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ARE THE BALTIC COUNTRIES READY TO ADOPT THE EURO?
A GENERALISED PURCHASING POWER PARITY APPROACH

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Abstract

This paper focuses on macroeconomic interdependencies between the Euro area and three transition economies (Estonia, Lithuania and Latvia), with the aim of establishing whether the latter are ready to adopt the Euro. The theoretical framework is based on the Generalised Purchasing Power Parity (GPPP) hypothesis, which is empirically tested within a Vector Error Correction (VEC) model. Using both monthly and quarterly data over the period 1993-2005, it is found that GPPP holds for the real exchange rate vis-à-vis the Euro of each Baltic country, reflecting a degree of real convergence consistent with Optimum Currency Area criteria. Further, the adopted joint modelling approach for the real exchange rates of the Baltic region outperforms a number of alternative models in terms of out-of-sample forecasts.

Keywords: *Transition economies, Euro area, (Generalised) Purchasing Power Parity, Vector Error Correction models*

JEL Classification: *C32, E00, F36.*

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1. Introduction

Following the European Union (EU) enlargement of 2004, the next step in the European integration process is expected to be the (possible) expansion of the European Monetary Union (EMU) to allow the new member states to join the Euro. How fast (slow) this process should be depends on several factors affecting both old and potential candidate members. What motivates the transition economies of Eastern Europe (Balkan, Baltic, Central Europe, and Commonwealth of Independent States (CIS)[1] countries) to seek EMU membership is their expectation of being able to enjoy the benefits of a monetary union, in particular policy discipline, deeper economic and financial integration with the incumbent countries, a reduction in transaction costs, and lower interest rates (Lättemäe and Randveer 2004).

Of these economies, Estonia, Latvia and Lithuania appear to be best prepared to adopt the Euro, especially in terms of nominal convergence, with Lithuania seeming to be the strongest candidate at present. As pointed out by Havrylyshyn and Wolf (1999), the Baltic countries had a higher rate of growth compared to other transition economies, owing to better initial conditions such as less distortions and their being closer to market systems. Although some reforms began in the last years of the Soviet Union, the main transition process in these economies started at the end of 1991, when they achieved political independence. According to Wachtel and Korhoen (2004), within the whole package of political and economic reforms, a key role was played by the adoption of fixed exchange rate regimes, which resulted in credibility for both the new currencies and the economic system as a whole and lower and stable inflation. Further, both their degree of openness and foreign direct investment flows contributed to their success relative to the other transition countries.

The issue of how ready for EMU membership the accession countries are can be addressed using two main sets of criteria: institutional ones, i.e. the convergence criteria of the Maastricht Treaty, and economic ones (Breuss et al., 2004). As fiscal policy in the Baltic countries in recent years has generally been consistent with the Maastricht criteria, joining EMU will not require significant further changes in this respect, and thus the benefits from EMU membership should outweigh the costs. In particular, their fiscal parameters broadly meet the EU requirements, which reflects policy choices being made during the transition process to achieve exchange rate stabilisation as well as the control of the inflation.

[1] The CIS includes the former Soviet republics of Armenia, Azerbaijan, Belarus, Georgia, Kyrgyz and Uzbekistan.

From a more theoretical perspective, price convergence represents a necessary condition to stabilise both the nominal (explicit policy target) and the real exchange rate (implicit policy target), safeguarding the intra-regional competitiveness of member countries and avoiding incentives to implement "beggar-thy-neighbour" policies. However, Ahn et al. (2002) stress that a Purchasing Power Parity (PPP)-based approach may capture only partially the significant changes in economic policies and the restructuring process in Europe. A more useful framework is given by the Generalised PPP (GPPP) theory, which suggests that the (possible) non-stationarity of real exchange rates may be related to that of their long-run macroeconomic determinants. Empirically, GPPP between the EMU and the Baltic countries holds if it is possible to identify (at least) one stationary linear combination of bilateral real exchange rates vis-à-vis the Euro. The existence of an equilibrium path for a linear combination of the real exchange rates rules out real asymmetries (Bayoumi and Taylor, 1995), and therefore the GPPP hypothesis can be interpreted in terms of long-run sustainability of a common currency area (see Mundell, 1961).

This paper aims to establish whether economic conditions in these three Baltic economies are in fact such that they are ready to join the Euro area, putting an end to the transition phase of the last fifteen years. Cointegration techniques are applied to test the GPPP hypothesis, after investigating the stationarity of each real exchange rate. The adopted econometric framework is a Vector Error Correction (VEC) model. More specifically, this paper aims to analyse in sequence: a) whether the GPPP hypothesis holds between the Euro area and the Baltic region; b) the short-run dynamics of real exchange rates for the three Baltic countries; c) the relative contribution of global and regional shocks to the international competitiveness of those economies; d) the out-of-sample forecasting performance of our multivariate time-series model relative to a number of alternative models. The overall picture emerging from the estimates using both monthly and quarterly data over the period 1993-2005 suggests that the GPPP hypothesis is not be rejected by the data. Global (symmetric) shocks seem to explain most of the real exchange rate dynamics of the Baltic area, even though the difference in the size of regional shocks between countries is not negligible. Further, the adopted specification generally outperforms rival models describing the behaviour of the real exchange rates of the Baltic countries.

The paper is organised as follows. In Section 2 the theoretical framework is outlined. Section 3 presents the estimations results. In Section 4, the results from dynamic simulations based on impulse response analysis are discussed, while robustness and forecasting analysis are reported in Section 5. Some final remarks follow in the concluding Section 6.

2. Some Theory

The desirability for a given country of joining a common currency area is generally assessed through “cost-benefit analysis”. Potential gains are mainly thought of in terms of higher economic efficiency, whereas potential losses are mostly attributed to the loss of macroeconomic policy instruments (such as the exchange rate) to respond to asymmetric shocks (see, among others, Mongelli, 2002). However, despite some recent developments (see, for example, Demopoulos and Yannacopoulos, 1999), the optimum currency area (OCA) paradigm, originally proposed by Mundell (1961), still does not provide any formal criterion to establish the optimal timing and ways of creating a currency area (Eichengreen, 1990). Recent contributions focus on the role of public finances instead of the role of real exchange rates as an output stabiliser. From this perspective, whether domestic money can be used by the government as a tool for budgetary finance through seigniorage rents or whether countries differ in the costs associated with the collection of taxes are crucial issues in establishing the conditions to join a currency union (Buiter and Eaton, 1984; Canzoneri and Rogers, 1990; Aizenman, 1992). Most researchers point out that these factors can hardly be identified empirically (Baldwin, 1991; Buiter, 2000). Moreover, there is disagreement on the economic effects of monetary integration with respect to income correlation among member countries and intra-area trade flows.[2]

Following Rose (2000), a number of studies have used gravity models (de Nardis and Vicarelli, 2003, among others) to analyse the effect of joining a common currency area by testing the hypothesis of growing intra-area trade flows induced by the introduction of a common currency. Here, instead, we focus mainly on the role of cross-country income correlations to establish whether the Baltic countries are ready to join the Euro area, this being the criterion advocated by Artis (2003) and other authors to decide upon the adoption of a common currency.

