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Persistence in US Interest Rates: Is it Stable Over Time?

Guglielmo Maria Caporale, Luis A. Gil-Alana

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Guglielmo Maria Caporale
Brunel University, London

and

Luis A. Gil-Alana
University of Navarra

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Abstract

This paper analyses persistence in US interest rates. It focuses on the Federal Funds effective rate, whose degree of persistence is modelled using fractional integration, monthly from July 1954 through March 2008. The full-sample estimates of the fractional differencing parameter appear to be very sensitive to the choice of the I(0) error term; specifically, the order of integration is strictly above 1 if the errors are uncorrelated, whilst it is strictly below 1 with autocorrelated disturbances. It is also found that the differencing parameter is not stable over the sample period examined, the degree of persistence of the series decreasing from the beginning of the 80s.

JEL Classification: C22, G1
Keywords: Fractional Integration; Interest Rates; Persistence

Corresponding author: Professor Guglielmo Maria Caporale, Centre for Empirical Finance, Brunel University, Uxbridge, Mddlesex UB8 3PH, UK. Tel.: +44 (0)1895 266713. Fax: +44 (0)1895 269770. *Email:* Guglielmo-Maria.Caporale@brunel.ac.uk

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1. Introduction

The statistical properties of interest rates have been extensively investigated in the literature. Earlier studies typically focused on whether interest rates can be characterised as an I(0) or I(1) series. For instance, Cox, Ingersoll and Ross (1985) concluded that the short-term nominal interest rate is a stationary and mean-reverting I(0) process, whereas authors such as Campbell and Shiller (1987) assumed a unit root. A drawback of the I(0) models is that they imply long-rates which are not volatile enough (Shiller, 1979) whereas a problem with the I(1) models is that they imply that the term premium necessarily increases with bond maturities (Campbell, Law and MacKinlay, 1997).[1]

Other studies analysed whether or not real rates are stationary, since a unit root in ex-ante real rates is inconsistent not only with the Fisher hypothesis but also with the consumption-based capital asset pricing model (CCAPM) of Lucas (1978) (see Rose, 1988). Various papers found a unit root in the real interest rate (see, e.g., Goodwin and Grennes, 1994; Phylaktis, 1999; Rapach and Wohar, 2004). However, the low power and limitations of traditional unit root tests are now well known. More recent studies have tried to deal with them by using long-horizon data (see, e.g., Sekioua and Zakane, 2007), or adopting a fractional integration approach. The latter offers much more flexibility compared to the usual I(0)/I(1) dichotomy, as it encompasses the intermediate case of the degree of integration being between 0 and 1, as well as above 1. This is particularly useful for series which, although mean-reverting, might exhibit long memory and therefore be characterised by a high degree of persistence. For example, Shea (1991) investigated the consequences of long memory in interest rates for tests of the expectations hypothesis of the term structure. He found that allowing for the possibility of long memory significantly improves the performance of the model, even

[1] Recently, Gil-Alana and Moreno (2008) have therefore proposed a fractional model for the short-term interest rate and the term premium. Also, Kozicki and Tinsley (2001) propose a model with shifting endpoints for short-term interest rates, while Ang and Beckaert (2002) develop a regime-switching model.

though the expectations hypothesis cannot be fully resurrected. In a related study, Backus and Zin (1993) observed that the volatility of bond yields does not decline exponentially when the maturity of the bond increases; in fact, they noticed that the decline is hyperbolic, which is consistent with the fractionally integrated specification. Lai (1997) and Phillips (1998) provided evidence based on semiparametric methods that ex-ante and ex-post US real interest rates are fractionally integrated. Tsay (2000) employed an Autoregressive Fractionally Integrated Moving Average (ARFIMA) model to show that the US real interest rate can be described as an I(d) process. Further evidence can be found in Barkoulas and Baum (1997), Tkacz (2001), Meade and Maier (2003), Sun and Phillips (2004), Gil-Alana (2004a, b), and Karanasos et al (2006). Couchman, Gounder and Su (2006) estimated ARFIMA models for ex-post and ex-ante interest rates in sixteen countries. Their results suggest that, for the majority of countries, the fractional differencing parameter lies between 0 and 1, and is considerably smaller for the ex-post than for the ex-ante real rates.

