	(8.871)	(6.420)	(3.013)	(11.027)	(3.120)	(4.503)
ΔlnY		0.680	0.565		0.702	0.680		0.367	0.455		0.781	0.617
		(8.527)	(7.042)		(9.633)	(11.364)		(2.858)	(3.431)		(10.978)	(6.867)
r	0.249		-0.018	-0.178		0.009	-0.0004		0.118	0.076		-0.205
	(4.240)		(-0.319)	(-2.294)		(0.315)	(-0.004)		(1.546)	(0.876)		(-2.254)
Adj R^2	0.045	0.664	0.702	-0.198	0.759	0.747	-0.075	0.347	0.384	-0.100	0.501	0.506
SBIC	-7.212	-8.257	-8.304	-5.217	-6.819	-6.703	-6.731	-7.229	-7.217	-7.247	-8.037	-7.978
J(DOF)	4.83(4)	2.91(3)	3.57(3)	6.65(6)	3.43(5)	3.54(5)	4.83(3)	3.15(2)	1.89(2)	6.23(4)	4.94(3)	4.97(3)
Residuals												
r(1)	0.540	-0.223	-0.046	0.230	-0.238	-0.229	0.268	0.089	-0.017	0.443	-0.004	0.103
r(2)	0.367	0.012	0.188	-0.219	-0.013	-0.019	-0.088	-0.059	-0.016	0.015	-0.491	-0.308
LB(6)	35.027	10.965	9.011	6.549	6.328	6.269	7.040	5.606	3.924	10.854	13.396	10.731

Country	JAPAN			NETHERLANDS			NORWAY			SPAIN		
Model	(6.2.26')	(6.2.34)	(6.2.34')	(6.2.26')	(6.2.34)	(6.2.34')	(6.2.26')	(6.2.34)	(6.2.34')	(6.2.26')	(6.2.34)	(6.2.34')
Instrumented Equation												
Adj R^2	0.386	0.629	0.659	0.354	0.346	0.327	0.550	0.265	0.244	0.373	0.601	0.590
$PrF(R)^2$	[0.001]	[0.000]	[0.000]	[0.003]	[0.002]	[0.005]	[0.000]	[0.014]	[0.028]	[0.002]	[0.000]	[0.000]
REPIH/RELCH Model												
Intercept	0.031	0.002	0.003	0.030	0.001	-0.007	0.024	0.010	-0.008	0.041	0.005	0.005
	(7.219)	(0.591)	(0.607)	(8.070)	(0.299)	(-0.712)	(6.965)	(1.410)	(-1.041)	(5.817)	(2.125)	(2.239)
ΔlnY		0.918	0.775		0.993	1.185		0.632	0.903		0.849	0.854
		(13.269)	(12.484)		(7.923)	(5.115)		(2.824)	(3.182)		(18.966)	(17.674)
r	-0.040		0.182	-0.432		0.101	-0.141		-0.289	-0.235		-0.020
	(-0.210)		(1.868)	(-2.653)		(0.491)	(-1.315)		(-3.051)	(-1.241)		(-0.421)
Adj R^2	-0.298	0.761	0.783	-0.046	0.520	0.284	-0.020	0.200	-0.124	-0.225	0.664	0.653
SBIC	-6.652	-8.341	-8.369	-7.444	-8.222	-7.752	-7.059	-7.303	-6.892	-6.424	-7.717	-7.615
J(DOF)	4.24(4)	1.79(3)	4.17(3)	3.64(4)	3.89(3)	3.22(3)	3.01(5)	3.08(4)	2.98(4)	4.55(4)	2.00(3)	1.86(3)
Residuals												
r(1)	0.539	0.100	0.109	0.478	-0.172	-0.208	0.117	0.043	0.041	0.405	-0.113	-0.119
r(2)	0.453	0.097	0.056	0.169	-0.084	-0.071	-0.113	-0.086	-0.072	0.352	-0.025	-0.023
LB(6)	44.197	1.644	1.113	10.269	5.684	4.133	6.263	2.204	2.601	13.774	2.189	2.291

TABLE 6.3 Continued

Country	SWEDEN			SWITZERLAND			UK			USA		
Model	(6.2.26')	(6.2.34)	(6.2.34')	(6.2.26')	(6.2.34)	(6.2.34')	(6.2.26')	(6.2.34)	(6.2.34')	(6.2.26')	(6.2.34)	(6.2.34')
Instrumented Equation												
Adj R^2	0.477	0.448	0.432	0.467	0.316	0.421	0.443	0.243	0.222	0.102	0.212	0.267
PrF(R)^2	[0.000]	[0.000]	[0.001]	[0.000]	[0.004]	[0.001]	[0.000]	[0.014]	[0.029]	[0.150]	[0.024]	[0.014]
REPIH/RELCH Model												
Intercept	0.022	0.004	0.003	0.017	0.000	0.002	0.018	0.016	0.011	0.025	0.005	-0.012
	(6.100)	(1.611)	(1.307)	(6.834)	(0.078)	(0.369)	(4.261)	(4.156)	(2.155)	(6.776)	(1.200)	(-1.561)
ΔlnY		0.798	0.824		0.853	0.822		0.331	0.403		0.806	1.139
		(6.133)	(6.274)		(8.379)	(8.079)		(2.213)	(1.597)		(5.291)	(5.052)
r	0.082		-0.064	0.481		0.230	0.418		0.288	0.067		0.438
	(1.014)		(-0.929)	(2.283)		(3.027)	(4.575)		(2.277)	(0.034)		(2.066)
Adj R^2	-0.232	0.235	0.225	-0.406	0.582	0.493	-0.053	0.413	0.432	-0.182	0.643	0.061
SBIC	-7.300	-7.778	-7.693	-7.578	-8.792	-8.527	-7.357	-7.941	-7.902	-7.843	-9.040	-8.003
J(DOF)	4.62(4)	4.80(3)	4.02(3)	4.59(4)	5.99(3)	2.38(3)	2.45(4)	5.74(3)	3.33(3)	4.73(4)	2.31(3)	1.86(3)
Residuals												
r(1)	0.499	0.231	0.216	0.701	0.233	0.429	0.493	0.409	0.491	0.367	0.007	0.111
r(2)	0.268	0.116	0.099	0.415	-0.027	0.213	0.071	-0.014	0.083	-0.086	-0.184	-0.074
LB(6)	14.308	6.611	6.731	27.535	8.790	19.964	11.814	14.759	13.141	7.885	6.264	4.599

Table 6.3 Notes. The top row of the table specifies the country to which the results relate. For each country there are three different column headings **(6.2.26')**, **(6.2.34)** and **(6.2.34')**, indicating the model to which the results relate. In the section headed "Instrumented Equation" the coefficient of determination adjusted for degrees of freedom, "Adj R^2", and the probability value of the F-test for the significance of the regression, "PrF(R^2)", are reported for all three equations. A probability value below 0.050 indicates that the instrumented equation is significant at the 5% level. The instrument set used for each country is detailed in the text. In the section headed "REPIH/RELCH Model" the estimated coefficients for the constant, "Intercept", income growth, "ΔlnY", and the real interest rate variable, "r", are reported with associated t-ratios given in brackets immediately below. Bold emphasis indicates an insignificant coefficient. The adjusted R^2, "Adj R^2", Schwartz's Bayesian Information Criterion, "SBIC" and the J-statistic to test the overidentification restrictions, "J(DOF)", with degrees of freedom in brackets (DOF), are also reported for each model. The χ^2(DOF) critical values for J are: 5.99(2), 7.81(3), 9.49(4), 11.07(5) and 12.59(6). The first and second autocorrelation coefficients, "r(1)" and "r(2)", for the residuals of each model along with the Ljung-Box statistic for the joint significance of the first six residual autocorrelation coefficients, "LB(6)", is also given in the section headed "Residuals". The 5% critical values for the residual autocorrelation coefficients are ±0.345 and for the Ljung-Box test is 12.59. Bold emphasis indicates significant autocorrelation. The growth of consumption is the dependent variable for all models and the estimation period is 1960-1994.

of current income consumers (headed **(6.2.34)**), or both (headed **(6.2.34')**). The first section (headed "Instrumented Equation") gives the adjusted R^2 and the probability of the F-test for the significance of the regression for the instrumented equations. In all cases the instrumented equations are significant at the five percent level (with all the F-test probabilities below 0.050), except the income growth equations for Australia (which are significant at the ten percent level, so we tentatively suggest these provide valid inference) and the interest rate equation for the

USA (which is significant at the fifteen percent level). The first criterion for valid instruments is satisfied in virtually all cases: the instruments are significantly correlated with the variable being instrumented. The next section (headed "REPIH/RELCH Model") gives the estimated parameters, adjusted R^2, SBIC and the J-statistic (with associated degrees of freedom, DOF, in brackets) J(DOF) for each of the three REPIH/RELCH models estimated for each country. The overidentifying restrictions cannot be rejected at the five percent level for any of the models with all the J-statistics below their critical values. Thus, the second criteria for instrument validity (validity of overidentifying restrictions) is met for all the estimated specifications. This suggests that inferences drawn are legitimate, except those regarding interest rates for the USA.

For only four of the twenty countries is the coefficient on interest rates positive and significant (Austria, Greece, Switzerland and the UK) in model **(6.2.26')**. The fit of all the equations is very low, with all but the Greek model featuring a negative adjusted R^2. Indeed, a better specified model can be found for all countries. These results reject intertemporal substitution, which is consistent with the majority of researchers' findings including Hall (1988), Campbell and Mankiw (1991), Jin (1994) and Bacchetta and Gerlach (1997). However, our measure of consumption is total expenditure and not purely non-durables and this may cause our results to be unfavourable to intertemporal substitution, as Hahm (1998) warns.[36]

Estimated versions of CM's formulation without interest rates, **(6.2.34)**, shows that current income is significant with a coefficient between zero and one for all twenty countries, including Australia. Thus, this model is consistent with there being a significant proportion of current income consumers in all countries. This proportion varies considerably across countries with a low of 33.1% in the UK and a high of 99.3% in the Netherlands relative to an average of 71.2%.[37] This current income model represents a superior description of the data compared to

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[36] To ascertain whether durables are more sensitive to interest rates than nondurables one would wish to use consumption disaggregated into these two expenditure categories. We do not pursue this because separate data on durables and nondurables is not currently available for the many of OECD countries over the long time span we consider here.

[37] The majority of OECD agents' consumption expenditures between 1960 and 1994 appear to be determined by current income.

the simple intertemporal substitution model of Hall (1988) with far greater fit (according to both the adjusted R^2 and SBIC) for every country. The adjusted R^2 for the current income specification ranges from Denmark's 17.9% to 76.1% for Japan with an average explanatory power of 52.6%. Only three countries (Denmark, Italy and the UK) exhibit significant autocorrelation (though the GMM estimates are robust to autocorrelation). Therefore, this model appears to provide a satisfactory description of the data for all twenty countries.

Adding interest rates to the current income CM model does not improve the description of the data for any country. The interest rate is invariably incorrectly signed and statistically insignificant while fit generally deteriorates and three countries' equations suffer from autocorrelation. However, there are five countries where we reserve judgement. For Finland, Switzerland and the USA both income growth and interest rates enter with positive significance, but these models exhibit worse fit relative to the specification solely containing income growth. The addition of interest rates improves the fit of the Japanese model and enters with the expected positive sign, if it is (just) insignificant. Finally, interest rates enter with positive significance in the UK equation and the adjusted R^2 increases. However, the income term becomes insignificant and the SBIC deteriorates. With these reservations in mind, it seems that, in general, there is little support for adding interest rates to the CM model. Departures from the REPIH/RELCH appear to be better characterised by a proportion of current income consumers than variable interest rates.

6.4.3 Instrumental Variables (IV) Estimation Results

In this sub-section we estimate the three modified REPIH/RELCH models considered above, **(6.2.26')**, **(6.2.34)** and **(6.2.34')**, with the IV method that explicitly incorporates an MA process. This allows us to jointly discern the extent of the role of durability, variable interest rates and current income consumers in the rejection of the REPIH/RELCH. The presence of durability will be indicated by the MA terms being (jointly) significant. The expected sign of the MA terms' sum is ambiguous, though we suspect it will likely be positive. When considering MA errors we use instruments lagged by two periods plus the order of the MA process to ensure the instruments are orthogonal of the error process. (We use the iterative IV/MA procedure in

Eviews 2.0 to estimate regressions with MA terms). Since we explicitly model autocorrelation with MA errors and do not expect heteroscedasticity with such time-series regressions using logarithmic variables, we do not adjust coefficient standard errors.

CM argue that the implied cross equation (overidentifying) restrictions cannot be tested using Sargan's (1964) statistic when the error term is autocorrelated and/or heteroscedastic. Many recent researchers who estimate CM-style models adopt the alternative Wald test that they suggested. Although we explicitly use an MA process to remove autocorrelation, we use CM's alternative test following previous researchers. This procedure involves the addition of all the instruments used, less at least one, to the IV equation and employing a Wald test to determine if the instruments are jointly significant.[38] This follows a chi-square distribution with the degrees of freedom equal to the number of instruments tested.

Due to the use of moving average error terms, different estimation method and different test for instrument validity, the instruments employed for any particular country may be different from those used with GMM. For example, the addition of moving average error terms causes further lagging of the instruments to help ensure that they are uncorrelated with the error term (see Flood and Garber 1980).[39] The instruments we choose satisfy both the overidentifying restrictions and secure significant instrumented equations, except for the UK where the overidentifying restrictions are rejected at the five percent, but not one percent, level. The instruments used are $\Delta\ln C_{t-2}$, $\Delta\ln Y_{t-2}$, $\ln C_{t-2}$ and $\ln Y_{t-2}$ for Austria, Canada, Greece, Japan, the Netherlands, Spain, Sweden and the UK. The instrument sets for the following countries are

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[38] One cannot add all of the instruments to the IV equation as the resultant perfect multicollinearity will prevent the model from being estimated and, therefore, the test from being conducted. Thus, at least one instrument must be excluded. Indeed, the problem of identification may dictate the further reduction of the number of instruments that can be added to the IV regression. The test is conducted to ensure the order condition is satisfied. In our application, we conduct two tests with different combinations of instruments to ensure all variables are tested for instrument validity.

[39] It is interesting to note that significant instrumented equations, made up of instruments dated no earlier than period t-3 or t-4, which also satisfy the overidentification restrictions could be obtained for all countries. However, in some cases, careful model reduction was required to achieve this.

given in brackets: Australia ($\Delta\ln C_{t-4}$, $\ln Y_{t-2}$); Belgium ($\Delta\ln Y_{t-2}$, $\ln C_{t-2}$, $\ln Y_{t-2}$); Denmark ($\ln Y_{t-2}$, $\Delta\Delta\ln P_{t-2}$, U_{t-2}); Finland ($\Delta\ln C_{t-2}$, $\Delta\ln Y_{t-2}$, $\ln C_{t-2}$, $\ln Y_{t-2}$, ΔU_{t-2}); France ($\Delta\ln C_{t-2}$, $\Delta\ln Y_{t-2}$, $\ln C_{t-2}$, $\ln Y_{t-2}$, U_{t-2}, $\Delta\Delta\ln P_{t-2}$); Germany ($\Delta\ln C_{t-2}$, $\Delta\ln Y_{t-2}$, $\ln C_{t-2}$, $\ln Y_{t-2}$, $\Delta\Delta\ln P_{t-2}$); Iceland ($\Delta\ln C_{t-2}$, $\Delta\ln Y_{t-2}$, $\ln C_{t-2}$, $\ln Y_{t-2}$, $\Delta\ln P_{t-2}$, U_{t-2}); Ireland ($\Delta\ln C_{t-4}$, $\Delta\ln P_{t-2}$, U_{t-3}, U_{t-4}); Italy ($\Delta\ln C_{t-4}$, $\Delta\ln Y_{t-4}$, $\ln C_{t-4}$, $\ln Y_{t-4}$); Norway ($\Delta\ln Y_{t-2}$, $\ln C_{t-2}$, $\ln Y_{t-2}$, ΔU_{t-2}); Switzerland ($\Delta\ln C_{t-3}$, $\Delta\ln Y_{t-3}$, $\ln C_{t-3}$, $\ln Y_{t-3}$); USA ($\Delta\ln C_{t-4}$, $\Delta\ln Y_{t-2}$, $\ln Y_{t-2}$, Δr_{t-2}). An intercept is included in all instrument sets.

Models incorporating interest rates, with and without income growth were tried. When included on their own the models typically suffered from autocorrelation and their coefficient was generally incorrectly signed and/or statistically insignificant. When included with income they never entered with both positive coefficient and statistical significance. This confirms the results, obtained using GMM, against intertemporal substitution. Interest rates do not feature in any of the countries' preferred models.

Table 6.4 presents the estimation results of the favoured IV specification adopted for each country. We only report results for **(6.2.34)** because models incorporating interest rates are not favoured. In all cases, the instrumented income growth equations are statistically significant and the validity of the overidentifying restrictions cannot be rejected according to two Wald tests [their probabilities are denoted Pr(Wald1) and Pr(Wald2)], although a 1% level is required to secure instrument validity for the UK. With the exception of the Danish, Swiss and UK models there is no evidence of statistically significant autocorrelation at the five percent level. Serial correlation is insignificant at the one percent level for Switzerland and the UK, but remains significant for Denmark.[40] We therefore present our results as providing valid inference, with reservations regarding autocorrelation in the Danish model. The regressions exhibit reasonable explanatory power with the adjusted R^2 ranging from 15.1% for Norway to 83.5% for Germany compared to an average of 56.3%. These equations fit the data slightly better than those using GMM. This may be expected given the use of moving average error terms for two countries.

Moving average error terms are only required for two of the twenty countries (Italy and

[40] Autocorrelation could not be satisfactorily removed with the inclusion of MA terms.