More in detail, our theoretical framework is based on Enders and Hurn (1994); it enables us to obtain a measure of convergence (through empirical validation of the GPPP

[2] The “specialization hypothesis” (Krugman, 1993, and Krugman and Venables, 1996) postulates that as countries become more and more integrated, their industrial structure will develop according to their comparative advantages (Eichengreen and Bayoumi, 1996). Consequently, the economic systems of each member country of an OCA should become more sectorally concentrated and vulnerable to supply shocks. Opposite implications arise from the “endogeneity hypothesis” (Frankel and Rose, 1997), which posits that a positive link between income correlation and trade integration exists, suggesting that countries joining a currency union may satisfy the properties of an OCA *ex-post* even if they do not *ex-ante*.

hypothesis) between the Baltic countries and the Euro area. According to the GPPP theory, bilateral real exchange rates which are individually non-stationary might be cointegrated if their long-run macroeconomic determinants are highly correlated. In that case, the parameters of the cointegrating vector(s) depend on the functional form of the national aggregate demand functions. If it is possible to identify (at least) one stationary linear combination of otherwise non-stationary real exchange rates, then the long-run equilibrium condition(s) can be interpreted in terms of the economic interdependencies between the Euro area and Baltic countries, such as commercial and financial transactions, technology transfers and migration flows.

2.1. A reduced-form model for the Baltic economies

As already mentioned, the Baltic countries have some distinguishing features with respect to the other Eastern European transition economies. After their separation from the former Soviet Union most of their trade was re-oriented towards the neighbouring Western European countries. Moreover, the main explicit policy objectives were defined taking into account the possibility of joining EMU in the short term. However, links with the US also played an important role in the transition period: for instance, Lithuania pegged its currency to the US Dollar and external debt was mainly denominated in US Dollars.[3]

Thus, our chosen framework analyses interdependencies between the Baltic economies allowing for the possible influence of both the US and the Euro area economies. Accordingly, we consider five economies (indexed by i= 0,…,4): the US (0), the Euro area (1), Estonia (2), Latvia (3), and Lithuania (4). The US is the reference country in the model. For each country i, we specify the following equation for aggregate output:

$$y_{it} = f\left(y_{it}^{*}, q_{it}^{*}, i_{0t}\right) \text{ for } i = 0,...,4 \qquad (1)$$

where y_{it} stands for domestic output; $y_{it}^{*} = \sum_{j=0, j\neq i}^{4} \xi_{ij} \cdot y_{it}$ indicates foreign output, calculated as a weighted average of other countries' output; $q_{it}^{*} = \sum_{j=0, j\neq i}^{4} \upsilon_{ij} \cdot q_{ijt}$ is the real effective exchange rate, constructed as a weighted average of the bilateral real exchange rates $q_{ij,t}$ between

[3] Estonia was the first to introduce in 1992 a new currency, the *kroon,* which was pegged to the German Mark through a currency board. In Lithuania the new currency was introduced in 1993. The higher inflation with respect to the other Baltic countries forced the authorities to adopt a currency board and the *litas* was linked to the US Dollar. In Latvia, after an initial period in which the temporary currency had a fixed parity with respect to the Russian rouble, in 1992 the authorities decided to adopt a managed float regime in order to prevent the hyperinflation experienced by Russia and then the *lat* was pegged to the IMF's Special Drawing Rights (SDRs).

country i and country j; ξ_{ij} and υ_{ij} are country-specific weights; i_{0t} represents the interest rate of the reference country and is an indicator of the monetary policy stance prevailing in the world at time t. Assuming a linear specification for each equation in (1), we have:

$$y_{it} = a_i \cdot y_{it}^* + b_i \cdot q_{it}^* + c_i \cdot i_{0t}$$

or equivalently:

$$y_{it} = a_i \cdot \sum_{j=0, j\neq i}^{4} \xi_{ij} \cdot y_{it} + b_i \cdot \sum_{j=0, j\neq i}^{4} \upsilon_{ij} \cdot q_{ijt} + c_i \cdot i_{0t} = \sum_{j=0, j\neq i}^{4} \eta_{ij} \cdot y_{it} + \sum_{j=0, j\neq i}^{4} \psi_{ij} \cdot q_{ijt} + c_i \cdot i_{0t} \qquad (2)$$

where $\eta_{ij} \equiv a_i \cdot \xi_{ij}$ and $\psi_{ij} \equiv b_i \cdot \upsilon_{ij}$.

2.2. Empirical implementation

We follow Enders and Hurn (1994) and use the multivariate specification of the five equations in (2) where aggregate demand in each country only depends on factors driving demand in the reference country. As Appendix A shows, it is possible to obtain the reduced-form solution for the four independent bilateral real exchange rates vis-à-vis the reference country:

$$\mathbf{q}_{0t} = \underset{4\times 5}{\mathbf{C}} \cdot \mathbf{y}_t \qquad (3)$$

where $\mathbf{C}$ is a full rank matrix.

Next, we assume that $\mathbf{y}_t$ evolves over time according to a Vector AutoRegressive (VAR) process of order p

$$\mathbf{y}_t = \sum_{l=1}^{p} \mathbf{A}_l^y \cdot \mathbf{y}_{t-l} + \boldsymbol{\varepsilon}_t$$

or, in its isomorphic Vector Error Correction (VEC) form

$$\Delta \mathbf{y}_t = \mathbf{\Pi}^y \cdot \mathbf{y}_{t-1} + \sum_{l=1}^{p-1} \mathbf{P}_l^y \cdot \Delta \mathbf{y}_{t-l} + \boldsymbol{\varepsilon}_t \quad , \qquad \boldsymbol{\varepsilon}_t \sim N\left(\mathbf{0}, \mathbf{\Sigma}_\varepsilon\right) \qquad (4)$$

where $\mathbf{\Sigma}_\varepsilon$ is the time-invariant variance-covariance matrix associated to the vector of residuals $\boldsymbol{\varepsilon}_t$. Combining condition (4) and (5), we obtain the k-variate baseline model used in the empirical analysis, which describes the joint dynamics of the bilateral real exchange rates vis-à-vis the reference country:

$$\Delta \mathbf{q}_{0t} = \mathbf{\Pi}^q \cdot \mathbf{q}_{0,t-1} + \sum_{l=1}^{p-1} \mathbf{P}_l^q \cdot \Delta \mathbf{q}_{0,t-l} + \mathbf{u}_t \qquad (5)$$

where $\mathbf{u}_t = \mathbf{C} \cdot \boldsymbol{\varepsilon}_t$, $\mathbf{A}_l^q \cdot \mathbf{C} = \mathbf{C} \cdot \mathbf{A}_l^y$, and $\mathbf{\Pi}^q \cdot \mathbf{C} = \mathbf{C} \cdot \mathbf{\Pi}^y$.

VEC modelling builds on the association between the economic concept of *long-run* and the statistical concept of *stationarity* and focuses on the identification of stationary linear combinations of the data, known as cointegration vectors. In the presence of cointegration $\mathbf{\Pi}^q$ has reduced rank $r < k = 4$ and can be decomposed as $\mathbf{\Pi}^q = \boldsymbol{\alpha} \cdot \boldsymbol{\beta}'$, where the matrix $\boldsymbol{\alpha}$ contains the feedback coefficients (loadings) and the matrix $\boldsymbol{\beta}$ the $r < 4$ theory-based long-run relationships to which the series converge, once all the effects of transitory shocks have been absorbed (Johansen, 1995). These cointegrating relationships are hit by $4 - r$ permanent shocks (the common trends).

On the basis of the rank of matrix $\mathbf{\Pi}^y$, and, thus, of matrix $\mathbf{\Pi}^q$, it is possible to distinguish between five different cases, which can be associated to a relevant economic hypothesis, as detailed in Table 1. All possible outcomes share the common feature that the joint dynamics of the real exchange rate (the observed variable) closely mirror those of domestic output (the relevant variables for policy considerations on the feasibility of a common currency area).