Fractional cointegration tests have also been applied in recent studies. Lardic and Mignon (2003) tested for fractional cointegration between nominal interest rates and inflation under the assumption that both individual series were I(1). They tested this hypothesis with standard unit root procedures (Dickey and Fuller, 1979; Phillips and Perron, 1988; Kwiatkowski et al., 1992). However, these methods have extremely low power if the alternatives are of a fractional form (Diebold and Rudebusch, 1991; Hassler and Wolters, 1994; Lee and Schmidt, 1996). Barkoulas and Baum (1997) also used fractional integration to model nominal interest rates and found evidence of long memory in the differenced series. Mean reversion in nominal rates is reported for Asian and emerging countries respectively in Gil-Alana (2004a) and Candelon and Gil-Alana (2006).

The present paper analyses the behaviour of the US Federal Funds effective rate and, similarly to various studies referenced above, uses a fractional integration framework to estimate its degree of persistence. However, it also makes the additional contribution of examining thoroughly, by means of recursive and rolling techniques, its stability over time, an issue which has not been addressed in the previous literature. The layout of the paper is the following. The econometric model and estimates are reported in Section 2, whilst the final Section 3 summarises the main findings and offers some concluding remarks.

2. Econometric model and estimates

A time series process $\{y_t, t = 0, \pm 1, \ldots\}$ is said to be integrated of order d, and denoted by $y_t \approx I(d)$, if it requires d-differences to render the series stationary I(0), that is,

$$\begin{aligned} (1 - L)^d \, y_t &= u_t, \quad t = 1, 2, ..., \\ y_t &= 0, \quad t \leq 0, \end{aligned} \tag{1}$$

where, u_t is an I(0) process, defined as a covariance stationary process with spectral density function that is positive and finite, and L is the backward shift operator.[2] In the event that d is not an integer, the series y_t requires fractional differencing in order to obtain a possibly stationary ARMA series. ARIMA(p,d,q) models in which d is a positive integer are then special cases of the general process in (1). If $d > 0$, y_t is said to have long memory because of the strong association between observations distant in time, and the higher the value of d is, the higher is the level of dependence in the time series behaviour.[3]

[2] The condition $y_t = 0$, $t \leq 0$ is required for the Type II definition of fractional integration. For an alternative definition (Type I), see Marinucci and Robinson (1999).
[3] See Robinson (1994a, 2003), Beran (1994), Baillie (1996), Doukhan et al. (2003) and Gil-Alana and Hualde (2008) for surveys of the literature on fractional integration.

The time series data analysed in this paper is the Federal Funds effective rate, monthly, (with overnight maturity), from July 1954 to March 2008, obtained from the Board of Governors of the Federal Reserve System (http://www.federalreserve.gov).

[INSERT FIGURE 1 ABOUT HERE]

Figure 1 shows a plot of this time series. Visual inspection suggests that it exhibits an upward trend until the beginning of the 80s, when the trend is reverted and becomes negative.

As a first step we compute an estimate of d for the whole sample period. For this purpose, we use a Whittle estimator in the frequency domain along with a testing procedure developed by Robinson (1994b). This method is parametric, and does not require preliminary differencing; it allows us to test any real value d, thus encompassing stationary and nonstationary hypotheses. We consider the following formulation

$$y_t = \alpha + \beta t + x_t, \qquad (2)$$

$$(1 - L)^d x_t = u_t, \qquad (3)$$

testing the null hypothesis,

$$H_o : d = d_o, \qquad (4)$$

in (2) and (3) for any real value d_o. Initially, we assume that α = β = 0 a priori in (2), that is, we suppose that there are no deterministic terms in the regression model (2), though we also consider the cases of an intercept (α = 0 and β unknown), and an intercept with a linear time trend (α and β unknown). Given the parametric nature of the tests, we must specify the functional form of the disturbance term, u_t. Here we consider the two cases of white noise and autocorrelated errors. In the latter case we first assume an AR(1) process for u_t. Then, the model of Bloomfield (1973) is also considered. This is a non-parametric approach to modelling the I(0) error term that produces autocorrelations decaying exponentially as in the AR(MA) case.[4]