TABLE 6.4: IV Estimates of Modified REPIH/RELCH Models, Equation (6.2.34)

Country→	Aul	Aut	Bel	Can	Den	Fin	Fra	Ger	Gre	Ice
Instrumented Equation										
Adj R^2	0.121	0.208	0.189	0.209	0.325	0.321	0.574	0.304	0.255	0.315
PrF(R^2)	[0.048]	[0.025]	[0.023]	[0.025]	[0.002]	[0.005]	[0.000]	[0.007]	[0.011]	[0.009]
REPIH/RELCH Model										
Int	0.009	0.006	0.012	0.003	0.009	0.006	0.008	0.004	0.007	0.009
	(2.518)	**(0.836)**	(2.292)	**(0.815)**	**(1.638)**	**(1.170)**	(2.687)	**(1.157)**	**(1.533)**	**(1.659)**
ΔlnY	0.594	0.729	0.476	0.815	0.545	0.808	0.685	0.887	0.687	0.731
	(3.221)	(3.687)	(2.804)	(4.940)	(3.185)	(5.029)	(7.182)	(8.536)	(6.703)	(6.984)
MA(1)										
MA(2)										
ΣMA	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
Adj R^2	0.433	0.450	0.556	0.660	0.279	0.629	0.640	0.781	0.662	0.759
SBIC	-8.989	-8.476	-8.592	-8.260	-7.264	-7.640	-9.201	-8.736	-8.250	-6.821
Pr(Wald1)	[0.780]	[0.524]	[0.264]	[0.223]	[0.410]	[0.220]	[0.666]	[0.302]	[0.846]	[0.778]
Pr(Wald2)	[0.797]	[0.628]	[0.417]	[0.213]	[0.332]	[0.529]	[0.291]	[0.333]	[0.737]	[0.748]
Pr[F(π_G)]	[0.837]	[0.747]	[0.569]	[0.809]	[0.746]	[0.309]	[0.580]	[0.590]	[0.941]	[0.781]
Residuals↓										
r(1)	-0.263	-0.182	-0.058	-0.110	**0.391**	-0.009	-0.005	-0.114	-0.228	-0.249
r(2)	0.159	-0.211	0.105	-0.014	0.040	-0.128	-0.038	0.162	0.006	0.001
LB(6-q)	7.192(6)	5.713(6)	7.482(6)	9.111(6)	**19.50(6)**	2.659(6)	5.773(6)	8.446(6)	11.035(6)	6.597(6)
Forecast↓										
Pr[F(11)]	[0.433]	[0.581]	[0.058]	**[0.012]**	[0.140]	**[0.002]**	[0.206]	**[0.001]**	[0.652]	[0.057]

Table 6.4 continued

Country→	Ire	Ita	Jap	Net	Nor	Spa	Swe	Swz	UK	USA
Instrumented Equation										
Adj R^2	0.226	0.409	0.629	0.346	0.271	0.601	0.448	0.330	0.243	0.212
PrF(R^2)	[0.019]	[0.000]	[0.000]	[0.002]	[0.009]	[0.000]	[0.000]	[0.003]	[0.014]	[0.024]
REPIH/RELCH Model										
Int	0.009	0.011	0.008	0.001	-0.001	0.006	0.006	0.003	0.007	0.005
	(1.375)	(3.754)	**(1.688)**	**(0.312)**	**(-0.079)**	**(1.079)**	**(1.316)**	**(0.966)**	**(1.447)**	**(1.176)**
ΔlnY	0.633	0.809	0.841	0.941	0.974	0.763	0.570	0.703	0.681	0.839
	(3.118)	(8.984)	(9.690)	(6.411)	(2.690)	(6.170)	(2.840)	(5.026)	(3.890)	(4.489)
MA(1)		-0.022						0.345		
		(-0.140)						(2.027)		
MA(2)		-0.539								
		(-3.482)								
ΣMA	0.000	-0.561	0.000	0.000	0.000	0.000	0.000	0.345	0.000	0.000
Adj R^2	0.416	0.601	0.783	0.564	0.151	0.665	0.299	0.697	0.626	0.643
SBIC	-7.341	-8.121	-8.442	-8.318	-7.243	-7.719	-7.864	-9.043	-8.392	-9.041
Pr(Wald1)	[0.351]	[0.291]	[0.723]	[0.825]	[0.087]	[0.460]	[0.240]	[0.757]	**[0.029]**	[0.354]
Pr(Wald2)	[0.487]	[0.525]	[0.483]	[0.861]	[0.207]	[1.000]	[0.461]	[0.295]	**[0.049]**	[0.472]
Pr[F(π_G)]	[0.199]	[0.751]	[0.384]	[0.726]	[0.352]	[0.494]	[0.264]	[0.293]	[0.054]	[0.861]
Residuals↓										
r(1)	-0.220	0.027	0.044	-0.153	0.001	-0.097	0.279	-0.026	**0.352**	-0.020
r(2)	0.039	-0.047	0.083	-0.079	-0.068	-0.019	0.184	-0.060	0.013	-0.170
LB_W(6-q)	6.819(6)	3.053(4)	1.624(6)	6.053(6)	1.797(6)	1.571(6)	10.09(6)	**11.809(5)**	11.73(6)	5.720(6)
Forecast↓										
Pr[F(11)]	[0.626]	[0.274]	[0.954]	[0.451]	**[0.010]**	[0.337]	**[0.015]**	[0.494]	**[0.000]**	[0.468]

Table 6.4 notes: The section headed "Instrumented Equation" gives the adjusted R^2, Adj R^2, and probability of an F-test for the significance of the regression, PrF(R^2), of the instrumented income equation. The instruments used for each country are detailed in the text. In the section headed "REPIH/RELCH Model" the coefficients corresponding to the intercept, income growth and the moving average terms are given in the rows denoted Int, ΔlnY, MA(1) and MA(2), respectively, with estimated t-ratios specified in brackets below. ΣMA specifies the sum of the estimated moving average terms, Adj R^2 is the adjusted coefficient of determination and SBIC is Schwartz's Bayesian Information Criterion. Pr(Wald1) and Pr(Wald2) denote the probabilities of Wald tests for the validity of the overidentification restrictions. Pr[F(π_G)] is the probability for an version of a Wald test for the equality of IV and GMM estimates of π. The section headed "Residuals" provides the first r(1) and second r(2) autocorrelation coefficients along with the Ljung-Box [LB(6-q)] statistic. The 5% critical values for the autocorrelation coefficients are ±0.345. Critical values for the Ljung-Box statistics vary with q. At the 5% (1%) level the critical values are 12.59 (16.81), 11.07 (15.09) and 9.49 (13.28), for q=0,1,2, respectively. Pr[F(11)] is the probability value of an F-test for parameter constancy when the sample is split in 1983/1984. Bold emphasis indicates significant autocorrelation, rejection of the overidentifying restrictions, equality of IV and GMM coefficient estimates and parameter stability and insignificant parameters (depending upon context).

Switzerland). For Switzerland the order of moving average process is one and the coefficient is positive and less than unity. In the Italian model the moving average error process is of order two and the sum of the coefficients is negative. This suggests evidence for the role of durability in these two countries.[41]

The coefficients on the income terms are statistically significant for all countries, falling between zero and one so are, in this sense, consistent with the proportion of current income consumers interpretation. This proportion varies considerably from country to country. It ranges from a low of 47.6% for Belgium to a high of 97.4% for Norway relative to an average value of 73.6% - the corresponding GMM average estimate is 71.2%. The average IV and GMM estimates are similar, the coefficients feature a simple correlation coefficient of 56.3%, and (the probability values of) a Wald test for the equivalence of IV and GMM estimates, denoted $\Pr[F(\pi_G)]$ in Table 6.4, indicates that they do not differ with statistical significance at the five percent level - see Holmes (1993) p. 1321 for the application of such a test to error correction models.[42]

Although the two estimators' estimates do not differ with statistical significance there are notable numerical differences for four countries. The estimates are 80.8% (IV) and 64.2% (GMM) for Finland, 63.3% (IV) and 36.0% (GMM) for Ireland, 97.4% (IV) and 63.2% (GMM) for Norway and 68.1% (IV) and 33.1% (GMM) for the UK. The potential causes of these different estimates are the use of different instruments, the inclusion of MA terms with IV and the different estimator used. For both Finland and Ireland the only difference between the GMM and IV estimates are the different estimation method employed. For Norway, the instrument set used for IV adds $\Delta \ln C_{t-2}$ relative to the set used for GMM and may cause differences beyond those

[41] Interestingly, if one compares the fit of the REPIH/RELCH models with the favoured standard adjustment ECMs, excluding dummies, estimated in Chapter five, we see that the former has a larger adjusted R^2 relative to the latter for the two countries incorporating an MA error process, being Italy and Switzerland. This indicates the importance of the MA process for these two countries. Australia is the only other country where the REPIH/RELCH model exhibits superior fit to the ECM.

[42] This Wald test applied to the GMM estimates indicates statistically different estimates from those obtained by IV at the five percent level for only two countries, Ireland and the UK. However, even for these two countries the estimates are not significantly different at the one percent level.

stemming from the estimator employed. For the UK both GMM and IV estimates are subject to first order autocorrelation, however, instrument validity is only secured at the 1% (and not 5%) with the IV estimates. Since the difference between estimates is largest for Norway and the UK, we suggest that use of different instruments may explain some of this divergence for these two countries. However, given that the estimation method appears to be the only difference for Finland and Ireland, we believe this to be a contributory factor for Norway and the UK as well. With very similar (average) explanatory power and generally similar estimates it is not obvious that the IV or GMM estimates are superior, so both sets will be used for cross country comparisons in Chapter seven.

We compare our estimates of the proportion of current income consumers with those provided by Jin's (1994) *time-series* analysis for the eighteen countries common to the two analyses. In general, Jin's (1994) estimates are much lower, with a minimum value of 10.9% to a maximum of 90.8% relative to an average value of 49.8%. Indeed, the correlations between our and Jin's estimates are negative![43] There are various potential reasons for this disparity. Firstly, Jin (1994) uses an approximate measure of disposable income whereas we use the more appropriate actual disposable income series. Secondly, unlike Jin (1994), our favoured model (excluding interest rates) only incorporates stationary regressors so is not subject to spurious/nonsense regression. Thirdly, Jin (1994) uses a shorter sample, 1964-1988, which excludes the recession of the early 1990s that dramatically affected the consumption patterns of many of the countries included in the sample. We use a longer time series which includes the 1990s downturn.[44] Thus, we believe our results to be more informative than those provided by Jin (1994).

[43] Our GMM (IV) estimates feature a negative correlation of -0.322 (-0.291) Jin's (1994) *time series* estimates. A negative correlation of -0.092 (-0.027) is found with Jin's (1994) *pooled* estimates.

[44] It should be noted that Jin (1994) avoids some of the approximating assumptions adopted in our logarithmic derivation which could affect the interpretation of our parameter. However, the GMM estimates we provide are essentially for the CM model with implicit allowance from HAC coefficient standard errors for any MA process potentially arising from durability. Jin (1994) implicitly accounts for durability, and any heteroscedasticity arising from his semi-logarithmic specification, with Newey and West (1987) standard errors.

The models estimated above assume that the proportion of current income consumers is constant through time. However, an implication of the presence of liquidity constraints and income uncertainty is that this proportion may vary through time, although these factors may offset each other leaving relatively constant parameters. Indeed, both CM and Bacchetta and Gerlach (1997), who consider five or six economies, find that π varies for some countries and not for others.

Four main methods have been employed in the literature to investigate time variation and all have drawbacks. First, one can estimate the model over sub-samples yielding information on the evolution of π over these periods. This assumes that π shifts at known dates which Bacchetta and Gerlach (1997) argue is implausible. Second, π can vary with a deterministic time trend. Once again this may be unrealistic because it suggests a gradual and continuous evolution in the proportion of current income consumers ruling out, for example, abrupt changes coinciding with the release of credit constraints. Third, recursive estimates, which assume constant parameters over an ever-changing sample, have been used. These models may be unable to capture short-term fluctuations in π. Fourth, one can allow π to evolve according to a random walk by using the Kalman Filter. One can do this, following McKiernan (1996), using a two-step IV procedure by substituting the fitted instrumented equation into the REPIH/RELCH model and then estimating the time-varying parameters, however, this treats the coefficients estimated in the instrumented equation as known with certainty, which they are not. Alternatively, one can follow Bacchetta and Gerlach (1997) and estimate a time-varying non-linear system of structural REPIH/RELCH and instrument equations simultaneously. Two drawbacks of this approach are the general greater sensitivity of system estimates to misspecification and that the estimates provided by the filter are "only approximate and suboptimal, and that little is known about its properties in *small* samples", (Bacchetta and Gerlach 1997, p. 229, my comments in italics).

Given the imperfections of all of these methods we simply aim to gauge whether π varies through time using a Wald test for each country's forecast accuracy for the IV estimates of equation **(6.2.34)**. The sample is split in 1983/1984 because the effects of financial deregulation took effect during the 1980s for most countries. The probability value of this Wald test statistic, Pr[F(11)], is reported in Table 6.4. For six countries (Canada, Finland, Germany, Norway, Sweden and the UK) there is evidence of time-varying parameters at the five percent level. Use of the one percent level suggests that π only varies for three countries (Finland, Germany and the UK). These results are consistent with previous work which suggest time-variation in the proportion of current income consumers for some countries but not others. It is also notable that the countries where time variation is prevalent are those where the severe effects of financial deregulation have been well documented - see for examples, Miles (1992) for the UK and Berg (1994) for the Nordic countries. Since there is no evident time-variation in the majority of countries' REPIH/RELCH models (fourteen at the five percent level and seventeen at the one percent level) we consider, in Chapter seven, whether liquidity constraints and/or income uncertainty can explain the variation in π across countrys rather than through time.[45]

6.5 Conclusion

We derive a model for total consumption in logarithmic form to allow for a proportion of current income consumers, durability, and an intertemporal substitution extension. We estimate various specifications nested within our model using both GMM and IV estimation. First we consider whether durability can, on its own, provide an explanation for the REPIH/RELCH rejection applied to total expenditures. Only half of the twenty countries' moving average models are consistent with durability being the sole cause of the REPIH/RELCH's rejection. This suggests that other factors are in operation for many countries.

Intertemporal substitution, in common with the majority of previous studies, does not provide a satisfactory explanation for the failure of the REPIH/RELCH. Generally, the coefficients on interest rates are statistically insignificant and/or negatively signed, even when durability is implicitly or explicitly accounted for and whether constrained consumers are controlled for or not. Models including interest rates were never favoured.

The REPIH/RELCH adjusted for a proportion of current income consumers provides a satisfactory explanation of total consumption growth for all countries. The estimated proportions

[45] Inspection of unreported recursive estimates of π for the countries where there is time-variation suggests that the fixed parameter estimates are representative of the whole sample.

of constrained consumers are plausible and vary considerably from country to country, however, we only find evidence for variation through time for three to six countries. The coefficient estimates obtained using the GMM and IV estimators are generally similar. However, both sets differ greatly from the estimates provided by Jin (1994). We believe that our longer sample, sole use of stationary regressors in our structural equation (when interest rates are excluded) and more appropriate measure of income makes our results more reliable and informative than those produced by Jin (1994). We find that the majority of consumers expenditures in the OECD are determined by current rather than expected income with an average proportion of current income consumers being approximately 70%. This is substantially higher than the average estimate of around 50% provided by Jin (1994). These estimates may better reflect the proportion of current income consumption relative to studies of non-durables because one may expect some current income expenditures to be on durables, partly explaining the large estimated proportions with this data. Other reasons for the large estimates include our sample covering the 1990s downturn, the use of statistically significant instrumented equations and a model specified using stationary regressors. Although large, such estimates are compatible with the (simultaneous) presence of liquidity constraints, income uncertainty and non-negligible information costs involved in rational expectations formation. Indeed, Lattimore provides evidence against substantive forward looking behaviour for Australia, Horioka (1996) cites evidence of the prevalence of a large proportion of constrained consumers in Japan while CM produce an estimate that 100% of French consumers are current income consumers. The explanation of the cross-country variation in our estimates, considered in Chapter seven, should help clarify the existence of such factors.

In addition to current income consumers the IV estimates reveal a significant MA process consistent with the importance of durability for Italy and Switzerland. In general, any MA process exhibited by the pure MA models become statistically insignificant when income growth is added, suggesting it is excess sensitivity to income rather than durability that causes the rejection of the REPIH/RELCH in most countries.

Obtaining valid instruments for income growth has proven troublesome in previous studies. A consensus of earlier work suggests that the lagged log-level of the APC is a crucial instrument. We find that use of the separate log-levels of consumption and income improve the validity of instrumented income growth equations. We suggest that this is likely to be due to the APC's nonstationarity and that separate inclusion of consumption and income allows the greater possibility of instruments forming stationary linear combinations of superior significance.

CHAPTER 7

EXPLAINING CROSS-COUNTRY DIFFERENCES IN CONSUMER BEHAVIOUR

7.1 Introduction

This Chapter seeks to explain the cross-country differences in the parameters estimated using the vector error correction model (VECM), error correction model (ECM), and rational expectations (REPIH/RELCH) consumption function reported in previous Chapters. These give us six sets of coefficients: the long and short run elasticities of consumption with respect to income and inflation, the adjustment of consumption to deviations from its estimated equilibrium, and the proportion of current income consumers. These parameters are divided into the following four groups. The short and long run elasticities of consumption to income; the short and long run elasticities of consumption to inflation; the adjustment coefficients; and the proportions of current income consumers.