[TABLE 1]

Assume, for instance, that $p = 1$ and $\mathbf{A}_1^y$ is an identity matrix. System (4) simplifies to $\mathbf{y}_t = \mathbf{I}_5 \cdot \mathbf{y}_{t-1} + \boldsymbol{\varepsilon}_t$. In this case, all the elements of the vector $\mathbf{y}_t$ are unit root processes, and the rank of matrix $\mathbf{\Pi}^y$, as well as of matrix $\mathbf{\Pi}^q$, will be equal to zero. In other words, the only way to deal with a stationary system [in terms of model (4) or (5)] is by first-differencing. Now, assume instead that $\mathbf{A}_1^y$ is a null matrix: $\mathbf{\Pi}^y$ will be equal to (minus) $\mathbf{I}_5$, i.e. it will be a full-rank matrix, consistently with the case of all real exchange rates being mean-reverting processes. Finally, any intermediate result between the full and the null rank assumption for matrix $\mathbf{\Pi}^y$ (or equivalently, matrix $\mathbf{\Pi}^q$) represents a validation of what Enders and Hurn (1994) call the GPPP hypothesis.

3. Estimation results

3.1. Data description and model specification

Monthly data for the real exchange rates of the Euro, and the currencies of Estonia, Latvia and Lithuania vis-à-vis the US are used to estimate system (5) over the period 1993:1-2005:12. Both nominal exchange rates and price data for the US and the Baltic countries are from the IMF's International Financial Statistics (IFS), while those for the Euro area are from the Eurostat Newcronos database. The real exchange rate is defined as the product of the nominal exchange rate (national currency per US dollars) and the ratio between US and domestic prices. Thus, an increase (decrease) in the real exchange rate means a real depreciation (appreciation). All the variables are expressed in constant prices (base year 2000=1).

As a preliminary step, we test for unit root behaviour of each of the four series. ADF (Dickey and Fuller, 1979) tests as well as unit root tests allowing for unknown breaks are performed on the four series, both in levels and first differences (upper part of Table 2).[4] The deterministic component includes an intercept and, when statistically significant, a linear trend, while the number of lags is chosen such that no residual autocorrelation is evident in the auxiliary regressions. Consistently with many studies of the PPP hypothesis in transition countries (see, among others, Sarno and Taylor 2001 and Égert 2004), in each case, we are unable to reject the unit root-null hypothesis at conventional nominal significance levels, even when controlling for breaks in the series. On the other hand, first-differencing the series appears to induce stationarity. The KPSS (Kwiatkowski et al., 1992) stationarity tests, presented in the lower part of Table 2, corroborate these conclusions. Given the evidence of $I(1)$-ness for all individual real exchange rate series, testing for cointegration among them is the logical next step.

[**TABLE 2**]

Estimating (5) requires two steps. First, the lag length p is chosen such that the estimated residuals resemble the multi-normal distribution as closely as possible, this being an essential requirement for correct statistical inference. Second, the long-term component of the model is identified on the basis of the trace and the maximum eigenvalue test (Johansen, 1995).

The general-to-specific procedure, with maximum order of autoregression set to 12, suggests choosing $p = 9$. The results of the main univariate (Table 3, upper part) and multivariate (Table 3, lower part) diagnostic tests indicate that the estimated residuals match the multi-normal distribution in a satisfactory way both at the single equation and the system level.[5]

[4] Critical values for these tests are provided by MacKinnon (1996) and Lanne , Lütkepohl and Saikkonen (2002), respectively. We have also carried out alternative unit root tests (Philips and Perron, 1998) to check for robustness. The results are qualitatively similar (available from the authors upon request).

[5] Although some departures from normality and some ARCH effects are found in the case of Lithuania, Gonzalo (1994) and Rahbek et al. (2002) have shown that the estimates of a VAR model are usually robust to these forms

[**TABLE 3**]

Trace and maximum eigenvalue test statistics suggest the presence of three cointegration relationships in the system at the 5 percent significance level [Table 4 - Panel (a)].[6] In the remainder of the Table exclusion, stationary and weakly exogeneity tests are reported. Testing separately the null hypothesis of each coefficient being equal to zero suggests that all variables are statistically different from zero [Panel (b)]. Further, no variable is stationary by itself in the cointegration space, consistently with the univariate unit root and stationarity tests [Panel (c)]. Finally, except for the Euro/US dollar equation, there is no clear evidence of weak exogeneity [Panel (d)].

[**TABLE 4**]

3.2. Modelling the long-run: the structure of the β matrix

A key issue in the empirical investigation is establishing whether the cointegration vectors can be identified in terms of the structure described in Table 1 above. If the restrictions cannot be rejected by the data, then each vector in the β matrix validates the PPP condition for each Baltic countries *vis-à-vis* the Euro area. This implies the following long-run for model (5):

$$\boldsymbol{\beta}' \cdot \mathbf{q}_{0,t-1} = \begin{bmatrix} 1 & -1 & 0 & 0 \\ 1 & 0 & -1 & 0 \\ 1 & 0 & 0 & -1 \end{bmatrix} \cdot \begin{bmatrix} q_{01,t-1} \\ q_{02,t-1} \\ q_{03,t-1} \\ q_{04,t-1} \end{bmatrix}$$

Using a standard LR ratio test with 3 degrees of freedom following a χ^2 distribution, the test statistics (4.63), calculated using the Bartlett small-sample correction (with an estimated factor of 4.72), indicate that the restrictions are not rejected by the data at the usual significance levels (p-value of 0.20).[7]

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of misspecification.

[6] The choice of the cointegration rank is also robust to a graphical analysis of the recursive trace tests. Since the trace statistic is given by $-T_j \ln(1-\lambda_i)$, with $j = T_1,....,T_t$, it grows over time as long as $\lambda_i \neq 0$, while it must be constant if $\lambda_i \to 0$. In practice, the first r trace statistics should grow linearly, while the others must be constant over time. Graphs (not included, but available from the authors upon request) show that the first three statistics in fact grow linearly as expected, while the fourth one is less clearly increasing.

[7] Note that these conclusions are robust to possible parameter instability, as when we compute the LR test statistic recursively the results remain stable. Specifically, we performed the estimation for the sub-sample 1993:1 to 1997:9, imposing the restrictions on the three vectors and computing the value of the LR test statistic. Then we added recursively one observation at a time till until the end of the sample (2005:12) and in each step we re-calculated the value of the test statistic. Plotting these values, we note that the restrictions are supported by the data, since the statistics are always below the value of the χ^2 with 3 degree of freedom. (The graph of the

This is one of the key findings of the present paper, as the existing literature usually concludes that real exchange rates in the case of transition countries are not stationary (for a comprehensive review, see Égert 2004).[8] To model deviations from PPP, many authors invoke the dominance of Harrod-Balassa-Samuelson (HBS) effects for these economies, which, during the transition process, display high productivity growth in the traded goods sector (see, among others, Coricelli and Jazbec 2004).[9] In the case of the Baltic economies, De Broeck e Sløk (2001) argue that even if productivity gains in the tradeable sector may lead to a real exchange rate appreciation, this does not erode their competitiveness level because of the initial undervaluation of the national currencies and of the process of controlling inflation. Moreover, Kocenda (2001) suggests that the Baltic countries show a higher degree of convergence both amongst themselves and with the Euro area in the main macroeconomic fundamentals compared to other groups of transition economies. Consequently, the empirical results favouring the joint PPP hypothesis vis-à-vis the Euro area could reflect both the weak role played by the traditional factors (i.e. HBS effects) and robust convergence of the macroeconomics fundamentals, which is the key condition for GPPP and, thus, joining a currency union.