Robinson's (1994b) test statistic is given by:

$$\hat{r} \;=\; \frac{T^{1/2}}{\hat{\sigma}^2}\,\hat{A}^{-1/2}\,\hat{a}\,, \tag{5}$$

where T is the sample size and

$$\hat{a} \;=\; \frac{-2\pi}{T}\sum_{j=1}^{T-1}\psi(\lambda_j)\,g(\lambda_j;\hat{\tau})^{-1}\,I(\lambda_j); \qquad \hat{\sigma}^2 \;=\; \sigma^2(\hat{\tau}) \;=\; \frac{2\pi}{T}\sum_{j=1}^{T-1}g(\lambda_j;\hat{\tau})^{-1}\,I(\lambda_j);$$

$$\hat{A} \;=\; \frac{2}{T}\left(\sum_{j=1}^{T-1}\psi(\lambda_j)^2 \;-\; \sum_{j=1}^{T-1}\psi(\lambda_j)\,\hat{\varepsilon}(\lambda_j)'\times\left(\sum_{j=1}^{T-1}\hat{\varepsilon}(\lambda_j)\,\hat{\varepsilon}(\lambda_j)'\right)^{-1}\times\sum_{j=1}^{T-1}\hat{\varepsilon}(\lambda_j)\psi(\lambda_j)\right)$$

$$\psi(\lambda_j) = \log\left|2\sin\frac{\lambda_j}{2}\right|; \quad \hat{\varepsilon}(\lambda_j) \;=\; \frac{\partial}{\partial\tau}\log g(\lambda_j;\hat{\tau}); \quad \lambda_j = \frac{2\pi\, j}{T}; \quad \hat{\tau} \;=\; \arg\min\sigma^2(\tau).$$

$\hat{a}$ and $\hat{A}$ in the above expressions are obtained through the first and second derivatives of the log-likelihood function with respect to d (see Robinson, 1994b, p. 1422, for further details). $I(\lambda_j)$ is the periodogram of u_t evaluated under the null, i.e.:

$$\hat{u}_t \;=\; (1 \;-\; L)^{d_o}\; y_t \;-\; \hat{\beta}'w_t;$$

$$\hat{\beta} \;=\; \left(\sum_{t=1}^{T} w_t\, w_t'\right)^{-1}\sum_{t=1}^{T} w_t\,(1 \;-\; L)^{d_o}\; y_t; \qquad w_t \;=\; (1 \;-\; L)^{d_o}\, z_t; \qquad z_t = (1,\, t)^T,$$

and g is a known function related to the spectral density function of u_t. Robinson (1994b) established that under certain regularity conditions:[5]

$$\hat{r} \;\;\rightarrow_d\;\; N(0,1) \qquad as \quad T \rightarrow \infty, \tag{6}$$

[4] Gil-Alana (2004c) showed that the model of Bloomfield (1973) approximates fairly well ARMA(p,q) structures in the context of fractional integration, where p and q are small values.

[5] These conditions are very mild regarding technical assumptions which are satisfied by (2) – (4).

and also the Pitman efficiency of the tests against local departures from the null.

[INSERT TABLE 1 ABOUT HERE]

Table 1 reports the estimates of d based on the Whittle function in the frequency domain along with the 95% confidence intervals of the values of d_o for which H_o cannot be rejected using Robinson's (1994b) parametric approach, the range of values considered for d_o being from 0 to 2 with 0.01 increments. It can be seen that in the case of white noise errors the unit root null (i.e. d = 1) is rejected in the three cases of no regressors, an intercept and a linear trend in favour of higher orders of integration. The Whittle estimate of d is found to be 1.27 in the three cases. However, when allowing for possible autocorrelation, the estimated values of d are substantially smaller in all cases, and the unit root null hypothesis is now rejected in favour of smaller degrees of integration, implying mean-reverting behaviour. If u_t is AR(1) the estimated values of d range between 0.71 and 0.90, whilst when using the exponential spectral model of Bloomfield (1973) they are slightly higher, ranging from 0.76 to 0.93. Thus, mean reversion (d < 1) is obtained in all cases if autocorrelation is permitted.