These estimated coefficients vary considerably from country to country, see Figures 7.1 to 7.4. The top half of Figure 7.1 plots the estimated long run consumption elasticity with respect to income. The values of these coefficients range from 0.569 for Italy to 1.464 for Denmark relative to an average value of 1.014 (with a standard deviation of 0.205 and coefficient of variation of 0.201). The bottom half of Figure 7.1 plots the corresponding estimated short run income elasticities. The values are generally lower and more variable than their long run counterparts, ranging from 0.205 for Italy to 0.882 for Finland with a mean of 0.603 (the standard deviation is 0.236 and the coefficient of variation is 0.392). Figure 7.2 plots the estimated long and short run inflation elasticities. The long run inflation elasticities, top graph of Figure 7.2, take values ranging from -3.645 for Italy to 1.926 for Denmark relative to an average value of -0.394 (with a standard deviation of 1.135 and coefficient of variation of 2.880). The Italian value is extremely low compared to the other values and may be regarded as an outlier. The bottom graph in Figure 7.2 shows the short run inflation elasticities varying from

FIGURE 7.1: Estimated Long and Short Run Elasticities of Consumption with Respect to Income (LRY and SRY)

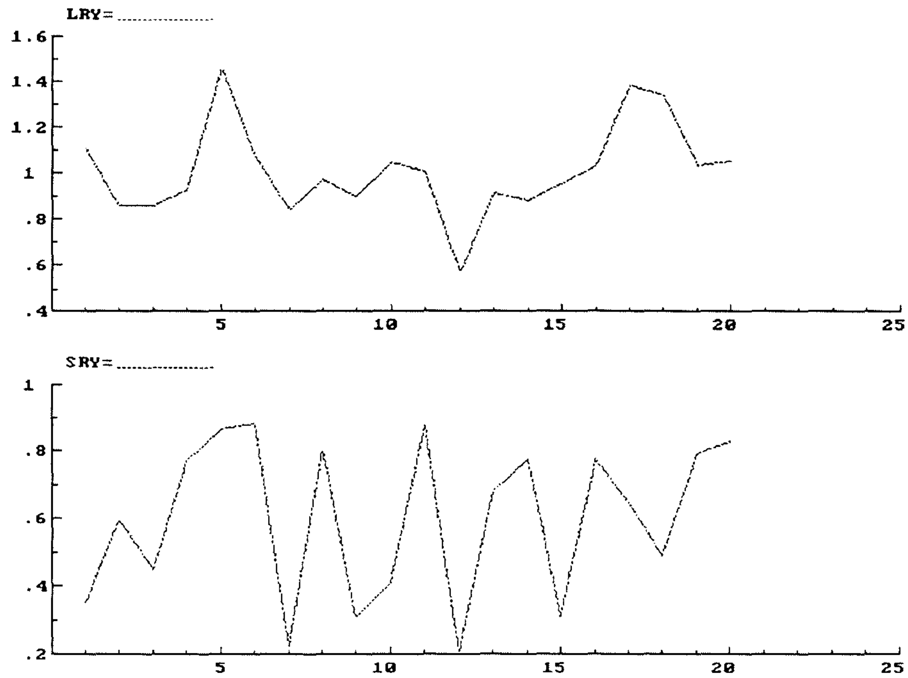
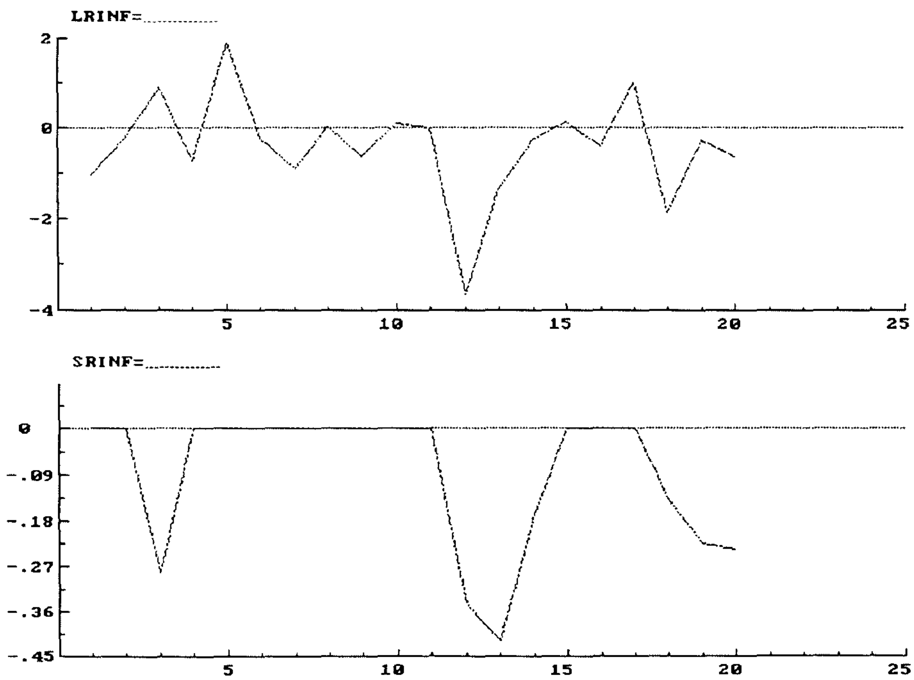


FIGURE 7.2: Estimated Long and Short Run Elasticities of Consumption with Respect to Inflation (LRINF and SRINF)



-0.419 for Japan to 0 for thirteen countries with a mean of -0.091 (the standard deviation is 0.138 and the coefficient of variation is 1.522). Only seven of the twenty countries exhibit a non-zero short run inflation elasticity because such effects were generally excluded from the favoured error correction models, hence the short run elasticities are generally smaller in magnitude and are less variable than their long run counterparts. Figure 7.3 shows the estimated adjustment coefficients. Their values range from -0.521 for Ireland to +0.042 for Sweden compared to an average value of -0.235 (with a standard deviation of 0.168 and coefficient of variation of 0.715). The positive values for Sweden and Switzerland are inconsistent with error correction behaviour while the Irish value is extremely low and quite non-typical. Figure 7.4 plots the GMM and IV estimates of the proportion of current income consumers. The GMM estimates, top half of Figure 7.4, range from 0.331 for the UK to 0.993 for the Netherlands relative to a mean of 0.712 (with a standard deviation of 0.175 and coefficient of variation of 0.246). The IV estimates, bottom half of Figure 7.4, are generally larger, with a mean of 0.736, but are less variable with values ranging from 0.476 (Belgium) to 0.974 (Norway) - the standard deviation is 0.132 and the coefficient of variation is 0.179. As pointed out in Chapter 6 the GMM and IV estimates are not significantly different, although there are some notable numerical differences for some countries, so we model both.[1]

We seek to identify factors which can explain the variations in these parameter estimates using cross-country regressions. For this we need to motivate potential explanatory factors. As far as we are aware the only study using cross-section regressions is Jin's (1994) analysis of cross-country variations in the proportions of credit constrained consumers.[2] We believe our parameter estimates to be superior to Jin's (1994) so an explanation of our estimates is important. We also expand upon the proxies of liquidity constraints used by Jin (1994) and additionally consider the role of precautionary saving. We are not aware of any previous cross-country analyses which

[1] In Chapter 6 we found that fourteen to seventeen countries' estimated proportions of current income consumers were relatively constant through time suggesting the estimates are representative of the sample so facilitating valid cross-country investigation.

[2] Both Jappelli and Pagano (1989) and Campbell and Mankiw (1991) note that the estimated proportions of credit constrained consumers appear to be lower in countries with more developed financial markets, however, no systematic regression analysis is conducted.

FIGURE 7.3: Estimated Adjustment Coefficient (ECM)

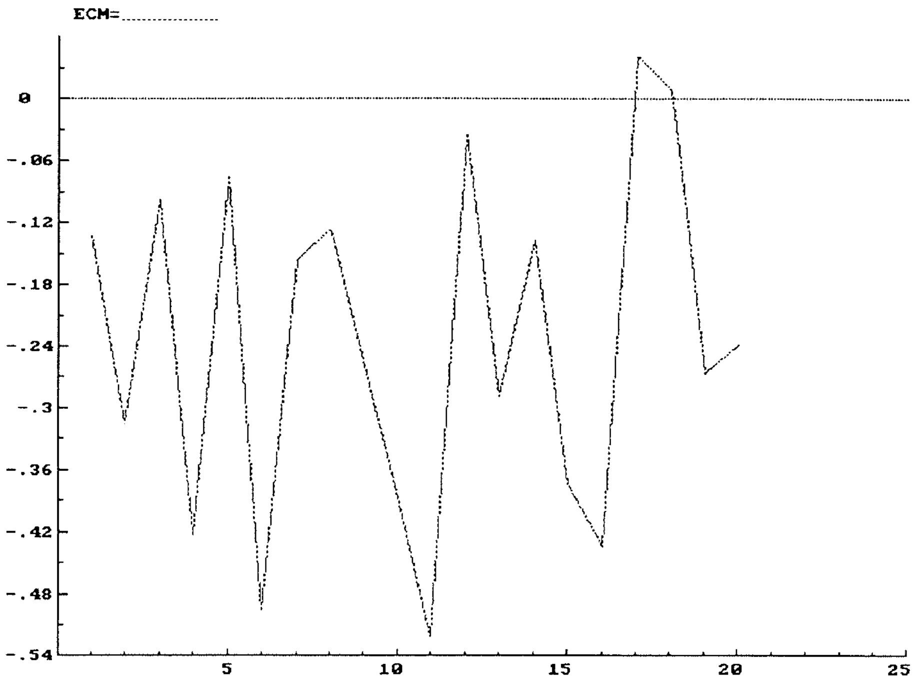
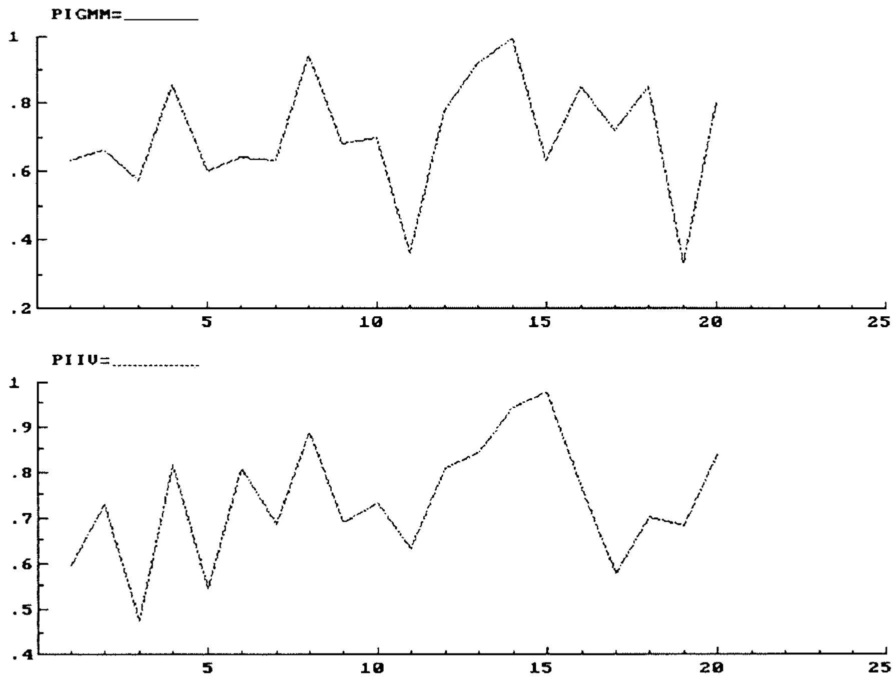


FIGURE 7.4: GMM and IV Estimates of the Proportion of Current Income Consumers (PIGMM and PIIV)



attempt to explain cross-country variations in the elasticity of consumption to income and inflation and the speed of adjustment of consumption to equilibrium. This represents an innovation of the present Chapter.

For each group of coefficients we outline potential explanatory factors and empirically assess their role in Sections 7.2 to 7.5. Section 7.6 presents conclusions.

7.2 Explaining Cross-Country Differences in the Response of Consumption to Income

In this section we empirically consider potential explanations for differences in the response of consumption to income, both short and long run. The theories reviewed do not directly refer to the variation of the estimated income elasticity parameters because there appears to be little explicit discussion on this topic. However, they do refer to responses of consumption to income and are, therefore, presented as providing potential explanatory factors for cross-country variations in the income elasticities.

7.2.1 Theoretical Considerations

Keynes (1936) conjectured that *windfall*, rather than planned, changes in non-human wealth,[3] large fluctuations in interest rates and alterations in income distribution could cause the marginal propensity to consume (MPC) to vary, though only in the short run. For example, Keynes is often attributed with the view that the short run influence of interest rates is relatively unimportant. However, he did recognise that there may be an important long run interest rate induced wealth effect. Changes in interest rates could affect consumption out of a given income to the extent that they altered the value of securities and other assets. Keynes believed that interest rates had an ambiguous impact upon consumption and rejected the classical economists' view of a clear negative association.

[3] Planned changes in wealth were, for example, regarded as the *result* rather than cause of consumers' savings decisions.

Keynes has also been attributed with the suggestion that the MPC falls as the *level* of income rises, with increased living standards reducing the proportion of income required to secure necessary consumption. This would appear to be a long run effect. Having said this, Hadjimatheou (1987) points out that "Keynes does not state in any explicit way that there is a secular trend of the propensity to consume to decline with income." (p. 2). Even though Keynes may not have pressed such a proposition, we regard it as an interesting hypothesis for the current cross-country analysis.

Duesenberry's (1949) relative income hypothesis (RIH) suggests that those with relatively lower living standards will attempt to emulate the consumption patterns of the better off. This implies that those commanding lower incomes will exhibit larger average propensities to consume (APCs) relative to higher income earners. If higher income earners dominate consumption the degree of income inequality *within* a country will be negatively associated with the proportion of income consumed. Duesenberry also argued that the savings rate was affected by changes in interest rates, income expectations, income growth and the age distribution of the population.

The development of the RIH into Brown's (1952) Habit Persistence model gives rise to consumption being a function of income and lagged consumption. Under the assumption that consumption and income grow at the same constant rate, it can be shown that the long run APC is negatively related to real income growth, which is consistent with the observed saving ratios of different countries through time - see Thomas (1994).

Friedman's (1957) permanent income hypothesis (PIH) suggests that permanent consumption is proportional to permanent income where the proportionality coefficient depends upon tastes, the return on wealth, age composition of the household and the ratio of non-human to human wealth. The latter ratio is expected to be positively associated with the proportionality coefficient because real wealth holdings provide a greater defence against an uncertain future than human capital. Thus, there is a precautionary motive. Deaton (1992) argues that an economy which exhibits greater *income* uncertainty may be expected to generate greater precautionary savings, and so exhibit a lower income elasticity compared to a country where consumers are more certain of their future incomes.

The life cycle hypothesis (LCH), as outlined by Ando and Modigliani (1963), suggests that consumption is a function of current income and wealth holdings, as well as expected future income. Assuming expected income is proportional to current income, the LCH implies that consumption is a function of current lifetime resources: income and wealth. In much of the (early) literature interest rate effects are ignored by assuming they are constant. When this assumption is relaxed, real interest rates are generally incorporated to account for intertemporal substitution suggesting a negative relationship with the APC. However, there may also be a positive relationship: if the income effect dominates the substitution effect - see Muellbauer (1994). However, Muellbauer (1994, p. 9) argues that interest rates will most likely have a small unstable impact in consumption functions, which is suggested to be consistent with the majority of empirical evidence. This unstable effect arises if the impact of interest rates varies over the business cycle and if the income and substitution effects offset each other to some degree. Both are suggested to be likely.

Modigliani (1986) outlines a simplified version of the LCH where a consumer attempts to maintain a constant level of consumption throughout their entire lifespan. This is achieved by saving a constant proportion of their income throughout their earning life to attain a level of wealth which is just sufficient to provide a constant flow of consumption throughout their retirement. Three of the six implications of this model outlined by Hadjimatheou (1987) are relevant for our present cross-country analysis. First, a country's APC is independent of the *level* of its per-capita income, which contrasts with the view that may be attributed to Keynes. Secondly, there is a negative relationship between an economy's APC and its income growth, which is consistent with the Habit Persistence version of the RIH. This arises because economic growth raises each generation's future income expectations and, therefore, their saving rates. Thus, at any moment in time, the larger is income growth the greater does the saving of the current generation exceed the saving of the previous generation. Thirdly, for a given income growth rate, the prevailing length of retirement is the major determinant of the APC. The longer is the length of retirement, the larger is the saving rate and the lower is the APC.

Modigliani (1990) postulates a basic LCH specification to characterise cross-country variations in the net *national* saving rate. The three explanatory factors are: GDP growth, the ratio of inflation-adjusted government saving to net national product and the dependency ratio. Following Modigliani (1986), GDP growth is expected to be positively (negatively) related to the national saving rate (APC).[4] Public saving will influence national saving if private sector saving does not fully adjust in response to public sector deficits or surpluses as suggested by Ricardian equivalence: fiscal policy has some degree of efficacy. Regarding our present analysis, if Ricardian equivalence holds, private saving will rise when public saving falls, suggesting a positive association between the income elasticity parameter and the fiscal surplus/deficit to GDP ratio. Without complete intergenerational altruism, fiscal policy can boost consumption, which would be consistent with a *small* positive, or a possible negative relationship, between the fiscal surplus/deficit and the income elasticity parameter.[5] Finally, an increase in the proportion of dependents in the population increases family needs, lowering saving, suggesting a positive relationship with the APC. However, a permanently higher dependency ratio implies a larger proportion of workers to retirees (the support ratio), which leads to increased saving: a negative relationship with the APC. Further, the support ratio may enter as well as, or instead of, the dependency ratio (justification for a broader set of demographic effects is outlined below).

Jappelli and Pagano (1994) extended Modigliani's (1990) model to consider whether liquidity constraints can explain international differences in (national) saving rates. They rationalise the inclusion of this variable by suggesting that, for example, the young may be unable to borrow upon the basis of their expected future income, and so are unable to follow their optimal lifetime consumption plan. Countries with less binding credit constraints will reduce such enforced saving suggesting a positive relationship between the APC and the degree of availability of credit. Using a panel of nineteen OECD countries over three decades (1960-1970, 1971-1980 and 1981-1987) they find general support for the extended Modigliani (1990) specification. That

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[4] Increased growth can, in a small open economy, reduce saving by stimulating the consumption of the young suggesting a positive relationship with the APC. Koskela and Viren (1989) implicitly find evidence for such a positive relationship, although the general evidence suggests the reverse association (see, for example, Jappelli and Pagano 1994).

[5] Since *complete* Ricardian equivalence requires a one for one replacement of private saving with public saving, a positive relationship between the fiscal surplus/deficit is consistent with efficacious fiscal policy provided that private saving only *partially* compensates for public saving.

is, GDP growth and government saving positively influence the ratio of net national saving to net national product. The latter provides evidence against Barro's (1974) hypothesis. The degree of credit constraints is found to be negatively associated with the national saving rate. However, the dependency ratio is found to exhibit no significant relationship.