Figure 1 presents the cointegration relationships from the R-model.[10] There appears to be a clearly identifiable cointegrating relationship in the case of Estonia, especially after 1996. By contrast the years 1993-1995 are characterised by slightly higher variability, corresponding to the first phase of economic reform at the beginning of 1992. Similar considerations apply to the Lithuanian case, although the period of high variability appears to be longer (until the end of 1999) because of a slower transition process. The graph of the cointegrating relationship for Latvia is qualitatively similar to that of Lithuania, the higher variability in the first years most likely being related to the initial currency peg to the Russian

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recursive LR test statistics is not included but is available on request).

[8] In order to check the robustness of our findings, we have also carried out the analysis in the context of three bivariate VEC models, each including one Baltic country and the real exchange rate of the Euro vis-à-vis the US Dollar. The general-to-specific procedure suggests an autoregression order of nine, one and six for the Estonia-Euro, Latvia-Euro and Lithuania-Euro model (in their isomorphic VAR representations), respectively, and an unrestricted constant for all systems. For each model the trace test statistic indicates the presence of a unique cointegrating vector at the 5 percent significance level. Imposing restrictions of the form [1,-1] on each of these vectors leads to a value of the χ^2 statistic (corrected using the Bartlett factor) of 0.44 [0.80], 5.56 [0.06], 0.82 [0.36], respectively, where p-values are reported in squared brackets. These results are fully consistent with the evidence reported in the main text, giving further evidence of the validity of the PPP condition for each Baltic countries *vis-à-vis* the Euro area.

[9] For an overview of the impact of the Balassa-Samuelson effect on long-run PPP deviations, see Sarno and Taylor (2001).

[10] The R-model is computed estimating the VEC representation of the system deleting all dummies and the short -run dynamics. The result is a model where only the long-run properties of the data are isolated (see, Johansen, 1995).

ruble and, subsequently, to the IMF's SDRs with the aim of preventing the hyperinflation experienced by Russia. Only most recently has the lat been pegged to the Euro.

[**FIGURE 1**]

3.3. Adjustment towards equilibrium

Once the cointegration space is identified, the long-run properties of system (5) are analysed by looking at their persistence profiles (Pesaran and Shin, 1996), which make it possible to assess how long the system takes to revert to its steady-state path, after being hit by a system-wide shock. By construction these profiles should tend to zero as the number of simulation periods increases only if the cointegration vector is genuinely stationary, while in the case of $I(1)$ (or "near integrated") series these can be different from zero for a long period. Figure 2 presents the absorption path of deviations from the PPP relationship between each Baltic country and the Euro area, over a simulation horizon of 60 months (5 years).[11]

[**FIGURE 2**]

In all cases, we observe convergence towards the steady-state, with the adjustment process being completed within the fifth year of the simulation. The half-life of the deviation from the steady-state for Estonia and Lithuania is close to four months, while it is larger for Latvia (15 months).[12]

Persistence profiles are also a useful tool to calculate convergence loss measures (CL) for pair-wise comparisons of the speed of convergence, as pointed out in Girardi and Paesani (2008). Once the CL is computed, the null of equivalent loss between the two entities of reference can be tested by means a standard two-sided t-test.[13] Natural symmetric loss functions are the absolute loss and the squared loss. The CL obtained from an absolute loss over the chosen simulation horizon for Estonia, Latvia and Lithuania are equal to 0.14, 0.24 and 0.09, respectively. The test of equality yields a p-value of 0.24 for the Estonia-Lithuania simulation, while it is equal to 0.01 and 0.00 in the Estonia-Latvia and Lithuania-Latvia cases, respectively.[14] This confirms that Estonia and Lithuania follow a similar dynamic path, while

[11] The size of all the shocks analysed in this section is set equal to one standard deviation.

[12] Half-life is defined as the number of months which have to pass before the deviation from the steady-state falls to half the size of the initial shock.

[13] To do this, we calculate the deviation from the steady state f at the simulation horizon k. Given a non-negative loss function L, we define the CL as $\frac{1}{k^*}\cdot\sum_{k=0}^{k^*} L(f_k)$, where k^* is a truncation lag such that $f_{k^*} \approx 0$.

[14] A F-test on the variance ratios indicates that all the distributions are heteroscedastic.

the absorption process for Latvia is somewhat slower relative to the one of the two other Baltic countries.

4. Modelling interdependencies between the Baltic countries

This Section analyses possible spillovers in the adjustment towards the long-run equilibrium condition(s) by focusing on the structure of the feedback matrix $\boldsymbol{\alpha}$.[15] It also considers a structural representation of the multivariate time-series model in order to ascertain the role of global and idiosyncratic shocks hitting the Baltic countries.

4.1. Modelling the short-run: the structure of the α matrix

The short-run dynamics of (5) are modelled using a parsimonious (subset) VEC model, obtained dropping the parameters of the matrices $\boldsymbol{\alpha}$ and $\mathbf{P}_l^q$ with p-values lower than a threshold,[16] according to the Sequential Elimination of the Regressors Testing Procedure (SER/TP) proposed by Brüggemann and Lütkepohl (2001). Specifically, the statistically significant parameters of $\boldsymbol{\alpha}$ give useful information about how our regional model moves around the long-run equilibrium path. Moreover, the rows of $\boldsymbol{\alpha}$ containing only zeroes allow to identify possible (weakly) exogenous variables.

Table 5 reports the coefficients estimated by 3SLS. The LR test does not reject the 80 zero-restrictions. Of these, only six concern the $\boldsymbol{\alpha}$ matrix; however, each regressor is present in (at least) one equation of the system, which supports the chosen model specification. The analysis of the elements of the loading coefficients matrix allows to highlight some interesting results: i) since $\alpha_{\Delta eu}^{\varepsilon_1} = \alpha_{\Delta eu}^{\varepsilon_2} = \alpha_{\Delta eu}^{\varepsilon_3} = 0$, the bilateral Euro/US dollar real exchange rate is a weakly exogenous variable (forcing variable) for the long-run parameters; ii) Δee, Δlv and Δlt are obviously affected by the cointegration residuals (ε_1, ε_2 and ε_3 respectively) which identify the PPP relationships vis-à-vis the Euro area; iii) the (absolute) values of the feedback coefficients indicate that the speed of adjustment towards equilibrium is higher for Estonia

[15] This analysis of spillovers is consistent with a situation where, although each real exchange rate may behave differently in the long run, there may be important relations among them in the short run reflecting similar macroeconomic conditions and/or transition processes. This property is directly related to the well-known notion of separation and cointegration developed by Konishi and Granger (1993), Granger and Swanson (1996) and Granger and Haldrup (1997). Furthermore, if some rows of the associated feedback contain only zeros, then it might be possible to identify weakly exogenous variables driving the system in the long run (see Pesaran et al., 2000).

[16] The AIC criterion with $t = 1.60$ is used as a significance threshold level for the short-run parameters. This is motivated by the idea that, in the reduction process of the model, it is preferable to keep the coefficients whose statistical significance is unclear.

(about 8 percent per month); iv) furthermore, there is evidence of some influence of ε_1 and ε_3 on Δlv and Δee, respectively.