Given the similarities between the results for the two cases of an intercept and an intercept with a linear time trend, it appears that a time trend might not required for these series. Thus, it may be of interest to consider a joint test of:

$$H_o : \; d \; = \; d_o \text{ and } \; \beta \; = \; 0. \qquad (7)$$

This possibility is not addressed in Robinson (1994b) though Gil-Alana and Robinson (1997) derived a LM test of (7) against the alternatives:

$$H_a : \; d \; \neq \; d_o \text{ or } \; \beta \; \neq \; 0, \qquad (8)$$

as follows. We consider the regression model (2) with the vector partitions $z_t = (z_{At}^T, z_{Bt}^T)^T$, $\beta = (\beta_A^T, \beta_B^T)^T$, and we want to test H_o: $d = d_o$ and $\beta_B = \beta_{B0}$. Then, a LM statistic may be shown to be $\hat{r}^2$ (see equation (5)) plus

$$\sum_{t=1}^{T} \hat{u}_t\, w_{Bt}^T \left(\sum_{t=1}^{T} w_{Bt} w_{Bt}^T - \sum_{t=1}^{T} w_{Bt} w_{At}^T \left(\sum_{t=1}^{T} w_{At} w_{At}^T \right)^{-1} \sum_{t=1}^{T} w_{At} w_{Bt}^T \right)^{-1} \sum_{t=1}^{T} \hat{u}_t w_{Bt} \qquad (9)$$

with $w_t = (w_{At}^T, w_{Bt}^T)^T = w_t = (1 - L)^{do} z_t$,

$$\hat{u}_t = (1-L)^{d_o}\, y_t - (\tilde{\beta}_A{}^T, \beta_{Bo}{}^T)\, w_t; \quad \tilde{\beta}_A = \left(\sum_{t=1}^{T} w_{At}\, w_{At}^T \right)^{-1} \sum_{t=1}^{T} w_{At}\, (1-L)^{d_o}\, y_t,$$

$\hat{\sigma}^2 = T^{-1} \sum_{t=1}^{T} \hat{u}_t{}^2$, and $\hat{r}^2$ is calculated as above but using the $\hat{u}_t$ just defined. If the dimension of z_{Bt} is q_B, then we compare (9) with the upper tail of the $\chi^2_{1+q_B}$ distribution. In our case, testing (7) against (8) in (2) and (3) with $z_t = (1, t)^T$, we have $q_B = 1$, $z_{At} = 1$, $z_{Bt} = t$ for $t \geq 1$. Although we do not report the results, the time trend was found not to be necessary in the three cases.

In what follows we examine if the fractional differencing parameter has remained constant over time. For this purpose we first re-compute the estimates of d along with Robinson's (1994b) tests, for different sample sizes, starting with 121 observations (July, 1954 – July 1964), and adding successively one observation at a time till the end of the sample (645 observations, from July, 1954 – March, 2008). We consider the model with an intercept, and the results for the three cases of white noise, AR(1) and Bloomfield disturbances are displayed in Figure 2.

[INSERT FIGURE 2 ABOUT HERE]

Starting with the white noise case (Figure 2(i)), it can be seen that the values are above 1 in all cases, increasing to 1.36 in the sample ending in November 1979 and dropping fast to 1.11 in May 1980. Then, the values stabilise around 1.26 till the end of the sample. When u_t is assumed to be autocorrelated, either as an AR process or using the Bloomfield model, (Figures 2(ii) and (iii)), the results again indicate a fall in the degree of persistence around May 1980. In fact, the values are around 1.15 before that day and drop drastically to around 0.78 afterwards. Overall, the results presented so far suggest the occurrence of a structural break around May 1980, with a sharp decline in the degree of persistence after that date.

The above results, however, are not directly comparable because of the different sample sizes used to obtain the recursive estimates, ranging from 120 to 645 observations. Therefore, we also take a rolling approach using a fixed-size window of 120 observations which shifts along the sample by one month at a time, the first subsample being July 1954 to August 1964, and so on until the final subsample from March 1998 to March 2008. The results are shown in Figure 3.

[INSERT FIGURE 3 ABOUT HERE]

Starting again with the white noise case, one can see that all the estimated values of d are above 1. They increase from 1 till almost 1.50 over the first 184 subsamples. Then there is a sharp decrease at subsample 185 (November, 1979 – November 1989), the estimate of d being only slightly above 1.20 until subsample 312 (May 1980 – May 1990); subsequently, it starts rising again.

Next we consider the case of autocorrelated disturbances and show in the bottom part of Figure 3 the values for the Bloomfield (1973) case. Again there is a substantial change in the estimated values of d for the subsamples ranging from 185 (November, 1979 – November 1989) to 312 (May 1980 – May 1990), suggesting once more the existence of a break around these dates.