Miles and Patel (1996) outline the importance of a broader set of demographic factors. They suggest the following characterisation of the consumer's life cycle. The (economically active) young aim to borrow, though may be frustrated by liquidity constraints. In mid life, people have families who need supporting, so spend to satisfy the needs of children (the dependency ratio effect). Only when free of dependents can the household begin to save for its retirement. Miles and Patel (1996) identify the age group of 50 to 64 as the period in an individual's life of substantial saving, and find strong empirical evidence to support this claim. Indeed, their results suggest that this age group could be 40 to 64; we therefore consider the age ranges 40-64, 45-64 and 50-64, to attempt to capture the impact of this high saving period of life. A parsimonious way of capturing these demographic effects may be to use the support ratio. Miles and Patel (1996) argue that the support ratio, the number of working age to the number of pensionable age, may be thought to be positively related to the saving rate, because it can, in simple terms, be viewed as representing the ratio of savers to dissavers. Thus, it would be negatively associated with the elasticity of consumption with respect to income.

Borooah and Sharpe (1986) argue that "consumption functions which treated all households as a single behavioural unit might be misleading" (p. 450) suggesting modification of the standard LCH-PIH. In particular, they argue that income distribution may affect the response of consumption to income. They cite Blinder (1975) as demonstrating that only under the LCH assuming no bequests are MPCs and APCs constant across income classes.[6] Using a DHSY-style consumption function for five equal UK income groupings Borooah and Sharpe (1986) find empirical support for the proposition that lower income earners typically exhibit larger propensities to consume than higher income earners. They conduct simulations demonstrating

[6] The invariance of consumption to income distribution also occurs with bequests if the marginal utility of bequests equal the marginal utility of income.

that policies which reduce income inequality raise aggregate consumption. Thus, a country with lower income inequality will likely feature a greater response of consumption to income relative to one with greater inequality.

7.2.2 A Cross-Country Model for the Elasticity of Consumption with Respect to Income

The discussion above provides a potential set of explanatory factors for the cross-country variation in the estimated income elasticities. Our general model is eclectic, drawing upon the predictions of different theories. These theories occasionally make opposing predictions about the influences of explanatory factors. Therefore, this general model helps us determine which theoretical aspects are important and which are not and, perhaps, which theoretical framework is most useful for the income elasticity under study. The general eclectic model is:

$$\beta_{Yi} = f(\underset{-(+)}{GRTH_i}, \underset{+/-}{DEF_i}, \underset{+(-)}{DEP_i}, \underset{+}{CRED_i}, \underset{-}{GAIN_i}, \underset{-(+)}{r_i}, \underset{-}{INEQ_i}, \underset{-}{f(INC_i)}, \underset{-}{UNCT_i}, \underset{-}{RTRE_i}, \underset{-}{SUPT_i}, \underset{-}{RSAV_i}) \quad \textbf{(7.2.1)}$$

β_{Yi} is the elasticity of consumption with respect to income for country i. $GRTH_i$ denotes an economy's income growth. DEF_i is the fiscal surplus/deficit to GDP ratio (capturing the impact of public saving on private saving). DEP_i is the dependency ratio. $CRED_i$ is private sector domestic credit to GDP ratio (a lower value indicates tighter credit constraints). The windfall capital gains and losses variable is denoted by $GAIN_i$. The real short term interest rate series is represented by r_i, income inequality by $INEQ_i$ and a function of the level of per-capita income, $f(INC_i)$. $UNCT_i$ represents income uncertainty, $RTRE_i$ denotes the length of retirement and $SUPT_i$ is the support ratio. Finally, $RSAV_i$ denotes the proportion of the population saving for retirement.

Regarding expected signs (given beneath the variables in **(7.2.1)**), Modigliani's LCH suggests that $GRTH_i$ should be negatively signed, but may also be positive (see footnote 4),[7] DEF_i will be positively signed (if there is some degree of Ricardian equivalence) and DEP_i will most likely feature a positive relationship but can also exhibit a negative association. Jappelli and Pagano

[7] Brown's (1952) Habit Persistence model suggests a negative relationship.

(1994) argue that $CRED_i$ will be positively signed. Keynes (1936) may be interpreted as suggesting a potential positive sign for $GAIN_i$. Typically a negative interest rate effect is expected, although both Keynes (1936) and the LCH can justify a positive coefficient. However, the former substitution effect is generally considered dominant while the presence of offsetting (income) effects may make the interest rate's influence insignificant.

The RIH and Borooah and Sharpe (1986) suggest a negative coefficient on $INEQ_i$. Keynes (1936), may be interpreted to suggest a negative sign for $f(INC_i)$. However, a linear relationship would imply that a continual rise in the *level* of per-capita income would cause an unbounded fall in the income elasticity, eventually making it negative, which is implausible. Therefore, one might expect a nonlinear relationship, allowing the income elasticity to decrease at a decreasing rate. Alternatively, $f(INC_i)$ may be unrelated to the income elasticity, as implied by Modigliani (1986).

The PIH suggests a negative coefficient on $UNCT_i$ while the same sign is expected for $RTRE_i$, (see Modigliani 1986). Age structure effects are predicted by the RIH, PIH and the LCH. For example, Miles and Patel (1996) suggests that one would expect both $SUPT_i$ and $RSAV_i$ to exhibit negative coefficients in **(7.2.1)**.

7.2.3 Measurement of Variables

For the empirical analysis we need to consider proxies for the variables in **(7.2.1)**. The dependent variable, β_{Yi}, is the estimated elasticity of consumption with respect to income. There are two models: one for the long run elasticity, with the coefficients obtained from the favoured cointegrating vectors reported in the Chapter 4; and one for the short run elasticity, being the sum of the parameters on the income growth terms from the preferred error correction models, reported in the Chapter 5.

The country-specific explanatory factors are averages of the variables for each country over the period 1960-1994. Precise definitions and sources are given in the data appendix (7.A).

Income growth may be approximated by both private disposable income growth and GDP growth. GDP growth is used by Modigliani (1990) and Jappelli and Pagano (1994) for the national saving to national income ratio. Since our investigation focuses upon the private sector, private income growth may be more appropriate.[8] We try both.

The central government fiscal surplus/deficit to GDP ratio is used to capture the impact of public saving upon private saving. A positive (negative) value denotes a surplus (deficit).

We have two proxies for credit constraints: the private sector credit to GDP ratio and the broad money (money plus quasi-money) to GDP ratio. The money to GDP ratio may reflect financial deregulation in a broad sense; however, it does not solely focus upon private sector credit conditions. In contrast, the private sector credit to GDP ratio does, though this variable is not without criticism. Jappelli and Pagano (1994) point out that this measure comprises credit available to both consumers and business. They note that, in some countries, the availability of credit to firms may be abundant while households may find loans difficult to secure. Unfortunately, superior measures such as consumer credit and the maximum loan to value ratio are not available with sufficient coverage to use here (see Jappelli and Pagano 1994).[9] We employ both the broad money and private sector credit to GDP ratios to gauge the effects of

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[8] Koskela and Viren (1989) use private disposable income growth in their cross-country analysis of the household saving ratio.

[9] Consumer credit would be a superior measure to private credit, however, it is less widely available. For example, Jappelli and Pagano (1994) report data for this variable for seventeen of the OECD countries used in our present study for the single year of 1980 and for far fewer countries for 1960 and 1970. Jappelli and Pagano (1994) also argue that such a variable need not reflect supply side constraints but may be demand determined. They suggest the use of an alternative supply side measure of credit constraints, being the maximum loan-to-value ratio (LTV). LTV is not subject to the problem of confusion over demand and supply side factors, it is a supply indicator, indicating the availability of credit to households: households must meet the down payment to obtain a mortgage regardless of their future ability to repay the loan. Meeting the down payment enforces saving reflecting a supply side constraint. Nevertheless, Jappelli and Pagano (1994) find that the LTV ratio has a strong positive correlation with the credit to income measures they use. These credit measures may be sufficiently supply determined to represent reasonable proxies of credit constraints. The LTV variable is not available over a sufficient time period or for an adequate number of countries to use here.

liquidity constraints.

Data on capital gains and losses for the twenty OECD countries is not available. The only proxy for which data is available is consumer price inflation. This variable may capture capital gains and loses on nominally fixed assets and, using this measure, suggests a negative relationship in **(7.2.1)**. However, we note that, beyond money holdings, one would additionally need the prices and quantities of assets to properly capture this effect.

We use the real interest rate, as defined in Chapters 3 and 6, to measure the real interest rate effect.

Income inequality is measured using the Gini coefficients reported in Atkinson (1995) - see Barrett and Pendakur (1995) for various means of this variable's construction. The larger is this coefficient the greater is the inequality, thus, this variable should be negatively related to the dependent variable in **(7.2.1)**. Because this data is only available for thirteen of the countries considered here its empirical implementation is limited to bilateral analysis and addition to favoured specifications on a sub-sample of observations.[10]

Per-capita income is measured using per-capita GDP in Geary-Khamis dollars, reported in Maddison (1995), to allow cross-country comparisons of living standards.[11] In addition to using this variable without transformation we also introduce its natural logarithm and its square root. These nonlinear transformations are to allow the elasticity parameter to decrease at a decreasing rate as the level of income increases. That is, a unit increase in the level of per-capita income causes less than a unit increase in the elasticity due to the logarithmic and/or square root transformation. Thus, the income elasticity need not become negative. One alternative would be

[10] The Gini coefficient is available in Atkinson 1995 p. 21 for Belgium, Finland, France, Germany, Ireland, Italy, the Netherlands, Norway, Spain, Sweden, Switzerland, the UK and the USA.

[11] Iceland's level of income in dollars was not available and was constructed using per-capita GDP in Kronur, the dollar-Kronur exchange rate and the US consumer price index. Some other adjustments were made to help achieve consistency with the GDP levels of other countries.

to use the log level of income *and* the squared value of the log-level. If the coefficient on the former is negative and the latter positive this would suggest the income elasticity falls initially and then rises as the level of income increases. If the eventual rise is modest this may *approximate* the nonlinear relationship between the *level* of income and the income elasticity. However, we have a preference for the logarithmic or square root forms because they do not imply an eventual rise.

Three measures of income uncertainty are considered: the rate of unemployment, its first difference and the absolute deviation of income from past trend - see Muellbauer, 1994, Muellbauer and Lattimore, 1995, and Malley and Moutos, 1996. The rate of unemployment *or* its first difference could be entered with the absolute deviation of income because they measure different aspects of income uncertainty - see Muellbauer, 1994. A larger value of each variable suggests greater uncertainty and so all are expected to be negatively correlated with the elasticity coefficient.

The length of retirement is measured as expected life length at birth (both sexes) minus the age of retirement (which is assumed to be 65 for all countries). We use two versions of this measure. The first is the average value for a particular country over the period 1960-1995. The second is the average value from 1950-1980. This effective lagging recognises that households only make expenditure decisions when they are economically active.

The support ratio is measured, following Miles and Patel (1996), as the number of the population of working age divided by the number of pensionable age. We take those aged 65 and over as of pensionable age. For the working age we consider two ranges, 15 to 64 and 20 to 64.

There are three measures for pre-retirement savers ($RSAV_i$), being the proportions of the total population in the age ranges 40 to 64, 45 to 64 and 50 to 64.

We have not distinguished between factors which influence short and long run elasticities although theory generally focuses upon the latter. We seek to determine whether the factors relevant for the long run are also pertinent for short term behaviour. We are not aware of any previous study which has attempted to separately identify short and long run cross-country variations. We will consider all the variables listed in **(7.2.1)** as potential influences for both and seek to provide initial insights into any differences or similarities.

7.2.4 Empirical Results

In our empirical analysis we employ the general-to-specific methodology in searching for a parsimonious multivariate form of **(7.2.1)**. All regressions are estimated with the OLS method. Since we only have twenty observations, we also estimate bivariate regressions to obtain initial insights into the theoretical plausibility and statistical significance of each factor. These insights are drawn upon in the model reduction process. Our favoured model is adopted on the basis of theoretical plausibility, best fit, and absence of misspecification. We only report regressions using the favoured proxy of each variable - being those measures which secure the most economically sensible and statistically valid results. Traditionally, heteroscedasticity is the primary problem in cross-section analysis. We therefore report t-ratios using White's (1980) heteroscedasticity consistent standard errors.

Table 7.1 reports the bivariate regression results for the long run elasticity of consumption with respect to income. The regressors are an intercept (INT) and an explanatory variable (EVAR) - given in the top row for each regression. The explanatory variables are defined in equation **(7.2.1)**. For those factors with more than one proxy the favoured measures are disposable income growth for GRTH, the proportion of the total population aged between 0 and 14 for DEP, the private credit to GDP ratio for CRED, the natural logarithm of per-capita income for f(INC), the rate of unemployment for UNCT, the expected retirement length averaged over the period 1960-1995 for RTRE, the age group 15 to 64 divided by those age 65 and over for SUPT and the proportion of the total population aged between 50 and 64 for RSAV. Unless otherwise specified these represent the favoured measures for all subsequent analysis. The reported statistics are the estimated coefficients with corresponding White's t-ratios given in brackets. The coefficient of determination adjusted for degrees of freedom (Adj R^2) is also reported. All regressions use twenty observations except that for income inequality, which uses thirteen observations (due to data constraints). Only income growth and the deficit to GDP ratio enter with the expected sign

TABLE 7.1: Bivariate Cross Country Models of the Long Run Income Elasticity of Consumption, Equation (7.2.1)

	GRTH	DEF	DEP	CRED	GAIN	r	INEQ	f(INC)	UNCT	RTRE	SUPT	RSAV
INT	1.470	1.197	1.171	0.894	1.017	1.015	1.143	0.258	1.080	0.567	0.810	0.952
	(8.904)	(17.60)	(3.168)	(8.771)	(15.06)	(22.05)	(3.093)	(0.658)	(11.26)	(1.927)	(3.469)	(2.631)
EVAR	-17.48	6.232	-0.676	0.229	-0.047	-0.049	-0.005	0.301	-0.013	0.050	1.095	0.400
	(-2.786)	(3.370)	(-0.474)	(1.379)	(-0.075)	(-0.026)	(-0.375)	(1.876)	(-0.890)	(1.461)	(0.799)	(0.157)
Adj R^2	0.334	0.454	-0.046	0.017	-0.055	-0.056	-0.078	0.055	-0.029	0.023	-0.029	-0.055

Table 7.1 notes. This table reports the results of bivariate cross section regressions for 20 countries (The regression including income inequality only uses 13 observations). The dependent variable is the long run consumption elasticity with respect to income. The regressors are an intercept (INT) and an explanatory variable (EVAR). The explanatory variable for each regression is given in the top row. The variables are income growth (GRTH), the fiscal deficit to GDP ratio (DEF), the dependency ratio (DEP), the degree of credit constraints (CRED), capital gains and losses (GAIN), the real interest rate (r), the degree of income inequality (INEQ), a function of the level of income (f(INC)), income uncertainty (UNCT), the expected length of retirement (RTRE), the support ratio (SUPT) and the proportion of the total population comprised of pre-retirement savers (RSAV). The favoured proxies are disposable income growth for GRTH, the proportion of the total population aged between 0 and 14 for DEP, the private credit to GDP ratio for CRED, the natural logarithm of income for f(INC), the rate of unemployment for UNCT, the expected retirement length averaged over the period 1960-1995 for RTRE, the age group 15 to 64 divided by 65 and over for SUPT and the proportion of the total population aged between 50 and 64 for RSAV. GAIN is proxied with inflation so is expected to exhibit a negative coefficient. The reported statistics are the estimated coefficients with corresponding t-ratios given in brackets, based upon White's (1980) heteroscedasticity consistent standard errors. The coefficient of determination adjusted for degrees of freedom (Adj R^2) is also reported. The approximate critical values for the t-ratios, assuming twenty degrees of freedom, are: ±2.85 (1 percent level), ±2.09 (5 percent level) and ±1.725 (10 percent level).

and are statistically significant. Interestingly, these are two of the four explanatory variables postulated in Jappelli and Pagano's (1994) extended version of Modigliani's (1990) model for national savings rates and the fiscal variable is that emphasised as a primary determinant of private savings by Pesaran, Haque and Sharma (1999).

Table 7.2 reports multivariate models for the long run income elasticity. All statistics and variables are the same as those reported for Table 7.1 with the addition of the probability values for the significance of the regression Pr[FR^2], first order serial correlation Pr[FSC1], non-linear functional form Pr[FFF1], non-normally distributed residuals Pr[χ^2N2] and heteroscedasticity Pr[FH1].[12] A probability value exceeding 0.05 indicates significant explanatory power and/or misspecification, depending upon context, at the five percent level.

[12] These are the standard misspecification tests automatically produced by Microfit 3.22.