[**TABLE 5**]

Overall, these results suggest that the Baltic countries exhibit heterogeneous paths of convergence towards their own equilibrium, with Lithuania appearing to be less affected by deviations from the long-run equilibrium condition than the two other Baltic countries, while Estonian reform policies seem to have insured a quicker transition process.

4.2. From a reduced-form to a structural representation

The model reduction process has two further implications. Firstly, dynamic simulations may differ, even markedly, from those derived from an unrestricted model (Brüggemann and Lütkepohl 2001). Secondly, dropping statistically coefficients can improve the quality of the forecasts generated by the model (Clements and Hendry, 2001, p. 119). Here we focus on the former issue, while the latter is discussed in the following Section.

As emphasised in the traditional OCA literature (see Artis, 2003), detecting the sources of shocks affecting an economic system has important implications: if economies are hit by dissimilar disturbances to the ones hitting their partner countries, the cost of joining a currency union can be correspondingly large. Thus, we analyse to what extent global, regional and domestic economic conditions affect domestic real exchange rate variability, considered as proxies of output fluctuations.

To do this, we carry out forecast error variance decomposition (FEVD), which aims at providing information on the relative importance of the forecast error variance of each shock as a function of the simulation horizon. The reduced form residuals in model $\mathbf{u}_t$ and the structural residuals $\mathbf{v}_t$ are linked through the relationship $\mathbf{u}_t = \mathbf{B} \cdot \mathbf{v}_t$, where $\mathbf{B}$ is a non-singular matrix (Warne, 1993). Retrieving the v's from u's requires the unique determination of the $k^2 = 16$ elements in $\mathbf{B}$. In our identification scheme, a first set of 10 constraints arises from assuming that structural shocks are orthonormal. The cointegration space gives $r(k-r) = 3$ additional restrictions and allows to distinguish transitory shocks (three in our case) from permanent (one) innovations. The remaining 3 restrictions are obtained by imposing a recursive scheme on the matrix of the transitory shocks where the causal order of the variables is chosen according to the size of the adjustment coefficients estimated previously. Thus, the causal order is the following: Lithuania, Estonia and Latvia. The permanent shock is derived from the permanent component of the system (that is, the common trend) and represents the global-external shocks that hit in a symmetric way all Baltic countries. By contrast, transient impulses hit in an asymmetric way each country according to their different degree of interdependency. Furthermore, temporary shocks are aggregated so as to quantify the overall relevance of regional factors in explaining real exchange rate fluctuations.

Table 6 shows the percentage of the variance of each variable of the system explained by global, regional and idiosyncratic shocks, where the latter are expressed as a percentage of regional impulses. The last column (mean) presents the average contribution of the shocks over the entire simulation period (60 months).

[**TABLE 6**]

As can be seen, the disturbance from the Euro area (the global shock), which represents the symmetric shock hitting the Baltic region, is the main driving force of real exchange rate movements for all Baltic countries, especially in the case of Estonia, where it explains about 97.6% of the forecast variance, compared to 88.4% in Latvia and 74.6% in Lithuania. In general, this shock accounts for a considerable percentage of the variance of the whole system (86.9%).

The relative importance of regional shocks in explaining the dynamics of each real exchange rate differs across countries. Its size is negligible in Estonia (2.42%), while it is bigger in Latvia (11.5%) and even more so in Lithuania (25.25%). Moreover, for the latter country the percentage of the regional shock due to the idiosyncratic component is equal to 91%. This means that the asymmetric shock plays a big role, while it is smaller for the other two countries.

In summary, although the symmetric shock is the biggest source of the variability of the real exchange rate for all the countries, the relative importance of regional shocks and of their idiosyncratic components varies considerably between them.

5. Robustness and extensions

5.1 Evidence from quarterly data

As a robustness check of our findings, we re-estimate model (5) using quarterly data. Unit root tests, even in the specification with unknown structural breaks in the deterministic component, and stationarity tests confirm that the series are realisations of integrated processes of order 1.[17] On the basis of the general-to-specific procedure, the specification of our quarterly model includes three lags (in the levels), consistently with the pre-estimation period used in the monthly specification (nine lags in the levels). Table 7 summarises the main diagnostic tests. As before, departures from normality are found in the case of Latvia, but now no evidence of ARCH effects in the residual is detected in any equation.

[TABLE 7]

Both the trace and the maximum eigenvalue statistics indicate that there are three cointegration relationships [see Table 8, Panel (a)], that all variable are jointly significant in the cointegrating space, and that no variable can be considered weakly exogenous and stationary by itself [Panel (b), Panel (c), and Panel (d), respectively].

[TABLE 8]

The over-identifying restrictions discussed in Section 3 above are still valid at the quarterly frequency: the Bartlett-corrected (with estimated correction factor of 4.53) χ^2-distributed LR test statistics is equal to 2.97 with a p-value of 0.40.[18] This finding provides further evidence of the stationarity of the bilateral real exchange rate, and thus of the joint validity of the PPP condition between each Baltic country and the Euro area.[19]

The persistence profiles plotted in Figure 4 are roughly similar in the case of the monthly estimates. For all the equations, the period required to absorb the shock completely is five years, as in the monthly simulations. In the case of Estonia, the half-life is close to 4 quarters, and greater than that shown by the monthly estimates. However, in the other two cointegration vectors, which represent the Latvian and Lithuania long-run equilibrium relationships, the half-lives are close to 3 and 1 quarters respectively, very similar results to the earlier ones. The absolute convergence losses are equal to 0.20, 0.23 and 0.08 for Estonia, Latvia and Lithuania, respectively. The former two are similar to those obtained in the monthly estimates, while the latter (i.e. Estonia) is higher. This might depend on the first

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[17] The results (available on request) are not reported for the sake of brevity.

[18] The three cointegration vectors are plotted in Figure 3. As in the previous model with monthly observations, the Estonian vector seems to be stationary, especially after the first period of transition, while the other two show higher variability of the long-run equilibrium conditions, as a consequence of a slower transition process.

[19] As before (see footnote 8), we carried out the analysis using three bivariate VEC models in which we can study separately each real exchange rate of the Baltic countries and the real exchange rate of the Euro vis-à-vis the US Dollar. The lag length in the three models for Estonia, Latvia and Lithuania suggested by usual general-to-specific procedure is three, two and two, respectively. We also include an unrestricted constant in all models. The trace test statistic indicates the presence of a unique cointegrating vector at the 5 percent significance level, confirming the results obtained in the general model with four variables. The restrictions of the form [1,-1] are still valid for the three different vectors, with a value of the χ^2 statistic (corrected using the Bartlett factor) of 0.13 [0.72], 2.66 [0.11], 1.63 [0.20], respectively (p-values in squared brackets). This further confirms the robustness of the results discussed in the main text, both at the monthly and quarterly frequency.

period of simulation which in the case of Estonia seems to be characterised by higher variability with respect to both the other two countries and the simulation carried using monthly observations.

[FIGURE 4]

Table 9 reports the coefficients of the subset VEC model estimated by 3SLS using the SER/TP method where statistically irrelevant parameters are deleted according to the AIC criterion with a significance threshold level of $t = 1.60$. The 26 zero-restrictions are not rejected by the data, the p-value of the LR test statistic with a χ^2 distribution being equal to 0.30. The restrictions on the feedback matrix $\boldsymbol{\alpha}$ confirm that the Euro real exchange rate is weakly exogenous; furthermore, spillover effects seem to be richer, with evidence of feedbacks in the case of Estonia and Latvia, but not in the Lithuanian one. As before, the speed of adjustment towards equilibrium is markedly higher for Estonia.