Given the possibility of a structural break in the data, we then perform a procedure due to Gil-Alana (2008) that enables us to estimate the date of the structural change along with the fractional differencing parameters and their associated coefficients for each subsample. This model is based on the following model,

$$y_t = \alpha_1 + x_t; \quad (1 - L)^{d_1} x_t = u_t, \quad t = 1,...,T_b \tag{10}$$

$$y_t = \alpha_2 + x_t; \quad (1 - L)^{d_2} x_t = u_t, \quad t = T_b+1,...,T, \tag{11}$$

where the α's are the coefficients corresponding to the intercepts; d_1 and d_2 can be any real value, and represent the orders of integration for each subsample, u_t is I(0) and T_b is the time of the break that is assumed to be unknown. This method is based on minimising the sum of squared residuals for a grid of (d_1, d_2) values. Using T_b-values from 50 (August, 1958) to 550 (June, 2004), the break was found to occur at observation 311, corresponding to May 1980, which is completely consistent with the results reported above. The resulting parameter estimates for the two cases of white noise and AR(1) disturbances are displayed in Table 2.

[INSERT TABLE 2 ABOUT HERE]

We note that when the disturbances are modelled as white noise the fractional differencing parameters are 1.28 and 1.49 for the first and second subsamples respectively. Thus, we observe an increase in the degree of dependence after the break. However, if u_t is assumed to follow an AR(1) process, the opposite happens, and the value of d decreases after the break: it is above 1 (1.32) before May 1980, and equal to

0.93 afterwards. We performed Beran's (1992) goodness-of-fit test for each of the four subsamples, and evidence in favour of the two autocorrelated models was found. Overall, it seems that the integration parameter representing the degree of persistence in the US Federal Funds effective rate has decreased from a value above 1 in the subsample ending in May 1980 to one strictly below 1 afterwards.

3. Conclusions

This paper analyses persistence in US interest rates. Specifically, using monthly data from July 1954 through March 2008, we estimate the order of integration of the US Federal Funds effective rate in a fractional integration framework. Moreover, unlike previous studies, we examine whether the degree of persistence of interest rates is stable over time. The results appear to be very sensitive to the specification chosen for the I(0) error term: when this is assumed to be a white noise, the fractional differencing parameter is found to be strictly above 1, while it is estimated to be strictly smaller than 1 in the case of autocorrelated disturbances. Further, recursive and rolling estimates indicate that the fractional differencing parameter has not remained constant over time, a sharp decline in the degree of persistence having occurred after May 1980. The existence of a structural break in May 1980, with the order of integration substantially decreasing in the presence of autocorrelated disturbances, is also confirmed by carrying out a procedure that enables us to determine the break date endogenously within a fractionally integrated framework (see Gil-Alana, 2008). In fact this break corresponds to the beginning of the announced Volcker disinflation, which resulted in greater instability in the response of interest rates to inflation. The reasons behind this instability and its policy implications will be investigated in future research.

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Figure 1: Federal Funds Effective Rate