TABLE 7.2: Multivariate Cross-Country Models of the Long Run Income Elasticity of Consumption, Equation (7.2.1)

	7.2.1a	**7.2.1b**	**7.2.1c**	**7.2.1d**	**7.2.1e**	**7.2.1f**	**7.2.1g**
Intercept	2.781	2.509	1.390	1.347	1.406	3.545	3.634
	(7.236)	(5.723)	(10.580)	(8.861)	(12.517)	(8.376)	(6.909)
GRTH	-17.866	-13.344	-10.453	-18.096	-9.785	-38.853	-41.903
	(-3.624)	(-2.684)	**(-1.873)**	(-3.309)	(-2.238)	(-5.640)	(-6.212)
DEF	5.382	6.502	4.325		4.665	1.516	
	(2.736)	(4.661)	**(1.610)**		(2.403)	**(0.730)**	
CRED	0.235		0.046	0.267		0.436	0.592
	(1.541)		**(0.258)**	(2.730)		(4.130)	(3.803)
f(INC)	-0.504	-0.380				-0.597	-0.601
	(-3.931)	(-2.651)				(-3.332)	(-3.044)
INEQ						-0.009	-0.013
						(-1.355)	(-2.909)
Adj R^2	0.606	0.587	0.489	0.399	0.517	0.824	0.831
Pr[FR^2]	[0.001]	[0.001]	[0.003]	[0.005]	[0.001]	[0.002]	[0.001]
Pr[FSC1]	[0.901]	[0.842]	[0.586]	[0.531]	[0.602]	[0.518]	[0.322]
Pr[FFF1]	[0.112]	[0.244]	**[0.033]**	[0.151]	**[0.037]**	[0.491]	[0.891]
Pr[χ^2N2]	[0.614]	[0.694]	[0.614]	[0.709]	[0.587]	[0.415]	[0.565]
Pr[FH1]	[0.354]	[0.103]	[0.754]	[0.769]	[0.614]	[0.072]	[0.059]

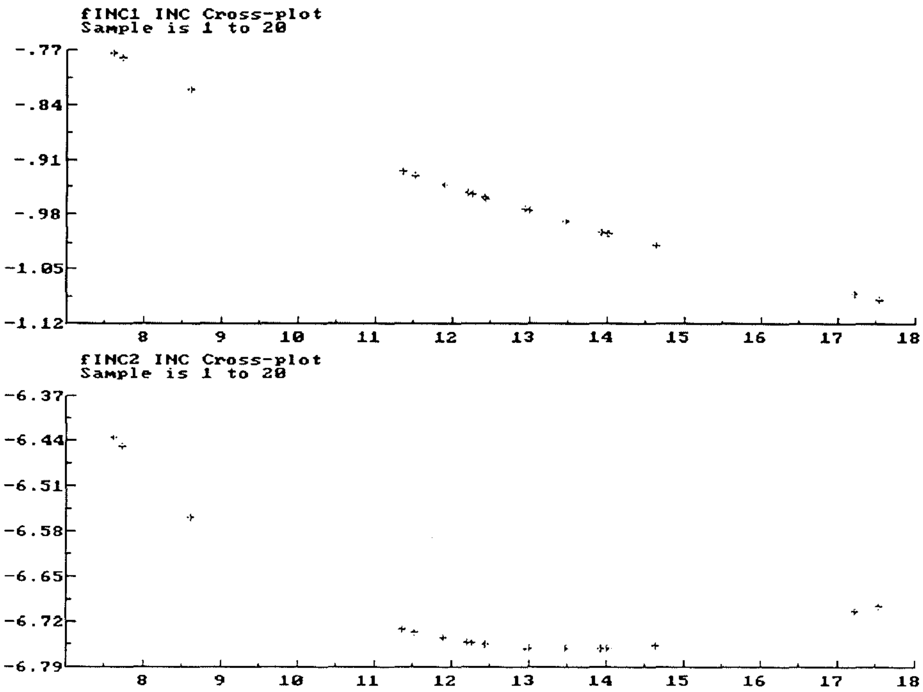
Table 7.2 notes. All statistics and variables are the same as those reported in Table 7.1 with the addition of the probability values for the statistical significance of the regression Pr[FR^2], first order serial correlation Pr[FSC1], non-linear functional form Pr[FFF1], non-normally distributed residuals Pr[χ^2N2] and heteroscedasticity Pr[FH1]. Bold emphasis indicates an insignificant regressor (and regression) and significant misspecification (depending upon context) at the 5 percent level. All regressions use 20 observations except those incorporating income inequality which employ 13.

Table 7.2 reports seven multivariate models for the estimated income elasticity which contain various combinations of, what appear to be, the five main explanatory factors: income growth, the fiscal surplus/deficit, private sector credit, the log of per-capita income (ln(INC)) and income inequality. All seven specifications exhibit statistically significant explanatory power and are free from evident misspecification at the five percent level except equations **7.2.1c** and **7.2.1e** which feature statistically significant evidence of nonlinear functional form at the five percent (but not one percent) level. We suggest that our inferences are legitimate except, perhaps, for these two specifications.

Equation **7.2.1a** includes GRTH, DEF, CRED and ln(INC). All exhibit the expected sign and are statistically significant at the five percent level, except CRED which is not significant (the removal of CRED causes the adjusted R^2 to drop marginally from 0.606 to 0.587). Excluding CRED from **7.2.1a** yields equation **7.2.1b**. All remaining variables are statistically significant and *correctly* signed. What is clear from these two regressions is that income growth is negatively associated with the income elasticity, as implied by Modigliani's (1986 and 1990) LCH and Brown's (1952) version of the RIH. The positive sign of the coefficient on DEF suggests that private saving rises when public saving falls, suggesting some degree of Ricardian equivalence. However, we are unable to assess whether the rise in private saving is of the same magnitude as the fall in public saving because our dependent variable is the elasticity of consumption with respect to income and not the ratio of savings to GDP. That per-capita income is negatively correlated with the income elasticity is consistent with Keynes's (1936) conjecture but inconsistent with the implications of Modigliani's LCH. This nonlinear relation allows the income elasticity to decrease at a decreasing rate as the level of per-capita income rises. The plot in the top half of Figure 7.5 graphs the contribution that the nonlinear function of per-capita income has on the long run income elasticity, fINC1, against the *level* of income. (This contribution, fINC1, is simply the estimated parameter on f(INC) in Equation **7.2.1b** multiplied by lnINC, that is: $fINC1_i = -0.38035 lnINC_i$). Figure 7.5 shows that as per-capita income increases the income elasticity decreases at a decreasing rate.[13]

[13] The square of the log-level of per-capita income enters with statistical significance when added to **7.2.1b**. However, plotting the nonlinear function of the log-level of income $\{fINC2_i = -5.1901(lnINC_i) + 0.99591(lnINC_i)^2\}$ against the level of income (INC_i) suggests that

FIGURE 7.5: Partial Non-Linear Relationships Between Per-Capita Income (INC) and the Long Run Income Elasticity (fINC1 and fINC2)



Due to the possible controversy surrounding the inclusion of such a variable and its statistical insignificance in the bivariate regressions reported in Table 7.1, we consider models excluding the log-level of income. Equation **7.2.1c** is obtained by excluding the level of income from **7.2.1a**. All variables are *correctly* signed, however, they are also statistically insignificant (except for the intercept). Equation **7.2.1d** excludes DEF from **7.2.1c**. Both GRTH and CRED are correctly signed and statistically significant at the five percent level. Replacing CRED with DEF in **7.2.1d** gives equation **7.2.1e**. Once again, both variables are correctly signed and

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its effect was to initially reduce the income elasticity and then increase it - see bottom half of Figure 7.5. We consider this eventual increase to be implausible so we do not favour the model incorporating this squared term - further, the introduction of this term induces evidence of misspecified functional form. However, we believe the squared term's significance suggests that the log-level of income does not, on its own, depict a fast enough slow down in the decrease of the income elasticity, rather than an eventual increase. We therefore believe that the evidence suggests that the income elasticity decreases at a decreasing rate, but that the precise form of the nonlinear function has not been found. Pursuing this form is inhibited by the sample of twenty observations. We suggest further investigation of this issue, using a larger sample, is warranted.

statistically significant at the five percent level. Comparing equations **7.2.1c**, **7.2.1d** and **7.2.1e**, we see that the adjusted R^2 dramatically falls from 0.489 to 0.399 with the exclusion of DEF but rises from 0.489 to 0.517 when CRED is removed. We therefore favour equation **7.2.1e** from these models, confirming our previous results suggesting the importance of GRTH and DEF. However, exclusion of the log of income, f(INC), causes a large (7.7%) fall in the adjusted R^2, from 0.587 for model **7.2.1b** to 0.517 for model **7.2.1e**, suggesting that this variable is also an important explanatory factor of the income elasticity.

Equation **7.2.1f** includes income inequality in the model. We find, in contrast to the corresponding bivariate regression results, that this variable is statistically significant and negative, which is consistent with the RIH and Borooah and Sharpe (1986). GRTH, CRED and f(INC) are also statistically significant and *correctly* signed, while DEF is highly insignificant. Excluding DEF gives equation **7.2.1g**. All retained variables are statistically significant at the 1% level and the fit of the equation rises to a high level, suggesting 83.1% explanatory power. However, this high value relative to previous regressions may be due, at least in part, to the use of 13 (rather than 20) observations. This model confirms the inferences drawn from the previous regressions regarding GRTH and f(INC) whilst suggesting an additional role for income inequality. However, unlike previous regressions it indicates that CRED is an important explanatory factor, which is implied by Jappelli and Pagano (1994), and that there is no role for DEF, which suggests that Ricardian equivalence does not hold.

Overall, our results suggest that income growth has a clear negative influence upon the income elasticity and is consistent with the LCH theory of Modigliani (1986 and 1990). The log of per-capita income also negatively influences the income elasticity. This might be regarded highly controversial if it were a simple linear relationship which implied that the elasticity would eventually become negative. However, the nonlinear logarithmic form allows the elasticity to decrease at a decreasing rate as the level of income rises which does not necessitate that the elasticity becomes negative. We argue that this is quite a plausible effect which supports Keynes's (1936) suggestion, if it contradicts an implication of Modigliani's (1986 and 1990) LCH. There is also some evidence indicating that increased income inequality reduces the income elasticity. The fiscal surplus/deficit generally exerts a positive and statistically significant influence on the income elasticity indicating some degree of Ricardian equivalence, however, we are unable to determine whether it is complete or not. This is consistent with previous work. There is also some evidence which supports the amount of credit available to the private sector having a positive impact upon the long run income elasticity.

TABLE 7.3: Bivariate Cross-Country Models of the Short Run Income Elasticity of Consumption, Equation (7.2.1)

	GRTH	DEF	DEP	CRED	GAIN	r	INEQ	INC	UNCT	RTRE	SUPT	RSAV
INT	0.732	0.681	0.448	0.519	0.725	0.511	0.561	0.508	0.586	1.091	0.772	0.815
	(4.138)	(7.646)	(1.049)	(3.923)	(10.255)	(9.023)	(1.325)	(0.751)	(4.788)	(3.305)	(2.765)	(1.561)
EVAR	-4.975	2.657	0.665	0.160	-1.819	7.276	0.002	0.038	0.847	-0.055	-0.908	-1.368
	(-0.758)	(0.986)	(0.362)	(0.745)	(-2.580)	(2.889)	(0.140)	(0.143)	(0.149)	(-1.529)	(-0.612)	(-0.409)
Adj R^2	-0.032	-0.014	-0.049	-0.029	0.032	0.234	-0.089	-0.054	-0.054	0.015	-0.042	-0.047

Table 7.3 notes. This table reports the results of bivariate cross section regressions for 20 countries. The dependent variable is the short run consumption elasticity with respect to income. The regressors are an intercept (INT) and a variable (EVAR). The explanatory variable for each regression is given in the top row. The variables and favoured proxies are the same as those defined in Tables 7.1 and 7.2, except the absolute deviation of income is used for UNCT. The reported statistics are the same as those defined in Table 7.1. The regression for income inequality only uses 13 observations.

Table 7.3 reports the bivariate regression results for the short run income elasticity. The variables and statistics are the same as those used in Table 7.1, except the absolute deviation of income (rather than the level of unemployment) is used to proxy income uncertainty. Only two variables are significant, capital gains (inflation) and real interest rates, and feature signs which are economically justifiable. The positive real interest rates effect suggests that the income effect dominates the substitution effect: increases in capital income raise the short run income elasticity. However, since this effect is unusual and Keynes (1936) argued that a positive interest rate effect was likely to prevail in the long rather than short run, we have doubts over the validity of this effect for the short run elasticity. The negative inflation effect is weak, the adjusted R^2 of this regression is 3.2%, so we do not draw strong inferences supporting a significant capital gains effect. Further, we are unable to construct satisfactory multivariate models which improve upon these bivariate results. Thus, we tentatively suggest that, to the extent that there are systematic variations in short run income elasticities, they appear to be positively related to real interest rates (income effect) and negatively related to inflation (capital gains effect). However, the evidence presented in favour of such short run effects is weak. We argue that no substantive explanation of the short run income elasticity has been revealed.

7.3 Explaining Cross-Country Differences in the Response of Consumption to Inflation

To the extent that inflation is approximating wealth effects, one might consider the variation in the response of consumption to inflation to be related to factors which affect the MPC out of assets. We consider such wealth effects as potential explanatory factors for cross-country differences in the response of consumption to inflation, assuming an inverse relation between inflation and asset effects.[14]

7.3.1 Theoretical Considerations

According to the LCH, older households will have a larger MPC out of assets relative to younger ones, especially the retired who are depicted as consuming totally out of accumulated saving. Hence, one might expect a positive (negative) association between the proportion of the retired population and the MPC out of wealth (inflation). However, the presence of a bequest motive may reduce the size of such a demographic effect and, in the extreme situation where parents obtain the same marginal utility from assets bequeathed to their own children as from their own consumption (as if infinitely lived), there will be no such effect. Such complete intergenerational altruism may be unlikely because parents are aware, for example, that the young, in a growing economy, have better income prospects: parents apply a larger discount factor to their children's utility relative to their own.

Uncertainty over life length can also affect the MPC out of wealth. Greater uncertainty and/or

[14] If inflation has an impact beyond approximating wealth one would need to explain cross-country differences in the measurement of income (see Hendry and Ungern Sternberg 1981) or consumers mistaking absolute price increases as relative rises (see Deaton 1977). The former is related to wealth effects, being the unaccounted inflationary losses on liquid (monetary or total) assets, so its variation may be reasonably well approximated by the factors cited above. How one explains the variation in the latter with observable data is not obvious and is regarded beyond the scope of this Chapter. We concentrate on variations of inflation as a proxy for wealth effects.

survival probability suggests a lower (higher) MPC out of assets (inflation). Indeed, the expected length of retirement may be thought to be negatively (positively) related to the wealth (inflation) elasticity of consumption.

Muellbauer and Lattimore (1995) expand upon the uncertainty over life length theme by suggesting an inverse relationship between remaining planning horizon and age. That is, the young have a small MPC out of assets, those in pre-retirement have a medium MPC out of assets while the retired have a large MPC out of assets. Thus, one might expect a negative (positive) relationship between the proportion of the population who are young (YNG_i) and the wealth (inflation) elasticity. Conversely, and as suggested above, a positive (negative) correlation may be expected between the retired proportion of the population ($65+_i$) and the elasticity out of assets (inflation). The direction of correlation of the middle aged with the MPC out of wealth is unclear and is further complicated by the needs of children, in both infancy and when attending college (see Banks *et al* 1994). However, one might expect assets to be run down when children are dependents suggesting a positive (negative) relationship between the dependency ratio and the MPC out of wealth (inflation). A reverse relationship may be expected if parents are saving for the education of their children. Further, once children have left home, the household may save for its retirement, wishing to accumulate, rather than run down, assets. Thus, the proportion of the population comprised of pre-retirement savers may be negatively (positively) related to the wealth (inflation) elasticity.

The precautionary motive derived from income uncertainty may reduce (raise) expenditure out of assets (inflation).

The less binding are credit constraints the more fungible (spendable) is wealth. For example, credit constraints may prevent consumers borrowing upon the basis of illiquid assets such as housing, frustrating their desired consumption plans. Relaxing such constraints would release this pent up demand, and has been suggested to be the cause of the UK consumer boom in the mid/late 1980s - see, for example, Miles (1992). Economies with greater access to credit may be expected to feature a larger (lower) MPC out of wealth (inflation).

Our general model for the estimated *inflation* elasticity, β_{Ii}, is given by **(7.3.1)** below.

$$\beta_{Ii} = \underset{-}{f(\text{CRED}_i}, \underset{+}{\text{UNCT}_i}, \underset{+}{\text{RTRE}}, \underset{-(+)}{\text{DEP}_i}, \underset{+}{\text{YNG}_i}, \underset{+}{\text{RSAV}_i}, \underset{-}{65+_i}) \qquad \textbf{(7.3.1)}$$

where β_{Ii} is the estimated inflation elasticity parameter for country i.

7.3.2 Empirical Analysis

Our modelling approach is similar to that used for the income elasticities. Bivariate regressions are run for each variable in **(7.3.1)** to obtain some initial insights into each variable's simple correlation with the estimated inflation elasticity. Multivariate models yielding partial correlations are developed based upon the general-to-specific methodology whilst bearing in mind the results of the bivariate analysis. Favoured specifications are based upon considerations of fit, theoretical plausibility and misspecification testing. This modelling strategy is applied to both long and short run inflation elasticities.

Table 7.4 reports the bivariate results for the long run inflation elasticities. The reported statistics and favoured proxies are the same as those used for Table 7.1, except the absolute income deviation (rather than the level of unemployment) proxies income uncertainty.[15] Initial regressions using all twenty countries revealed severe departures from normality due to a large outlying observation for Italy, identified in Figure 7.2. We therefore exclude the Italian observation from our regressions, restricting the sample to 19 observations, to secure valid inference. The variables CRED, YNG and 65+ feature theoretically plausible signs and are statistically significant at the 1% level. All other variables are statistically insignificant.

Table 7.5 presents the only satisfactory multivariate model that could be secured for the long run inflation elasticity. The definitions of the variables used are the same as for Tables 7.4 and the statistics are the same as those employed in Table 7.2. As before the regression uses 19

[15] The favoured proxies for YNG and RSAV are the proportions of the population aged 15-34 and 50-64, respectively.

TABLE 7.4: Bivariate Cross-Country Models of the Long Run Inflation Elasticity of Consumption, Equation (7.3.1)

	CRED	UNCT	RTRE	DEP	YNG	RSAV	65+
INT	0.787	-0.777	-0.011	1.176	8.980	-3.032	-2.812
	(1.975)	(-1.949)	(-0.009)	(0.690)	(3.108)	(-1.796)	(-3.411)
EVAR	-1.904	28.186	-0.024	-5.999	-30.678	18.150	21.444
	(-3.595)	(1.434)	(-0.159)	(-0.877)	(-3.219)	(1.622)	(2.907)
Adj R^2	0.236	0.056	-0.058	-0.015	0.209	0.055	0.195

Table 7.4 notes. This table reports the results of bivariate cross section regressions for 19 countries. The Italian observations is omitted because it consistently causes a severe departure from normality when included in the data set. The dependent variable is the long run consumption elasticity with respect to inflation. The regressors are an intercept (INT) and an explanatory variable (EVAR) . The explanatory variable for each regression is given in the top row. The variables are the private credit to GDP ratio (CRED), income uncertainty (UNCT), the length of retirement (RTRE), the proportion of the population who are young, aged 15-34, (YNG), the pre-retirement savers in the population, aged 50-64, (RSAV), and the proportion of the population aged 65 and over (65+). The reported statistics are the same as for Table 7.1.