[TABLE 9]

In summary, robustness analysis confirms the main results obtained from estimating the base model with monthly observations.

5.2 Out-of-Sample Forecasting

In this Section the out-of-sample forecasting performance of the model is evaluated in order to establish whether the estimated subset VEC models can describe satisfactorily real exchange rate movements for the Baltic countries over the period 2006:1-2007:12. In particular, we compare their forecasting performance against three alternative models: a similar VEC model without restrictions on the short-run parameters; a VAR including the three real exchange rates of the Baltic countries vis-à-vis the Euro; finally, a random walk[20]. The first comparison is with a model also including the variables dropped from our chosen specification on the grounds of statistical insignificance. The VAR models[21] are estimated to

[20] For each of the three real exchange rates, a random walk model with drift is estimated. The sample period goes from 1993:10 to 2005:12, and has been chosen for consistency with the VEC models, where first-differencing and the lag length chosen shorten the original sample by 9 observations.

21 The choice of a VAR as an alternative model follows directly from the statistical properties of the three Baltic real exchange rates with respect to the Euro. We consider a simple three-variable system of the real exchange rates of the Baltic currencies vis-à-vis the Euro, and we expect that each variable included in the system is stationary. In order to test this type of hypothesis, we estimate a VAR model with nine lags (in the levels) and an unrestricted constant term. Then, we set up the Taylor and Sarno's (1998) multivariate unit-root test, labelled as

evaluate the adequacy and the robustness of both the GPPP-representation and the restrictions on the cointegrating space. Finally, assessing the subset VEC model against a random walk model gives us information on how well the data are described by stationary processes rather than by unit root behaviour.

The dataset available at time 2005:12 (the last month of the estimation period) is used to project the cointegrating relations over the entire forecast horizon. Actual data from 2006:1 to 2007:12 are used to evaluate the system forecasts.

Table 10 shows the results of the forecasting exercises for the real exchange rates of Estonia (part a), Latvia (part b) and Lithuania (part c). The mean square error (MSE) and mean average error (MAE) values for the subset VEC model are given in levels (column 1), while for the others, as in Clarida and Taylor (1997), we consider their ratio for the VEC model and for each alternative model (column 3,4 and 5 in the table). Thus, a ratio of less than one indicates that the forecasting performance of the subset VEC specification is superior to the alternative ones. Moreover, p-values of the Diebold-Mariano test are also reported.

[TABLE 10]

For all three subset VEC models, the values of the MSE and the MAE are increasing with the size of the temporal window used for the forecast. Thus, for example, the value of the MSE (MAE) for Estonia rises from 1.29e-04 (9.57e-03) for the one-month horizon to 4.03e-03 (3.97e-02) for the twenty-four-month one. This is still true in the case of Latvia and Lithuania, where the values for the one-month horizon are 4.64e-04 (1.74e-02) and 1.68e-04 (1.00e-02), while for the twenty-four-month horizon they are 4.04e-03 (5.27e-02) and 1.24e-03 (3.15e-02), respectively.

The forecasting analysis for the Estonian real exchange rate shows that the subset VEC model generally outperforms the rival models. In particular, only in the case of the one- and six-month horizon do measure of forecasting accuracy indicate that the unrestricted VEC model is preferable.

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the JLR test, in which the null hypothesis is that at least one of the series in the system is non-stationary. As they suggest, we compare the test statistics with the critical values of a $\chi 2(1)$ distribution adjusted by the factor (T/T(-p-k)) to take into account the finite sample bias (where p is the number of lags included in the VAR and k is the number of countries). The results (not reported) confirm that there are three stationary series, which can be viewed as the PPP conditions for each of the Baltic countries *vis-à-vis* the Euro area.

In the Latvian case, the results also suggest that the forecasting performance of the subset VEC specification is superior. The ratios of the MSE and MAE for the subset VEC and the unrestricted VEC model is always less than one, while only in three cases[22] it is bigger than one when comparing the results of the subset VEC model to those of the VAR specification.

The results for the Lithuania real exchange rate are less supportive of the subset VEC specification. The restricted VEC model always outperforms the unrestricted specification, but the VAR model is found to be superior in most cases. Nevertheless, five out of eight times the Diebold-Mariano test indicates that the values are not statistically different. In all other cases this test confirms the previous results, always suggesting that the forecast accuracy of the subset VEC model is substantially better the ratio being less than one.

Finally, the subset VEC model generally beats the random walk model in the case of the Estonian and Latvian real exchange rates, while the forecasting accuracy of the *naïve* specification is always better in the case of the Lithuanian equation. In all cases, as expected, the unit root behaviour dominates the stationary subset VEC specification as the forecasting temporal window increases (that is, the value of the ratio increases).

Overall, the results of Table 11 provide strong support for the choice of the subset VEC model as the appropriate one for the real exchange rates of the Baltic countries.

6. Conclusions

This paper focuses on macroeconomic interdependencies between the Euro area and a group of three transition countries, namely Estonia, Lithuania and Latvia. Its aim is to test, using a theoretical framework based on the GPPP hypothesis, if the transition process undergone by these economies in recent years has made them strong candidates to join the Euro area in the near future. Cointegration techniques are used to test the GPPP hypothesis between each Baltic country and the Euro area. Persistence profiles and impulse responses are also computed in order to take into account the possible influence of global and regional shocks which hit the equilibrium conditions of each real exchange rate.

The results suggest that the GPPP hypothesis holds for each of the real exchange rates of the Baltic countries over the period 1993:1-2005:12, confirming that, as stressed De

[22] Note that in the case of the 12-month horizon the results are mixed. In fact, using the MAE criterion suggests that the forecast performance of the restricted VEC model is superior to that of the VAR one, while using the MAE criterion yields the opposite result.

Broeck e Sløk (2001), Harrod- Balassa-Samuelson effects have played a modest rule during the transition process of these economies, and that instead there has been significant real convergence towards the Euro area. In particular, we find evidence of more pronounced mean-reversion properties in the equation describing the behaviour of the Estonian kroon - Euro real exchange rate, with the presence of spillover effects between the equilibrium conditions, especially in the case of Estonia and Latvia. Quarterly data produce qualitatively similar results, confirming the robustness of all our main findings.

Moreover, forecast error variance decomposition analysis shows that symmetric shocks are the main driving force of the real exchange rates in the Baltic countries, with asymmetric shocks playing a more important role in explaining their variability in the case of Lithuania and Latvia (less so in Estonia), and out-of-sample forecasting analysis generally shows that the selected subset VEC specifications outperform rival models of the real exchange rates of the Baltic countries.

It is important to note that the empirical validation of the GPPP hypothesis is only one of possible criteria for an OCA - other studies have relied instead on measures such as rolling correlations or on structural VAR approaches to analyse the degree of synchronisation of shocks (for an extensive literature review, see Fidrmuc and Korhonen 2006). The fact that GPPP (or PPP as in our case) holds does not entail that the economies are synchronised, but it does imply that the exchange rates are driven by economic fundamentals and share common trends in the presence of a high degree of convergence. Therefore, although this is only an implicit convergence measure, it does suggest that the transition process in the Baltic economies has resulted in a stable and credible economic framework which makes these countries strong candidates for EMU membership.