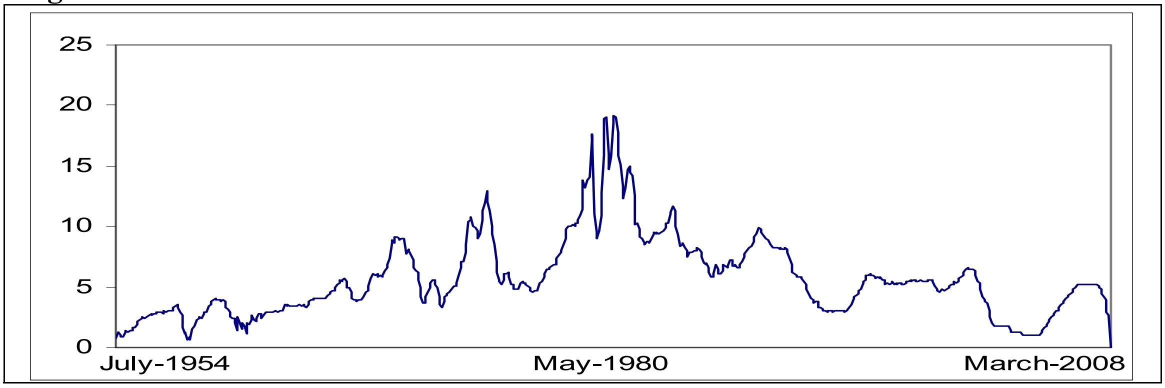
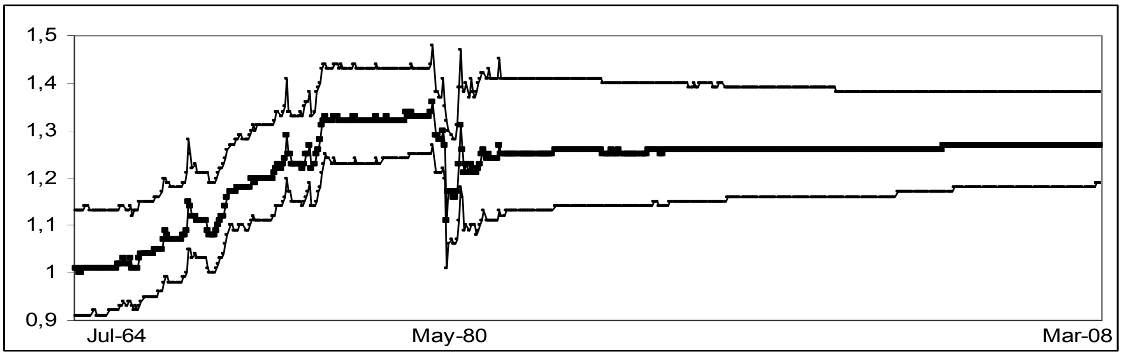


Table 1: Estimates of d using the whole sample

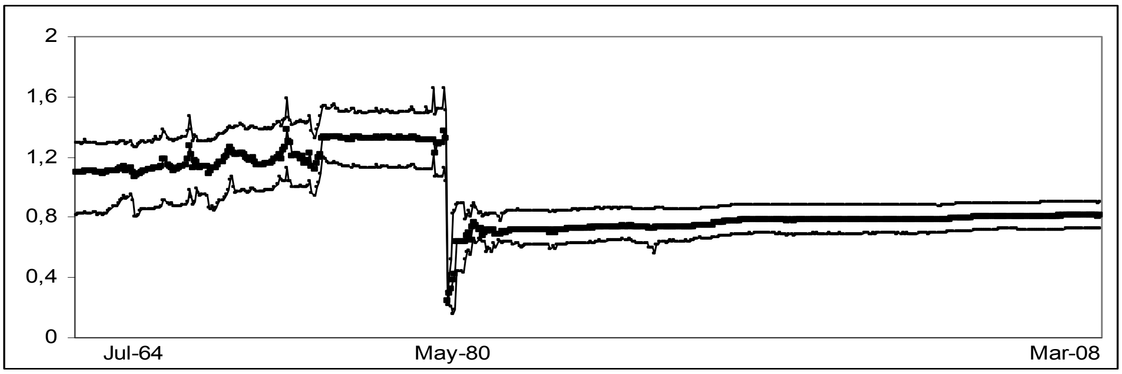
	No regressors	An intercept	A linear time trend
White noise	[1.18 **(1.27)** 1.38]	[1.19 **(1.27)** 1.38]	[1.19 **(1.27)** 1.38]
AR (1)	[0.71 **(0.79)** 0.88]	[0.73 **(0.80)** 0.90]	[0.73 **(0.80)** 0.90]
Bloomfield (m = 1)	[0.76 **(0.83)** 0.93]	[0.76 **(0.83)** 0.93]	[0.76 **(0.83)** 0.93]