TABLE 7.5: Multivariate Cross-Country Models of the Long Run Inflation Elasticity of Consumption, Equation (7.3.1)

Int	CRED	UNCT	DEP	Adj R^2	Pr[FR^2]	Pr[FSC1]	Pr[FFF1]	Pr[χ^2N2]	Pr[FH1]
4.792	-2.334	39.211	-19.510	0.551	[0.002]	[0.936]	[0.860]	[0.582]	[0.885]
(2.961)	(-3.788)	(2.448)	(-2.861)						

Table 7.5 notes. As for Table 7.4 the reported multivariate regressions use 19 observations - Italy is excluded. All statistics are the same as those defined in Table 7.2 and the variables are the same as those given in Table 7.3.

TABLE 7.6: Bivariate Cross-Country Models of the Short Run Inflation Elasticity of Consumption, Equation (7.3.1)

	CRED	UNCT	RTRE	DEP	RSAV	YNG	65+
INT	-0.011	-0.173	-0.110	-0.309	0.054	0.415	-0.195
	(-0.141)	(-2.828)	(-0.495)	(-1.638)	(0.203)	(0.527)	(-0.819)
EVAR	-0.153	4.199	0.002	0.940	-0.935	-1.688	0.858
	(-1.082)	(2.084)	(0.085)	(1.264)	(-0.537)	(-0.637)	(0.460)
Adj R^2	0.015	0.089	-0.055	-0.016	-0.044	-0.026	-0.041

Table 7.6 notes. This table reports the results of bivariate cross section regressions for 20 countries. The dependent variable is the short run consumption elasticity with respect to inflation. The regressors are an intercept (INT) and a variable (EVAR). The explanatory variable for each regression is given in the top row. The variables are the same as those defined in Tables 7.4 and 7.5. The reported statistics are the same as those defined in Table 7.1.

observations. There is no evident misspecification according to the reported diagnostics suggesting inference is valid. UNCT exhibits a positive and statistically significant impact upon the inflation elasticity while CRED and DEP feature negative and significant correlations.[16] The model provides significant explanatory power with a 55.1% fit. The estimated coefficients' signs are all consistent with cross-country variations expected if inflation were approximating wealth effects in the long run consumption function. Thus, we suggest that this provides evidence favouring this interpretation of inflation's role in consumption functions.[17]

Table 7.6 presents bivariate regressions for the short run inflation elasticity.[18] All variables and statistics are the same as those defined in Table 7.4. None of the variables enter with statistical significance and no multivariate models exhibiting significant relationships could be developed. Therefore, we find no explanation for the cross-country variations in the short run inflation elasticities.

7.4 Explaining Cross-Country Differences in the Speed of Adjustment Towards Equilibrium

The speed of adjustment towards equilibrium may be determined by the ability of consumers to change their consumption. We are aware of no theory explaining variations in speed of adjustment and simply offer some conjecture and empirical evidence. The main explanatory factors we postulate are habits and adjustment costs and the availability of credit.

A country where consumers are more habitual (higher adjustment costs) in their spending

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[16] It is noticeable that the demographic factor relevant in the multivariate regression (DEP) is different from those suggested in the bivariate regressions (YNG and 65+). Further, UNCT, which was not statistically significant in the bivariate regressions is in the favoured multivariate model.

[17] This is consistent with Lattimore's (1994) finding that inflation has no role in an Australian consumption function when well defined wealth variables are used.

[18] The majority of these elasticities are zero because there are no short run inflation effects in many of the countries' error correction models.

patterns may be expected to adjust expenditures more slowly than an economy with lower adjustment costs. We have no direct habit variable but suggest that an economy's adjustment costs will likely be related to country specific institutional factors. One such country specific factor that we can proxy is the financial development of an economy.

The degree of credit constraints may determine the speed of adjustment. For example, binding credit constraints may prevent rapid adjustments towards equilibrium. On the other hand, greater availability of credit may enable consumers to adjust more quickly towards their optimal level of consumption. This would suggest a negative association between the value of the adjustment coefficient and the degree of financial liberalisation.[19]

We can approximate the availability of credit by many factors. For example, we may expect the adjustment coefficient to be negatively associated with the availability of credit and positively related to the rate of unemployment (U_i) - the greater the unemployment the greater the number of constrained consumers. We also consider the change in the rate of unemployment to proxy this effect. A negative relationship may also be expected with the rate of interest (higher interest rates suggest larger costs for borrowing) and the proportion of the population who are young (assuming the young have less access to credit).

The general model for the adjustment coefficient is:

$$\alpha_i = f(\underset{-}{CRED_i}, \underset{+}{U_i}, \underset{+}{\Delta U_i}, \underset{+}{r_i}, \underset{+}{YNG_i}) \qquad \textbf{(7.4.1)}$$

Since the estimated adjustment coefficients for Sweden and Switzerland are positive, which is inconsistent with error correction behaviour, we also run regressions excluding these two countries from the sample.

Table 7.7 reports bivariate results for equation **(7.4.1)** using the full twenty observations. The

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[19] For valid error correction behaviour the adjustment coefficient should be negative which means that larger (less negative) values imply faster adjustment.

TABLE 7.7: Bivariate Cross-Country Models of the Adjustment Coefficient, Equation (7.4.1)

	CRED	U	ΔU	r	YNG
INT	-0.323	-0.122	-0.194	-0.239	0.830
	(-3.569)	(-1.271)	(-3.625)	(-5.353)	(1.137)
EVAR	0.169	-0.023	-0.199	0.269	-3.551
	(1.197)	(-1.412)	(-0.786)	(0.116)	(-1.478)
Adj R^2	0.003	0.061	-0.018	-0.055	0.034

Table 7.7 notes. This table reports the results of bivariate cross section regressions for 20 countries. The dependent variable is the adjustment coefficient. The regressors are an intercept (INT) and a variable (EVAR). The explanatory variable for each regression is given in the top row. The variables are as defined in previous Tables with U and ΔU being the rate of unemployment and its change, respectively.

TABLE 7.8: Bivariate Cross-Country Models of the Adjustment Coefficient (continued), Equation (7.4.1)

	CRED	U	ΔU	r	YNG
INT	-0.271	-0.211	-0.193	0.278	0.408
	(-2.535)	(-2.445)	(-4.375)	(-9.192)	(0.579)
EVAR	0.015	-0.010	-0.366	1.066	-2.234
	(0.081)	(-0.633)	(-2.410)	(0.612)	(-0.965)
Adj R^2	-0.062	-0.035	0.089	-0.046	-0.021

Table 7.8 notes. This table reports the results of bivariate cross section regressions for 18 observations on the adjustment coefficient. The Swedish and Swiss adjustment coefficients are excluded because they feature theoretically indefensible positive signs. Variables and statistics are the same as for Table 7.7.

variables and statitistics reported are the same as those presented in previous tables. None of the variables enter with statistical significance. Multivariate models with significant explanatory power could not be developed. To consider whether this was due to the *implausible* Swedish and Swiss observations we reestimated the bivariate regressions excluding these two countries. Table 7.8 reports these regressions using the remaining eighteen observations. The only variable which is statistically significant is the change in unemployment, however, it enters with a theoretically unanticipated negative sign. As before, plausible multivariate models which provide significant explanatory power could not be developed. We conclude that the evidence suggests no systematic relationship between the adjustment coefficient and credit constraints.

7.5 Explaining Cross-Country Differences in the Proportion of Current Income Consumers

Evident excess sensitivity has led to the modification of the rational expectations permanent income/life cycle hypotheses (REPIH/RELCH) to allow for current income consumers. Originally this excess sensitivity was considered due to the presence of liquidity constraints. Recent literature has suggested other potential reasons as well. "The Euler equation approach has generated a large empirical literature, much of which has dealt with the issue of whether consumption depends on predictable changes in current income. The weight of existing empirical results indicates that it does, although the reason for this dependence - whether it reflects liquidity constraints or precautionary saving - remains in dispute." (Bayoumi and Masson, 1998, p. 1035). Indeed, Hahm and Steigerwald (1999) present evidence which suggests that income uncertainty, operating through precautionary saving, partially explains the excess sensitivity of consumption to current income.[20] We will consider whether the excess sensitivity found for OECD countries is due to liquidity constraints and/or precautionary saving.

Many studies which estimate models allowing for current income consumers have sought to determine whether this proportion reflects the degree of liquidity constraints (see, for examples, Jappelli and Pagano, 1989; Campbell and Mankiw, 1991; and Jin, 1994). Jappelli and Pagano (1989) present tentative evidence suggesting that the countries with the lowest proportion of current income consumers also exhibit the largest levels of consumer credit (such as Sweden and the USA). Campbell and Mankiw (1991) also assert that their estimated proportions of current income consumers are smaller for economies with better developed consumer credit markets. This suggests that the availability of consumer credit is negatively related to the proportion of current income consumers.

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[20] Hahm and Steigerwald (1999) also argue that durable consumption may be particularly sensitive to consumer sentiment and, therefore, income uncertainty. Since we use total consumer expenditure we might expect income uncertainty to influence the cross-country variation in the proportion of current income consumers.

Jin (1994) suggests the following potential explanatory factors for the cross-country variation in the proportion of current income consumers. Economies with higher unemployment rates are likely to feature more people unable to access capital markets. Higher expected income growth suggests a steeper earnings profile through time, leading to people being liquidity constrained for longer because their actual income is lower than their optimal income level for longer. The young will likely be more liquidity constrained suggesting the larger is this proportion of the population in this age group the greater is the liquidity constraint. Economies with faster population growth rates will feature a larger proportion of young people and should, therefore, be subject to more binding liquidity constraints. Countries with lower savings (rates) are likely to be more liquidity constrained than those with higher savings. While high interest rate economies will hinder borrowing.

Jin (1994) produces estimates of the proportions of current income consumers for nineteen countries, over the period 1965-1988, using an income measure which incorporates both private and public sector income and total expenditure measures consumption. Using bivariate cross-country regressions, Jin (1994) finds that the proportion of current income consumers only features a statistically significant and theoretically expected relationship with the rate of unemployment and the savings rate. No significant relationship is found between this proportion and population growth, the fraction of the population who are young, income growth or the real rate of interest.[21] More reliable estimation is believed to be obtained using a pooled regression.[22] In this regression all the variables considered by Jin (1994) are significant, except interest rates. Income growth, unemployment and the proportion of the economically active population who are young are positively related with the proportion of current income consumers whilst

[21] These results are based upon six separate regressions of the estimated coefficient for the proportion of current income consumers against the sample mean (for each country) of each variable.

[22] Jin(1994) uses a pooled regression by substituting $\pi_i = b_0 + \Sigma b_i Z_i$, i=1,2...6; into his modified REPIH/RELCH model, which features a nonlinear semi-logarithmic form, to yield: $\Delta \ln C_{it} = \mu_i + [b_0 + \Sigma b_i Z_i](Y_{it-1}/C_{it-1})\Delta \ln Y_t - \mu[b_0 + \Sigma b_i Z_i](Y_{it-1}/C_{it-1}) + z_{it}$, where Z_i denotes the sample mean of each of the six variables used to proxy credit constraints. This is estimated with non-linear three stage least squares using the second lag of consumption growth from each country and a constant as instruments.

population growth and the savings rate exhibit a negative association. This is argued to confirm that liquidity constraints explain the variation in the proportion of current income consumers.[23] Thus, Jin (1994) provides evidence which, in general, supports the REPIH/RELCH, modified to allow for liquidity constrained consumers.

Although Acemoglu and Scott (1994) find that precautionary saving rather than liquidity constraints explain the rejection of the REPIH/RELCH for the UK, we are not aware of any previous cross-country analyses of the relationship between the proportion of current income consumers and income uncertainty. To measure income uncertainty we follow Muellbauer (1994) and Malley and Moutos (1996) by using the rate of unemployment, its difference and the absolute deviation of income. Larger values of all three measures indicate greater income uncertainty. If greater income uncertainty reduces consumers' confidence about expected future income, they may prefer to base their consumption decisions on current rather than future income: greater uncertainty raises the proportion of current income consumers. Thus, one would expect a positive association between the proportion of current income consumers and the specified measures of uncertainty. An alternative hypothesis is that income uncertainty may reduce the *expenditure* of current income consumers (allowing them to save), which is what π_i measures in our model.[24] Under this hypothesis, our measures of income uncertainty would be negatively correlated with the proportion of current income consumers.

We base our model of potential explanatory factors on the liquidity constraint variables employed by Jappelli and Pagano (1989) and Jin (1994) and on income uncertainty measures used by, for example, Muellbauer (1994) and Malley and Moutos (1996). However, we note that unemployment (and its change) may approximate both liquidity constraints and income

[23] All variables exhibit the correct sign except population growth which should be positive. It is argued that this effect should be considered in combination with the proportion of the population who are young, however, it is not obvious that their combined impact is positive, so an anomaly remains.

[24] This would suggest the relaxation of the assumption that current income consumers' consumption must equal their income each period in the derivation of the modified rational expectations model - at least across countries. This could easily be done by relaxing the assumption that current income consumers are not allowed to accrue precautionary savings.

uncertainty. We expect the proportion of current income consumers (π_i) to be negatively related to private sector debt to GDP ratio ($CRED_i$) and the saving rate (SY_i). We also expect a positive or negative association with the rate of unemployment (U_i), and its difference (ΔU_i) and the absolute deviation of income (AD_i). A positive correlation is anticipated with (disposable) income growth ($GRTH_i$), the proportion of the *total* population aged between 15 and 34 (YNG_i), population growth ($GPOP_i$) and real interest rates (r_i).[25] Equation **(7.5.1)** summarises this model:

$$\pi_i = f(\underset{-}{CRED_i},\ \underset{-}{SY_i},\ \underset{\pm}{U_i},\ \underset{\pm}{\Delta U_i},\ \underset{\pm}{AD_i},\ \underset{+}{GRTH_i},\ \underset{+}{YNG_i},\ \underset{+}{GPOP_i},\ \underset{+}{r_i}). \qquad \textbf{(7.5.1)}$$

In Chapter 6 we produced two sets of estimates for π_i: one set using Generalised Methods of Moments (GMM) which implicitly accounted for durability and another employing Instrumental Variables (IV) which explicitly introduced moving average error terms to allow for durables. The estimates produced by these two methods are similar so we apply **(7.5.1)** to both sets of estimates.

IV estimates of bivariate models are presented in Table 7.9. None of the variables form statistically significant bivariate relationships (at the five percent level) with the proportion of current income consumers, although YNG is *correctly* signed and statistically significant at the ten percent level. The corresponding bivariate models for the GMM estimates are reported in Table 7.10. YNG enters with the expected positive sign and is statistically significant at the five percent level while GPOP is also correctly signed though only significant at the ten percent level. CRED is statistically significant but features a positive sign which is not consistent with increased credit *reducing* the proportion of constrained consumers.[26]

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[25] All variables proxy liquidity constraints except AD_i which captures income uncertainty effects. U_i and ΔU_i proxy both liquidity constraints and income uncertainty effects. To be consistent with the liquidity constraint interpretation these variables should be positively related with the income elasticity, whereas, both a positive and negative relation is consonant with income uncertainty effects.

[26] This positive relationship is consistent with increased credit reducing enforced saving and increasing consumption out of aggregate income. Such an interpretation would suggest that the estimated parameter represents an income elasticity rather than the proportion of current income consumers. However, because neither the bivariate IV nor the nonlinear three stage least

TABLE 7.9: Bivariate Cross-Country Models of Current Income Consumers, IV Estimates, Equation (7.5.1)

	CRED	SY	U	ΔU	AD	GRTH	YNG	GPOP	r
INT	0.646	0.747	0.785	0.753	0.780	0.614	-0.378	0.674	0.742
	(8.325)	(10.173)	(13.657)	(18.448)	(13.253)	(7.394)	(-0.623)	(9.078)	(22.152)
EVAR	0.172	-0.091	-0.010	-0.084	-2.234	4.690	3.716	8.116	-0.514
	(1.278)	(-0.196)	(-1.031)	(-0.747)	(-0.965)	(1.656)	(1.872)	(0.916)	(-0.369)
Adj R^2	0.044	-0.054	-0.020	-0.045	-0.026	0.012	0.104	0.006	-0.051

Table 7.9 notes. This table reports the results of bivariate cross section regressions for 20 countries. The dependent variable is the proportion of current income consumers (IV estimates). The regressors are an intercept (INT) and a variable (EVAR). The explanatory variable for each regression is given in the top row. The variables and favoured proxies are the private sector debt to GDP ratio ($CRED_i$), the saving rate (SY_i), the rate of unemployment (U_i) and its first difference (ΔU_i), the absolute deviation of income (AD_i), disposable income growth ($GRTH_i$), the proportion of the *total* population aged between 15 and 34 (YNG_i), population growth ($GPOP_i$) and real interest rates (r_i). T-ratios use White's (1980) heteroscedasticity consistent standard errors.

TABLE 7.10: Bivariate Cross-Country Models of Current Income Consumers, GMM Estimates, Equation (7.5.1)

	CRED	SY	U	ΔU	AD	GRTH	YNG	GPOP	r
INT	0.519	0.644	0.824	0.724	0.768	0.599	-1.204	0.576	0.713
	(6.523)	(7.656)	(10.287)	(13.062)	(11.752)	(4.819)	(-1.546)	(7.024)	(20.584)
EVAR	0.369	0.550	-0.023	-0.054	-2.843	4.350	6.390	18.006	-0.073
	(2.946)	(1.088)	(-1.227)	(-0.262)	(-1.077)	(1.044)	(2.363)	(1.807)	(-0.062)
Adj R^2	0.201	-0.021	0.049	-0.053	-0.029	-0.023	0.210	0.116	-0.056

Table 7.10 notes. This table reports the results of bivariate cross section regressions for 20 countries. The dependent variable is the proportion of current income consumers (GMM estimates). The variables and favoured proxies are the same as for Table 7.9.