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Appendix A

Assuming linearity for each equation in (2) and taking into account that $q_{ij,t} \equiv -q_{ji,t}$, $q_{ik,t} \equiv q_{ij,t} - q_{kj,t}$ by construction, we can rewrite the five equations in (2) as a multivariate system where aggregate demand in each country only depends on the factors driving demand in the reference country.

$$\underset{5\times5}{\mathbf{\Xi}} \cdot \mathbf{y}_t = \underset{5\times4}{\mathbf{\Gamma}} \cdot \mathbf{q}_{0t} + \underset{5\times1}{\boldsymbol{\gamma}} \cdot i_{0t} \qquad \text{(A1)}$$

where $\mathbf{y}_t = \begin{bmatrix} y_{0t} & y_{1t} & y_{2t} & y_{3t} & y_{4t} \end{bmatrix}'$, $\mathbf{q}_{0t} = \begin{bmatrix} q_{01} & q_{02} & q_{03} & q_{04} \end{bmatrix}'$, $\boldsymbol{\gamma}' = \begin{bmatrix} c_0 & c_1 & c_2 & c_3 & c_4 \end{bmatrix}$, $\mathbf{\Gamma} = \underset{5\times20}{\mathbf{\Psi}} \cdot \underset{20\times4}{\mathbf{F}}$, and

$$\mathbf{\Xi} = \begin{bmatrix} 1 & -\eta_{01} & -\eta_{02} & -\eta_{03} & -\eta_{04} \\ -\eta_{10} & 1 & -\eta_{12} & -\eta_{13} & -\eta_{14} \\ -\eta_{20} & -\eta_{21} & 1 & -\eta_{23} & -\eta_{24} \\ -\eta_{30} & -\eta_{31} & -\eta_{32} & 1 & -\eta_{34} \\ -\eta_{40} & -\eta_{41} & -\eta_{42} & -\eta_{43} & 1 \end{bmatrix},$$

$$\mathbf{\Psi} = \begin{bmatrix} \psi_{01} & \psi_{02} & \psi_{03} & \psi_{04} & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & \psi_{10} & \psi_{12} & \psi_{13} & \psi_{14} & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & \psi_{20} & \psi_{21} & \psi_{23} & \psi_{24} & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & \psi_{30} & \psi_{31} & \psi_{32} & \psi_{34} & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & \psi_{40} & \psi_{41} & \psi_{42} & \psi_{43} \end{bmatrix} \quad \text{and}$$

$$\mathbf{F}' = \begin{bmatrix} 1 & 0 & 0 & 0 & -1 & -1 & -1 & -1 & 0 & 1 & 0 & 0 & 0 & 1 & 0 & 0 & 0 & 1 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 & 1 & 0 & 0 & -1 & -1 & -1 & -1 & 0 & 0 & 1 & 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 1 & 0 & 0 & 0 & 1 & 0 & 0 & 0 & 1 & 0 & -1 & -1 & -1 & -1 & 0 & 0 & 0 & 1 \\ 0 & 0 & 0 & 1 & 0 & 0 & 0 & 1 & 0 & 0 & 0 & 1 & 0 & 0 & 0 & 1 & -1 & -1 & -1 & -1 \end{bmatrix}.$$

Recasting system (A1) as

$$\underset{5\times5}{\mathbf{\Xi}} \cdot \mathbf{y}_t = \begin{bmatrix} \underset{5\times4}{\mathbf{\Gamma}} & \vdots & \underset{5\times1}{\boldsymbol{\gamma}} \end{bmatrix} \cdot \begin{bmatrix} \mathbf{q}_{0t} \\ \cdots \\ i_{0t} \end{bmatrix}$$

we obtain

$$\begin{bmatrix} \mathbf{q}_{0t} \\ \cdots \\ i_{0t} \end{bmatrix} = \begin{bmatrix} \underset{5\times4}{\mathbf{\Gamma}} & \vdots & \underset{5\times1}{\boldsymbol{\gamma}} \end{bmatrix}^{-1} \cdot \underset{5\times5}{\mathbf{\Xi}} \cdot \mathbf{y}_t = \underset{5\times5}{\mathbf{C}^*} \cdot \mathbf{y}_t \qquad \text{(A2)}$$

Tables and Figures

Table 1: Long-Run Restrictions

	Matrix			Hypothesis to test	Economic interpretation
	$\mathbf{\Pi}^y$	$\mathbf{C}\cdot\mathbf{\Pi}^y = \mathbf{\Pi}^q\cdot\mathbf{C}$	$\mathbf{\Pi}^q$		
Rank	5	4	4	$\beta_{ri} = 1$, $r, i = 1, \dots, 4$	*PPP*
	4	4	4		
	3	3	3	$\beta_{11} = -\beta_{12}$, $\beta_{21} = -\beta_{13}$, $\beta_{31} = -\beta_{14}$ where $\beta_{r1} = 1$, $r = 1, \dots, 3$	*GPPP* and *PPP*
	2	2	2	$\beta_{11} = -\beta_{1j}$ and $\beta_{21} = -\sum_{i=2, i\neq j}^{4} \beta_{2j}$ where $\beta_{r1} = 1$, $r = 1, 2$	*GPPP* and *PPP*
	1	1	1	$\beta_1 = -\sum_{i=2}^{4} \beta_i$ where $\beta_1 = 1$	*GPPP* and *PPP*
	0	0	0		*PPP* rejection

Note. In the table, we associate to each possible rank of the matrices $\mathbf{\Pi}^y$ and $\mathbf{\Pi}^q$ a relevant economic hypothesis and the restrictions that can be tested. In particular, if the matrixes have full rank (first two rows), we obtain a system of four stationary variables. Conversely, if the rank is equal to zero (last row), each variable can be considered a unit root process. All other possible ranks of the matrices can be considered as a validation of the *GPPP* hypothesis, and in the most restricted case, that is when also the symmetric and proportional conditions hold in each vector of the system, as a validation of the standard *PPP* hypothesis.

Table 2 – Unit Root Tests

(a) ADF tests	*eu*	*ee*	*lv*	*lt*
Deterministic component	c	c,t	c	c,t
Test statistics	-1.33	-2.81	-2.29	-1.75
	Δeu	Δee	Δlv	Δlt
Deterministic part	-	c	-	c
Test statistics	-11,25	-10.24	-4.41	-5.07
(b) UR tests with unknown breaks	*eu*	*ee*	*lv*	*lt*
Deterministic component	c	c,t	c	c,t
Test statistics	-1.16	-1.09	-5.20	-0.81
Detected break	2000:12	1994:3	1993:11	1993:10

Note. (a) The statistics are the augmented Dickey–Fuller test statistics for the null hypothesis of a unit root process; *eu*, *ee*, *lv* and *lt* denote the log level of the bilateral real exchange rate *vis-à-vis* the US for the Euro area, Estonia, Latvia and Lithuania respectively. Δ is the first difference operator. The critical value at the 1% level of significance is –4.02 to two decimal places if there is a constant (*c)* and a trend (*t)* in the regression, –3.48 if there is only a constant (*c*), and -2.58 if no deterministics components are included in the regression, while at the 5% level of significance these values are -3.44 ,-2.88 and -1.94, respectively (MacKinnon, 1996). (b) The number of lags in each regression is chosen according to the AIC criterion. The critical values at the 1% ,5% and 10% level of significance are equal to -3.55, -3.03 and -2.76, respectively (Lanne *et al.* 2002) when there is a constant (*c*) and trend in a regression, while are respectively equal to -3.48, -2.88 and -2.58, if only a constant term is included. The last row shows the detected break modelled as a shift dummy.