Figure 2: Recursive Estimates of d and 95% Confidence Bands

i) White noise case



ii) AR(1) case



iii) Bloomfield (m = 1) case

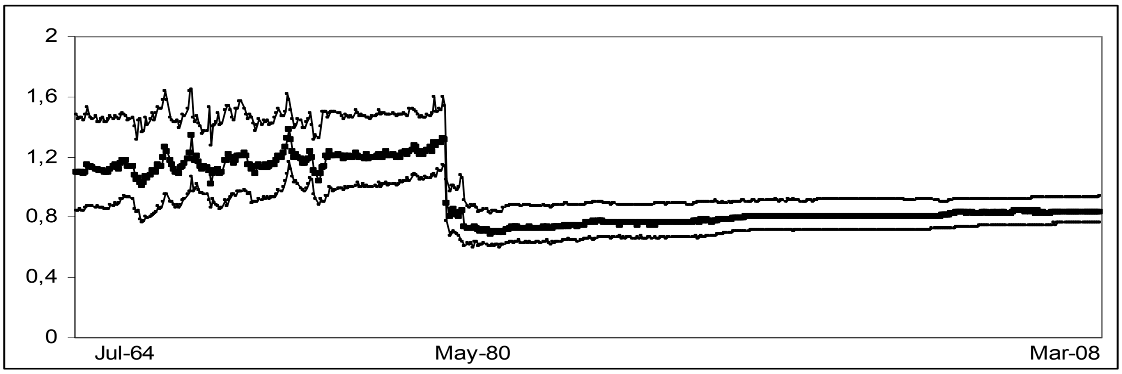
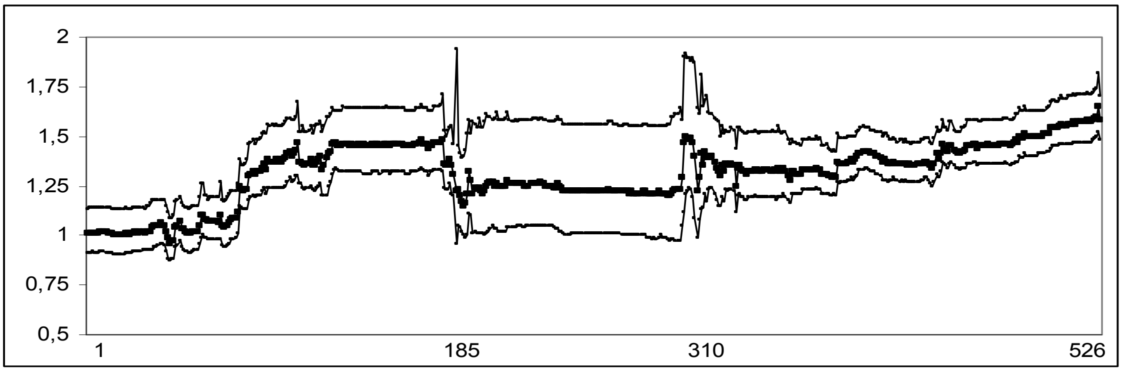


Figure 3: Rolling Estimates of d and 95% Confidence Bands

i) White noise case



ii) Bloomfield (m = 1) case

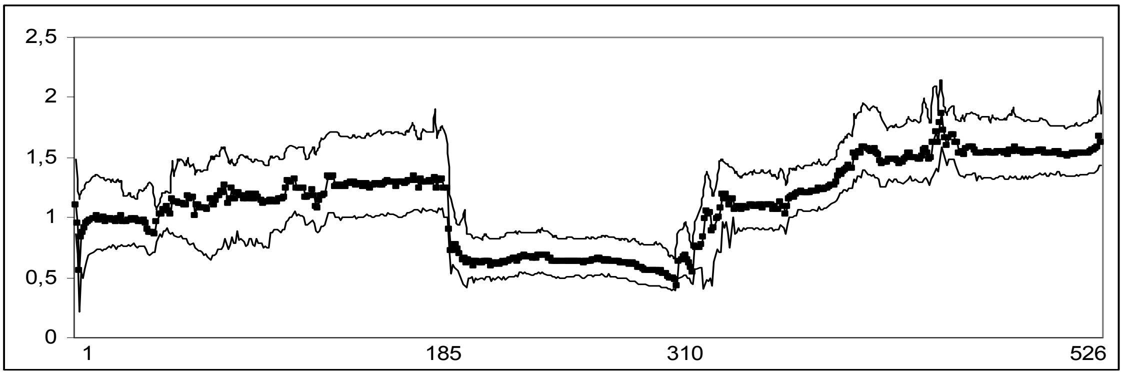


Table 2: Estimates of d and the associated parameters in a model with a single break

T* = 311 (May, 1980)	First subsample			Second subsample		
	d	intercept	AR coeff.	d	intercept	AR coeff.
White noise	1.28	0.713 (1.821)	-----	1.49	11.421 (27.311)	-----
AR (1)	1.32	-0.022 (-0.106)	-0.067	0.93	14.071 (16.251)	0.566