Table 7.11 presents multivariate models for the proportion of current income consumers. The statistics are the same as those reported in Table 7.2 and the variables used are the same as those employed in Tables 7.9 and 7.10. Multivariate regressions are developed using the general to specific methodology. The preferred specifications are based around the three variables, GRTH, YNG and GPOP, identified as promising above plus SY. All of the four reported equations are

squares estimates support a statistically significant and positive relationship, we do not regard the GMM estimates as providing powerful evidence rejecting the modified REPIH/RELCH.

TABLE 7.11: Multivariate Cross-Country Models of Current Income Consumers, Equation (7.5.1)

	IV	GMM	GMM	GMM
	7.5.1a	**7.5.1b**	**7.5.1c**	**7.5.1d**
Intercept	-0.499	-1.220	-0.966	0.366
	(-0.891)	**(-1.404)**	**(-0.765)**	(2.009)
SY	-0.651			
	(-1.157)			
GRTH	6.305	2.771		7.155
	(1.842)	**(0.754)**		**(1.514)**
YNG	3.841	6.203	5.481	
	(1.997)	(2.297)	**(1.231)**	
GPOP			4.553	21.111
			(0.296)	(2.181)
Adj R^2	0.113	0.178	0.170	0.153
Pr[FR^2]	**[0.186]**	**[0.074]**	**[0.080]**	**[0.095]**
Pr[FSC1]	[0.496]	[0.788]	[0.636]	[0.914]
Pr[FFF1]	[0.359]	[0.561]	[0.615]	[0.751]
Pr[χ^2N2]	[0.768]	[0.981]	[0.842]	[0.779]
Pr[FH1]	[0.461]	[0.051]	[0.301]	[0.434]

Table 7.11 notes. This table reports the results of multivariate cross section regressions for 20 countries. The dependent variable is the proportion of current income consumers (both IV and GMM estimates). All statistics are the same as those reported in Table 7.2. All variables are as specified for Tables 7.9 and 7.10.

free from evident misspecification indicating valid inferences may be drawn.

Equation **7.5.1a** is the best multivariate model that could be developed using the IV estimates of π_i. The regression does not exhibit statistically significant explanatory power and all the variables, SY, GRTH and YNG, are individually insignificant at the five percent level. However, they are all *correctly* signed and GRTH and YNG are significant at the ten percent level.

The three remaining models are for the GMM estimates of π_i. These are exclusively based on combinations of the variables GRTH, YNG and GPOP, although all of these variables could not be included together. All three equations are statistically significant at the ten percent level, though not at the five percent level. Equation **7.5.1b** includes GRTH and YNG which are both correctly signed, however, only the latter is individually significant. Equation **7.5.1c** includes YNG and GPOP which are both correctly signed if individually insignificant. Equation **7.5.1d** includes GRTH and GPOP. Both feature the expected signs with the former being statistically insignificant and the latter statistically significant at the five percent level. These regressions provide evidence suggesting that some combination of GRTH, YNG and GPOP explain the variation in π_i. The best combination, according to individual t-ratios, appears to be GRTH and GPOP.

Both GMM and IV cross-section estimates provide some evidence supporting the view that π represents the proportion of liquidity constrained consumers. Following Jin (1994), we estimate the REPIH/RELCH model in a system, to exploit the larger sample provided by the pooled data, and explicitly allow the parameters to vary with the explanatory factors outlined in **(7.5.1)**.[27] That is, we estimate our modified REPIH/RELCH model using the equation:

(7.5.2)

$$\Delta \ln C_{it} = \mu_i + [b_0 + b_1 CRED_i + b_2 SY_i + b_3 U_i + b_4 \Delta U_i + b_5 AD_i + b_6 GRTH_i + b_7 YNG_i + b_8 GPOP_i + b_9 r_i] \Delta \ln Y_{it} + u_{it}.$$

Following Jin (1994) we estimate **(7.5.2)** using a system estimator. However, instead of simply using nonlinear three stage least squares (NL3SLS), as Jin (1994) does, we employ *iterative* NL3SLS.[28] Kennedy (1985) p. 140 cites Monte Carlo evidence which demonstrates the marked superiority of *iterative* 3SLS over 3SLS. 3SLS is the systems counterpart of the IV estimator (two stage least squares, 2SLS). In the current context this procedure involves estimating **(7.5.2)** by IV for the i countries – provided they are (over) identified. The *structural* equation coefficient estimates are retrieved from the reduced form parameters and are used to construct the

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[27] The greater information provided by the panel of data should enhance our inference.

[28] **(7.5.2)** could not be estimated using GMM, Eviews 2.0 suggested collinearity problems, although exactly the same equation could be estimated using 3SLS. We therefore provide 3SLS rather than GMM estimates.

variance/covariance matrix of the *structural* equation's residuals. Using this estimated variance/covariance matrix one estimates **(7.5.2)**, as a single equation, using generalised least squares (GLS), to produce the 3SLS estimates.[29] If the covariances are zero the 3SLS and IV estimators will produce identical estimates; however, if they are non-zero, 3SLS will be more efficient. The 3SLS estimates are consistent, even in the face of first order autocorrelation, provided we use appropriately dated instruments.[30] The instruments used are the second lags on consumption growth, income growth and the log-levels of consumption and income. The results of Chapter 6 suggest these to be the most relevant instruments for the majority of countries' REPIH/RELCH specifications.

Table 7.12 reports NL3SLS estimates of three versions of equation **(7.5.2)**.[31] The first, **7.5.2a** is the most general model including all the variables specified in **(7.5.2)**. All variables enter with theoretically justifiable signs and are statistically significant at the five percent level, except SY, AD and YNG, which are all highly insignificant. Removing these insignificant variables yields equation **7.5.2b**. All retained variables feature theoretically plausible signs and are statistically significant at the one percent level. It therefore represents our preferred model for drawing inference. This preferred equation suggests that CRED, U, ΔU, GRTH, GPOP and r, explain the cross-country variation in the proportion of current income consumers which is consistent with both liquidity constraints *and* precautionary saving explaining the variation in π_i.

Table 7.13 reports these systems' determinant of the residual variance/covariance matrix, $|\Omega|$,

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[29] *Iterative* 3SLS uses the 3SLS estimates to produce new estimates of the structural equations' error terms and, therefore, the system's error variance/covariance matrix. The latter is then used to produce new GLS parameter estimates. Calculation of the error variance/covariance matrix and GLS parameter estimates is repeated until these parameter estimates converge.

[30] Our results from Chapter 6 suggest that almost all equations are free of significant autocorrelation.

[31] To save space we do not report the estimated country specific intercepts.

TABLE 7.12: Iterative NL3SLS Estimates of the Cross-Country Variation in the Proportion of Current Income Consumers, Equation (7.5.2)

	7.5.2a	**7.5.2b**
Intercept (b_0)	0.836	0.484
	(1.438)	(3.899)
CRED (b_1)	-0.275	-0.366
	(-2.120)	(-3.548)
SY (b_2)	-0.211	
	(-0.289)	
U (b_3)	-0.073	-0.078
	(-4.324)	(-4.879)
ΔU (b_4)	0.584	0.648
	(2.905)	(3.756)
AD (b_5)	3.325	
	(0.885)	
GRTH (b_6)	12.561	12.682
	(3.740)	(4.386)
YNG (b_7)	-1.489	
	(-0.625)	
GPOP (b_8)	25.993	26.874
	(2.664)	(3.645)
r (b_9)	6.857	5.865
	(3.071)	(3.708)

Table 7.12 notes. NL3SLS estimates of cross country variation in the proportion of current income consumers, equation (7.5.2). T-ratios are reported below the estimated parameters. Bold emphasis indicates an insignificant parameter. Different intercepts are allowed for each country (fixed effects) but are not reported to save space. The same instrument set is used for each country being: $\Delta \ln C_{t-2}$, $\Delta \ln Y_{t-2}$, $\ln C_{t-2}$, $\ln Y_{t-2}$ and an intercept.

TABLE 7.13: (Implied) Iterative (NL)3SLS Estimates of the Proportion of Current Income Consumers, Equations (6.2.34) and (7.5.2)

	6.2.34a			7.5.2a		7.5.2b	
	π_i (PI1)	t	Adj R^2	π_i (PI2)	Adj R^2	π_i (PI3)	Adj R^2
AUL	0.911	(4.874)	**-0.165**	0.805	**-0.241**	0.841	**-0.202**
AUT	0.881	(5.333)	0.363	0.732	0.243	0.746	0.320
BEL	0.473	(4.375)	0.554	0.530	0.410	0.525	0.474
CAN	0.669	(7.388)	0.659	0.682	0.534	0.708	0.589
DEN	0.708	(4.383)	0.148	0.629	**-0.068**	0.561	0.108
FIN	0.863	(8.126)	0.625	0.918	0.469	0.889	0.536
FRA	0.693	(10.095)	0.637	0.654	0.514	0.655	0.568
GER	0.770	(9.990)	0.810	0.833	0.725	0.787	0.766
GRE	0.680	(8.395)	0.670	0.685	0.538	0.701	0.571
ICE	0.701	(6.639)	0.759	0.730	0.669	0.737	0.705
IRE	0.693	**(1.682)**	0.402	0.462	0.179	0.411	0.246
ITA	0.583	(5.667)	0.528	0.524	0.333	0.544	0.414
JAP	0.869	(11.963)	0.779	0.889	0.690	0.906	0.719
NET	0.740	(8.009)	0.653	0.656	0.526	0.682	0.580
NOR	0.832	(2.501)	0.190	0.615	**-0.092**	0.625	0.030
SPA	0.727	(8.132)	0.659	0.592	0.468	0.577	0.517
SWE	0.837	(7.075)	0.223	0.675	0.019	0.647	0.135
SWZ	0.588	(10.147)	0.690	0.625	0.575	0.620	0.622
UK	0.511	(5.508)	0.559	0.440	0.330	0.397	0.362
USA	0.397	(2.287)	0.466	0.459	0.335	0.461	0.409
$\|\Omega\|$	6.15×10^{-78}			3.09×10^{-78}		3.63×10^{-78}	
SBIC	-177.411			-178.193		-178.060	

Table 7.13 notes. The table reports the proportion of credit constrained consumers, π_i, with associated t-ratios, t, and adjusted R^2, Adj R^2, for equation **(6.2.34)**, without MA error, and the adj R^2 and value of π_i implied by equations **7.5.2a** and **7.5.2b**, reported in Table 7.12. $|\Omega|$ denotes the determinant of the estimated residual variance/covariance matrix for each system, where SBIC=$\ln|\Omega|$ + p[ln(NT)/NT], with N=20 countries, T=35 periods and p is the number of estimated parameters.

with corresponding SBIC,[32] the adjusted R^2 of each individual equation in each system and the proportion of current income consumers *implied* by each system's parameter estimates.[33] We also report the iterative 3SLS estimates of equation **(6.2.34)**, which does not allow π_i to vary with explanatory factors (denoted **6.2.34a**), for comparative purposes.[34] All three systems feature plausible estimates of π_i, in the sense that they fall between zero and one for all countries, although the estimate for Ireland is statistically insignificant for system **6.2.34a**. Both the systems where π_i varies with explanatory factors, **7.5.2a** and **7.5.2b**, feature notably superior SBICs to the system where π_i is not related to variables, **6.2.34a**. Thus, **7.5.2a** and **7.5.2b** are preferred to **6.2.34a**. System **7.5.2a** features a slightly better SBIC compared to **7.5.2b**, however, three of **7.5.2a**'s individual equations' adjusted R^2s are negative while only one is negative for **7.5.2b** (one country's adjusted R^2 is also negative in system **6.2.34a**). Further, and as suggested above, three of the factors explaining the variation in π_i are statistically insignificant in **7.5.2a** while all such factors are significant at the one percent level for system **7.5.2b**. Overall, we favour system **7.5.2b** for inference, with its well determined parameters and fewer negative adjusted R^2s, despite a *slightly* inferior SBIC relative to **7.5.2a**.

As can be seen from Figure 7.6, which visually plots the actual and implied estimates of π_i for these three systems (where PI1, PI2 and PI3 denote the values of π_i for systems **6.2.34a**, **7.5.2a** and **7.5.2b**, respectively), the estimates for systems **7.5.2a** and **7.5.2b** are extremely similar. This suggests that the three explanatory factors excluded from **7.5.2a** to obtain **7.5.2b** have virtually no impact on the implied estimates of π_i. However, the estimates of π_i for **6.2.34a** are clearly different from those of the other two systems, indicating the impact of allowing π_i to vary with explanatory factors.

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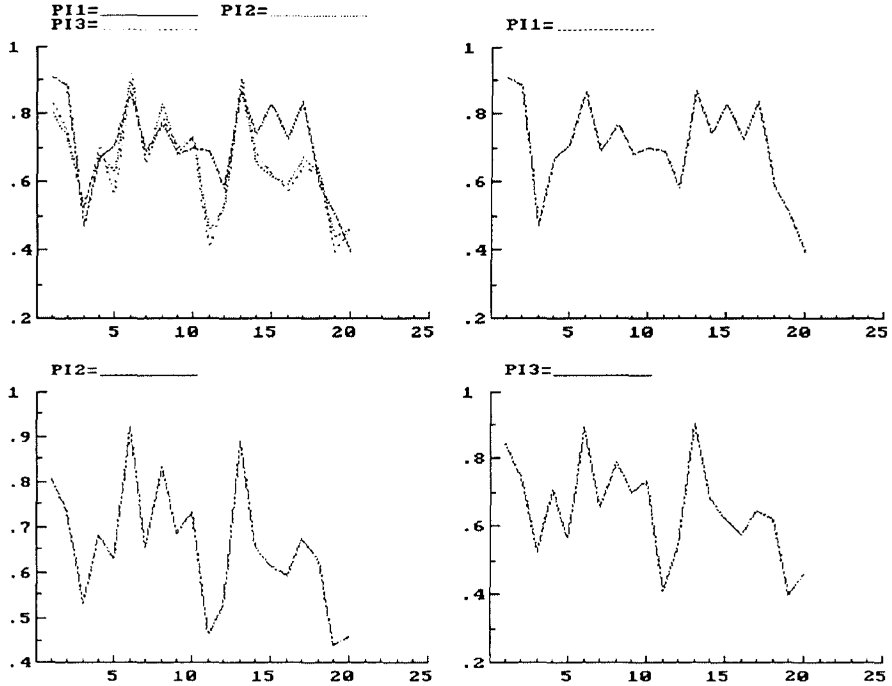
[32] $SBIC = \ln|\Omega| + p[(\ln NT)/NT]$, where N=20 (countries), T=35 (time periods) and p is the number of parameters in each system (p=40 for **6.2.34a**, p=30 for **7.5.2a** and p=27 for **7.5.2b**).

[33] We calculate the implied proportions of current income consumers by substituting the cross-country estimates from Table 7.12 into the formula $\pi_i = [b_0 + b_1CRED_i + b_2SY_i + b_3U_i + b_4\Delta U_i + b_5AD_i + b_6GRTH_i + b_7YNG_i + b_8GPOP_i + b_9r_i]$.

[34] This is equivalent to estimating **(7.5.2)** imposing the restriction $b_i = 0$, for i=1,...,9.

FIGURE 7.6: Iterative NL3SLS Etimates of the Proportion of Current Income Consumers (PI)



The superiority of **7.5.2b** suggests that π_i varies across countries with credit, the level and change in unemployment, income and population growth and the rate of interest. Exploitation of the information incorporated in the full panel of the data facilitates clarification of the multivariate time series results reported in Table 7.11. In particular, we are able to clarify that factors beyond income and population growth influence the proportion of current income consumers and that the proprtion of the population who are young and the saving rate do not explain the cross-country variation in π_i. Further, our preferred system for drawing inference, **7.5.2b**, suggests that up to five proxies for liquidity constraints, CRED, ΔU, GRTH, GPOP and r, explain the cross-country variation in the proportion of current income consumers. Of course, ΔU is also consistent with precautionary saving explaining the variation in π_i while the sign of U is *only* consistent with a precautionary saving interpretation. Thus, our evidence suggests that the proportion of current income consumers and, therefore, the excess sensitivity of consumption to predictable changes in income growth, systematically vary with the intensity of liquidity constraints and the degree of income uncertainty.[35]

7.6 Conclusions

In this Chapter we have attempted to explain the cross-country variation of long and short run consumption elasticities with respect to income and inflation, the speed of adjustment to disequilibrium (adjustment coefficient) and the proportion of current income consumers. To this end, bivariate and multivariate cross section regressions, relating each of these parameters to potential explanatory factors, have been employed. We are not aware of previous analyses which seek to explain cross-country variations in these parameters, except for the proportion of current income consumers, and this represents the main innovation of this Chapter. Further, we extend the variables considered by Jin (1994) for explaining cross-country variation in the proportion of current income consumers to include a direct measure of private sector credit and to consider a precautionary saving interpretation.

The long run income elasticities are found to be negatively related to income growth. This is consistent with Modigliani's (1986 and 1990) LCH theory of the APC and reinforces the evidence provided by Jappelli and Pagano (1994). We find the fiscal surplus/deficit to GDP ratio to be positively associated with the income elasticity suggesting at least partial Ricardian equivalence, which is consistent with recent studies (see Jappelli and Pagano 1994 and Pesaran, Haque and Sharma 1999). We provide some evidence that holdings of private sector credit positively influence the income elasticity which is consistent with the findings of Jappelli and Pagano (1994).

Our results indicate that the log of per-capita income is negatively correlated with the income elasticity. The use of the natural logarithm means that the income elasticity does not necessarily

[35] Data constraints have prevented us from considering consumer confidence variables to approximate income uncertainty effects, as employed by Acemoglu and Scott (1994), Carroll, Fuhrer and Wilcox (1994) and Fan and Wong (1998). Use of such data, when it becomes available for a broad range of countries, should enable one to further clarify the relationship between precautionary saving and excess sensitivity.

become negative as the *level* of income rises, which removes the objections that can be made against a simple linear relationship. This provides support for a negative relationship between per-capita income and the APC often attributed to Keynes's (1936) AIH and contradicts an implication of Modigliani's (1986) LCH. This finding represents a particular innovation of the current study.

There is also some evidence indicating that increased income inequality reduces the income elasticity. This is consistent with the RIH and the simulation results of Borooah and Sharpe (1986).

We find that the credit available to the private sector and the dependency ratio have negative impacts upon the long run inflation elasticity while income uncertainty exhibits a positive association. These results are consistent with inflation approximating wealth effects (through a negative correlation) in the long run consumption function and supports the evidence produced by, for example, Lattimore (1994). That is, the results imply that increased credit raises the spendability of assets, increased uncertainty raises the accumulation of wealth while an increased dependency ratio raises spending out of wealth to support the needs of children.

We are unable to uncover any systematic explanation of the variation in the short run response of consumption to income or inflation, except a possible, if weak and unconvincing, interest rate or capital gains effect for the income elasticity. This lack of empirical evidence is consistent with the general shortage of theory on the short run.

We find no evidence to suggest that the adjustment coefficient systematically varies with the degree of credit constraints. This may also reflect a lack of theory regarding the variation in this parameter.

We find evidence suggesting that both liquidity constraints and income uncertainty explain the cross-country variation in the proportion of current income consumers. Liquidity constraints unambiguously influence the proportion of current income consumers through private sector credit, income and population growth and the rate of interest. Liquidity constraints and/or precautionary saving may also influence this proportion through the change in the rate of unemployment, while the negative association of the rate of unemployment with the proportion of current income consumers is only consistent with a precautionary saving interpretation. The finding of a role for precautionary saving, in addition to liquidity constraints, in explaining the *cross-country* variation in the proportion of current income consumers is also an innovation of this study.

7.A Data Appendix

The variables used to explain the cross-country variations in the estimated parameters of this Chapter are the averages, over the period 1960-1994 (unless otherwise stated), of the variables listed below.

CRED is proxied by two different variables: the money to GDP ratio (**M/G**) and the private sector domestic credit to GDP ratio (**PSDC/G**). Where (broad) money, **M**, is defined as money plus quasi-money, reported as lines 34 and 35, respectively, in the data source International Monetary Fund International Financial Statistics (IMFIFS). Sources for, and construction of, GDP, **G**, is detailed in Chapter 3. Private sector domestic credit, **PSDC**, is obtained from line 32d of IMFIFS.

DEF = **D/G**. Where the Central Government fiscal surplus/deficit, **D**, is obtained from line 80 of IMFIFS. Data is available over the period 1960-1994 for all countries except Iceland (1960-64 and 1972-94), Japan (1960-93) and Spain (1962-94). The periods over which the data is available is given brackets. When data over the full sample (1960-1994) is not available, we take the mean of the available data, for the cross-country comparisons. Sources for, and construction of, GDP, **G**, is detailed in Chapter 3.

f(INC) This is a function of the level of per capita income in a common currency. Where the measure of income, **INC**, is real (1990) per-capita GDP in thousands of Geary-Khamis US dollars. This is available for all countries, except Iceland, from Maddison 1995. For Iceland we use real GDP in Kronur (from UN/OECD National Accounts) and convert it into dollars by dividing by the Kronur/dollar exchange rate (line rf from IMFIFS). We then multiply by the ratio of Iceland's consumer price index to that of the USA. This is then multiplied by the ratio of Norway's average GDP level in Geary-Khamis US dollars to its GDP level converted into US dollars: (12268.5/14585.2). This adjusts, to some degree, the difference in the two calculations for Iceland (being the adjustment for purchasing power parity). Our functions are based upon the following four transformations. The first is simply the untransformed variable, **INC**. The second is its natural logarithm, ln **INC**. The third is its square root, $\sqrt{\textbf{INC}}$. The fourth is the squared value of the log, $(\ln\textbf{INC})^2$.

GAIN ~ $\Delta \ln P_t$, where P_t is the price level (1990~100). See Chapter 3 for sources and construction of prices.

GRTH $= \ln Y_t - \ln Y_{t-1}$, where Y_t is per-capita real private disposable income at 1990 prices. We also try per-capita GDP growth, however, this is not the favoured measure. Sources and construction of income and GDP are outlined in Chapter 3.

INEQ Gini coefficient of income inequality. This is reported on p. 21 of Atkinson (1996). It is only available for thirteen countries: Belgium, Finland, France, Germany, Ireland, Italy, the Netherlands, Norway, Spain, Sweden, Switzerland, the UK and the USA.

GPOP is population growth. See Chapter 3 for sources and construction of population.

r The natural logarithm of one plus the short term real interest rate is exactly defined as $\ln[1+(I_t/100)]-\Delta \ln P_t$, where P_t is the price level (1990~100) and I_t is the nominal rate of interest. Sources and construction of prices and interest rates are outlined in Chapter 3.

RTRE Retirement Length. The average life expectancy at birth (years) for both sexes, **LE**, minus 65. It is assumed that 65 is the age of retirement for both sexes in all countries. Where **LE** is reported in Table A.15 of UN World Population Prospectus the 1992 Revision (1993). Averages for the periods 1950-55, 1955-60,..., 1990-95 are reported for all 20 OECD countries. We try two measures, the first averages the data from 1960-1994. The second averages **LE** over the period 1950-80 to recognise that only the economically active make expenditure decisions.

SY $= (Y_t - C_t)/Y_t$. Where C_t and Y_t are real per-capita disposable income and total consumption, respectively. See Chapter 3 for details of sources and construction of consumption and income.

U is the standardised rate of unemployment. See Chapter 3 for sources and details of construction of this variable.

ΔU $= U_t - U_{t-1}$.

UNCT We use three measures for **UNCT**: **U**, **ΔU** and **AD**. The first two are outlined above. The third, **AD**, the absolute deviation of income growth is defined as: $|\Delta \ln Y - \text{"trend"}|$, where "trend" is, following Muellbauer (1994) and Muellbauer and Lattimore (1995), the average trend growth over the past 5 years. That is, an MA(5) of income growth.

Age Structure variables are based upon the age distribution of the population, **AGE**, which is broken down into those (of both sexes) falling in the inclusive age groupings: 0-4, 5-9, 10-14, 15-19, 20-24, 25-29, 30-34, 35-39, 40-44, 45-49, 50-54, 55-59, 60-64, 65-69, 70-74, 75-79 and 80+; for the years 1950, 1955, 1960, 1965, 1970, 1975, 1980, 1985, 1990 and 1995. Where the 1995 observation is a "medium variant" projection. We take the averages of appropriate combinations of these age groupings over the period 1960-1995 for our cross-country analysis. The data source is *The Sex and Age Distribution of the World Populations the 1994 Revision* UN 1994. The variables obtained from this source are listed below.

DEP The dependency ratio. The proportion of the *total* population aged between 0 and 15 years. We also consider the proportion of the population aged between 0 and 19. The former measure is favoured.

YNG The proportion of the population who are young. We try three measures for this: the proportions of the *total* population aged between 15 and 24 years, between 15 and 29 years and between 15 and 34 years. The later is the favoured measure.

RSAV The proportion of the population who are saving for retirement. We consider three measures for this: the proportions of the *total* population aged between 40 and 64 years, between 45 and 64 years and between 50 and 64 years. We favour the latter measure.

65+ The proportion of the *total* population aged 65 and over. The retired.

SUPT The support ratio. The number of the population of working age (15-64) divided by the number of retirement age (65 and over).

CHAPTER 8

CONCLUSIONS

Research into consumer behaviour over the past twenty years has been dominated by two modelling methodologies: error correction mechanisms (ECMs) and (modified) rational expectations permanent income hypothesis/life cycle hypothesis (REPIH/RELCH) models. Empirical investigations of these two methods have primarily focussed upon the UK and the USA although a small number of recent analyses have examined a broader range of countries. Those analyses which consider a broad range of economies always focus on one model only and any cross-country investigation conducted is generally limited. Furthermore, the substantive findings of these recent international analyses have been based upon income data incorporating the public sector.

The present thesis has sought to empirically investigate the ECM and REPIH/RELCH methodologies for twenty OECD economies over the period 1960-1994 using income data exclusively based upon the private sector. This provides the longest time-series of data exclusively employing the more appropriate private disposable income measure for all twenty OECD countries compared to any previous study. In Chapter three we argued that transformations of variables based upon GDP and private disposable income featured, at times, very low correlations, and could potentially yield quite disparate inference. Since private disposable income is closer to the concepts of Hicksian and labour income compared to measures incorporating public sector income, such as GDP, we argue that inference from models using private disposable income will be superior.

We also employ more flexible specifications relative to any previous international comparative study. Time-series ECMs, based upon the log of consumption, the log of income and inflation, are estimated - in Chapter three these three variables were generally found to be I(1) across countries therefore satisfying the necessary condition for forming a cointegrating relation. These are the variables employed by Davidson *et al* (1978) - DHSY - and this specification may be interpreted as an approximation of Ando and Modigliani's (1963) life cycle hypothesis (LCH) formulation with inflation capturing asset effects. Our specifications allow short run dynamics to be heterogeneous across countries, intercepts to be included or excluded from the long run consumption functions and inflation to be incorporated or omitted from both the long and short run component of the model. We also consider dynamic models facilitating two forms of asymmetric adjustment towards long run equilibrium: the partitioned and (reduced) cubic forms.

In Chapter four we present evidence for the existence of a cointegrating vector for all twenty countries. All countries' cointegrating vectors represent reasonable approximations of long run consumption functions according to statistical and theoretical criteria, except for the Swedish and Swiss long run models which are not convincing as equilibrium consumption functions (with incorrectly signed adjustment coefficients in the consumption growth equation).

The long run income elasticities estimated in Chapter four vary substantially across countries. There is evidence of a below unit income elasticity for six countries, a unit income elasticity for nine countries and an above unit income elasticity for five countries. The evidence for a below unit income elasticity may be underestimated due to omitted variable bias raising some of the countries' elasticities and the poor determination of some countries' income elasticities. Thus, we argue that one should not automatically assume that consumption is homogeneous of degree one in income for any particular OECD country. This is consistent with recent evidence which suggests that OECD countries' consumption-income ratios are nonstationary.

In Chapter seven we provide evidence indicating that the cross-country variation in the estimated long run income elasticities are negatively related to income growth and the log-level of income (which allows the income elasticity to decline at a decreasing rate as the level of income rises). The former is consistent with Modigliani's (1986 and 1990) LCH although the latter is not, being more consistent with a postulate often attributed to Keynes (1936). There is an evident positive relationship between the fiscal surplus/deficit and the long run income elasticity which is consistent with *at least* incomplete Ricardian equivalence. Thus, fiscal policy may have little, if any, influence on aggregate demand because the impact of an increased fiscal deficit on national income will be (partially) offsett by a reduction in the propensity to consume out of the private sector's *post-tax* income. This is consistent with the typical finding of partial, but not complete, Ricardian equivalence. We also find that increased income inequality reduces the income elasticity. This indicates that policies aimed at reducing (raising) income inequality can boost (lower) aggregate consumer demand; in particular, taxation policies. We also find evidence that the availability of credit influences the income elasticity. Thus, economies where institutions provide greater access to credit feature higher levels of aggregate consumption at given levels of income. This implies that in countries where consumers are subject to binding liquidity constraints, policies which increase credit availability can boost consumption. For example, the removal of ceilings and guidelines on bank lending (especially in the housing market) and the removal of interest rate and exchange controls - see Miles (1992) and Berg (1994). Indeed, European Monetary Union (EMU) may influence the EMU countries' consumer demand through the greater integration of their financial markets. Tax cutting policies could also boost consumption when households face binding liquidity constraints - this would be consistent with incomplete Ricardian equivalence. We find that demographic factors, inflation (capital gains), interest rates and income uncertainty do not influence the long run income elasticity.

The long run inflation elasticities estimated in Chapter four are statistically significant and negative in only seven of the twenty countries' long run consumption functions and inflation is excluded from the Irish equilibrium equation. This suggests that inflation may not be a fundamental explanatory factor of long run consumer behaviour for many OECD countries. Nevertheless, in Chapter seven we present evidence which is consistent with inflation approximating wealth effects (through a negative correlation) in the long run consumption function. That is, we may interpret our cross-country analysis of the long run inflation elasticities as indicating that increased credit raises the spendability of assets, increased income uncertainty raises the accumulation of wealth while an increased dependency ratio raises spending out of wealth. Thus, policies which raise credit availability (financial deregulation policies are discussed above) and reduce income uncertainty may promote aggregate consumer demand through wealth effects. Policies regarding the latter would include legislation aimed at reducing Trade Union's power when strikes are a cause of instability, as in the UK during the 1970s, reform of social security provision (for example, a move away from state provision of pensions), and promoting consumer confidence via stable economic growth (for example, giving control over monetary policy to independent central banks). Malley and Moutos (1996) argue that the success of monetary union might be hindered without first removing differences in EMU countries' social security provision, where such differences will cause differences in the degree of these economies' precautionary saving.

The cointegration evidence from Chapter four is reflected in Chapter five where we are able to develop error correction models which are consistent with consumption being continuously forced towards its long run equilibrium for all twenty countries except Sweden and Switzerland. The models for Sweden and Switzerland are interpreted as short run consumption functions. These models feature good explanatory power and provide valid inference according to a range of misspecification tests. Wald tests provide evidence against asymmetric adjustment towards equilibrium for partitioned and full cubic nonlinear error correction specifications, however, this is probably due to this test's low power. In contrast, this test does provide evidence suggesting statistically significant nonlinear adjustment using the reduced cubic specification for thirteen countries, however, because linear error correction terms can be excluded this does not necessarily reflect a preference for nonlinear adjustment over linear adjustment. To determine the favoured form of adjustment we compare the fit of the various specifications. Linear symmetric adjustment is preferred for six countries, the partitioned form for one country, the full cubic model for three countries and the reduced cubic specification for eight countries. This provides evidence favouring nonlinear/asymmetric adjustment towards equilibrium for just over half of the twenty OECD countries considered here. Further, the (reduced) cubic nonlinear form of adjustment is virtually always preferred to the partitioned form, suggesting that the speed of adjustment is related to the degree of disequilibrium. This might be expected of models explaining total consumer expenditures which embody a durable component because there may be a threshold which determines when tolerable disequilibrium durable expenditures become intolerable. This indicates a need to consider the role of adjustment costs, particularly associated with durability, in the theoretical and empirical specification of models of consumer behaviour. Indeed, the evident presence of asymmetries suggests that this may represent an omitted parameterisation from standard consumption functions.

Although the adjustment coefficient and short run income and inflation elasticities, estimated in Chapter five, vary considerably from country to country we are unable to explain the cross-country variation in these parameters (see Chapter seven). We note that only seven countries' ECMs incorporate short run inflation effects, providing further evidence that inflation may not be a fundamental explanatory factor of many OECD economies' consumer behaviour. This contrasts with short run income effects which appear in all twenty countries' ECMs.

In Chapter six we develop a REPIH/RELCH model in logarithmic form which explicitly accounts for current income consumers, a component of durable expenditures and intertemporal substitution. No previous study develops a modified REPIH/RELCH specification to allow for all three of these features simultaneously. Using the ARIMA method we find that only ten countries' models are consistent with durability being the sole cause of the rejection of the REPIH/RELCH. We find that intertemporal substitution, in common with the majority of previous studies, does not provide a satisfactory explanation for the failure of the REPIH/RELCH, either on its own or when added to a model incorporating predictable income growth. In contrast, the REPIH/RELCH model modified to allow for a proportion of current income consumers does provide a satisfactory explanation of the data for all twenty OECD countries. The instrumented income growth term's coefficient is statistically significant and between zero and one for all countries. The estimates provided using GMM and IV-MA estimation methods do not exhibit statistically significant differences. The average proportion of current income consumers is 65%-73%, depending upon estimation method, which is larger than previous studies' estimates and probably reflects our use of total consumption as the dependent variable, allowing current income consumers to purchase durables, and that the sample covers the economic downturn experienced by many economies in the 1990s. This high proportion of current income consumers would also explain our finding against consumption smoothing for many OECD countries (see Chapter three) and is consistent with the notions of bounded rationality, liquidity constraints and precautionary saving. In addition to current income consumers we find evidence of a moving average error process, consistent with the importance of durability, for Italy and Switzerland. Thus, it is the excess sensitivity of consumption to income that is the primary explanation for the failure of the REPIH/RELCH for our twenty OECD countries. There is evidence that the proportion of current income consumers varies through time for three to six economies - typically the UK and Nordic countries. Future work may focus on characterising this temporal variation and identifying the factors which explain it in these countries.

The factors which cause a large proportion of current income consumers to exist across countries will determine the policy implications of this phenomenon. In Chapter seven we present evidence suggesting that both liquidity constraints and income uncertainty explain the cross-country variation in the proportion of current income consumers, with liquidity constraints appearing to be the more dominant factor. Liquidity constraints unambiguously influence the proportion of current income consumers through private sector credit, income and population growth and the rate of interest though not through the saving rate and the proportion of the population who are young. Liquidity constraints and/or income uncertainty may also influence this proportion through the change in the rate of unemployment, while the negative association of the rate of unemployment with the proportion of current income consumers is only consistent with a precautionary saving interpretation - there is no uncertainty effect through the absolute deviation in income. The finding of a role for precautionary saving, in addition to liquidity constraints, in explaining the cross-country variation in the proportion of current income consumers is an innovation of this study. These results reinforce the conclusion drawn above, that policies which successfully affect the supply of credit to consumers and future income uncertainty can influence aggregate consumer demand.

Future empirical investigation into international consumer behaviour would benefit from the use of direct measures of wealth, rather than relying upon the inflation proxy, especially in a structural model such as an ECM. Indeed, this would enable clarification of the extent of non-wealth inflation effects. Obtaining international inference using the most theoretically appropriate labour income measure, especially for REPIH/RELCH models, would also be beneficial as would considering whether assumptions of Hall's model, other than those considered here, can help explain the failure of REPIH/RELCH in OECD countries. The separate modelling of non-durable and durable expenditures for a broad range of economies would also be of interest. Such work should become feasible as international data availability increases.

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