Table 3 – Misspecification tests

(a) Univariate misspecification tests

	EU	EE	LV	LT
$AR_{(1-7)}$	1.5911 [0.1463]	1.1532 [0.3363]	0.5849 [0.7668]	0.3374 [0.9351]
Normality	3.4745 [0.1760]	4.0960 [0.1290]	1.1350 [0.5669]	16.108 [0.0003]
$ARCH_{(1-7)}$	0.9031 [0.5075]	1.1315 [0.3500]	1.5206 [0.1694]	5.1289 [0.0001]
Heteroscedasticity	0.5027 [0.9936]	0.6682 [0.9278]	0.8185 [0.7687]	0.6991 [0.9027]
(b) Multivariate misspecification tests				
$AR_{(1-7)}$	1.1765 [0.1397]			
Normality	13.024 [0.1110]			
Heteroscedasticity	0.5572 [1.0000]			

Note: p-values in square brackets

Table 4 – Long-run properties

(a) Cointegration rank

			Trace test		Maximum eigenvalue test	
p-r	r	Eigenvalue	Statistics	95% cv	Statistics	95% cv
4	0	0.3253	107.616	47.21	57.84	27.07
3	1	0.1933	49.772	15.41	31.58	20.97
2	2	0.0957	18.196	15.408	14.81	14.07
1	3	0.0227	3.382	3.76	3.38	3.76

(b)Test of exclusion

r	dgf	5% .v.	EU		EE		LV		LT	
3	3	7.815	33.486	(0.000)	28.980	(0.000)	19.257	(0.000)	17.970	(0.000)

(c)Test of stationarity

r	dgf	5%c.v.	EU		EE		LV		LT	
3	1	3.841	9.848	(0.002)	10.090	(0.001)	8.656	(0.003)	11.420	(0.001)

(d) Test of weak exogeneity

r	dgf	5%c.v.	EU		EE		LV		LT	
3	3	7.815	10.545	(0.014)	11.861	(0.008)	22.323	(0.000)	17.592	(0.000)

Note:(a) The critical values for trace test and maximun eigenvalue statistics are from Pesaran et al. (2000); (b) p-value in round brackets.

Table 5 – VEC model estimated by 3SLS

	Δeu		Δee		Δlv		Δlt	
Intercept	.		-0.004	(0.002)	.		.	
$\varepsilon_{1,t-1}$	.		0.081	(0.023)	0.060	(0.010)	.	
$\varepsilon_{2,t-1}$	.		.		0.029	(0.007)	.	
$\varepsilon_{3,t-1}$	.		-0.023	(0.013)	.		0.052	(0.009)
Δeu_{t-1}	.		.		0.060	(0.029)	.	
Δee_{t-1}	.		.		.		.	
Δlv_{t-1}	.		.		.		.	
Δlt_{t-1}	.		.		.		0.092	(0.056)
Δeu_{t-2}	0.582	(0.196)	0.677	(0.208)	0.331	(0.109)	.	
Δee_{t-2}	-0.578	(0.186)	-0.635	(0.190)	-0.281	(0.100)	0.136	(0.064)
Δlv_{t-2}	0.146	(0.078)	.		.		-0.223	(0.113)
Δlt_{t-2}	.		0.072	(0.049)	.		0.113	(0.065)
Δeu_{t-3}	.		-0.089	(0.075)	.		.	
Δee_{t-3}	.		.		.		.	
Δlv_{t-3}	.		0.119	(0.075)	.		.	
Δlt_{t-3}	.		.		.		-0.106	(0.056)
Δeu_{t-4}	-0.375	(0.145)	-0.471	(0.152)	.		.	
Δee_{t-4}	0.230	(0.129)	0.318	(0.136)	.		.	
Δlv_{t-4}	.		.		-0.144	(0.052)	.	
Δlt_{t-4}	0.126	(0.065)	0.149	(0.070)	.		.	
Δeu_{t-5}	-0.354	(0.148)	-0.384	(0.153)	.		0.117	(0.058)
Δee_{t-5}	0.348	(0.131)	0.362	(0.136)	.		.	
Δlv_{t-5}	.		.		.		-0.324	(0.094)
Δlt_{t-5}	.		0.089	(0.042)	0.091	(0.037)	0.192	(0.053)
Δeu_{t-6}	.		.		0.230	(0.085)	0.409	(0.130)
Δee_{t-6}	.		.		-0.213	(0.079)	-0.317	(0.119)
Δlv_{t-6}	0.308	(0.106)	0.314	(0.106)	.		.	
Δlt_{t-6}	-0.179	(0.063)	-0.300	(0.065)	.		.	
Δeu_{t-7}	0.187	(0.094)	0.296	(0.094)	0.234	(0.092)	0.327	(0.136)
Δee_{t-7}	.		.		-0.147	(0.075)	-0.463	(0.117)
Δlv_{t-7}	-0.336	(0.128)	-0.551	(0.130)	-0.237	(0.070)	0.260	(0.089)
Δlt_{t-7}	0.108	(0.064)	0.111	(0.067)	.		.	
Δeu_{t-8}	.		0.085	(0.040)	0.212	(0.083)	.	
Δee_{t-8}	.		.		-0.232	(0.073)	.	
Δlv_{t-8}	.		-0.110	(0.062)	.		.	
Δlt_{t-8}	.		0.064	(0.040)	0.082	(0.034)	.	

$\chi^2(80) = 73.9399$
$\chi^2(80)$ 10% cv = 96.58; 5% cv = 101.88
Log-likelihood = 1.768842e+03

Notes. Standard errors in round brackets.

Table 6 –Forecast error variance decompositions

	Δee	Δlv	Δlt	Mean
Global shock	97.60	88.40	74.65	86.88
Regional shock	2.42	11.50	25.25	13.06
Idiosyncratic shock	0.72	0.78	0.91	

Note. The percentage of the variance of each variable of the system explained by global, regional and idiosyncratic shocks, where the latter are expressed as a percentage of regional disturbances. The simulation horizon is 60 months.

Table 7 –Robustness analysis: misspecification tests (quarterly data)

(a) Univariate misspecification tests

	EU	EE	LV	LT
$AR_{(1-7)}$	0.9785 [0.4330]	1.0234 [0.4102]	0.4191 [0.7936]	0.6677 [0.6191]
Normality	4.2437 [0.1198]	4.3969 [0.1110]	5.8442 [0.0544]	16.178 [0.0003]
$ARCH_{(1-7)}$	0.3136 [0.8664]	0.5372 [0.7096]	0.3763 [0.8235]	0.3112 [0.8681]
Heteroscedasticity	0.2438 [0.9981]	0.2306 [0.9987]	0.1939 [0.9996]	0.2104 [0.9993]
(b) Multivariate misspecification tests				
$AR_{(1-7)}$	1.7537 [0.0117]			
Normality	12.360 [0.1358]			
Heteroscedasticity	0.2488 [1.0000]			

Note: p-values in square brackets

Table 8 – Robustness analysis: long-run properties (quarterly data)

(a) Cointegration rank

			Trace test		Maximum eigenvalue test	
p-r	r	Eigenvalue	Statistics	95% cv	Statistics	95% cv
4	0	0.6642	90.07	47.21	53.47	27.07
3	1	0.3323	36.60	15.41	19.79	20.97
2	2	0.2508	16.81	15.408	14.15	14.07
1	3	0.0529	2.66	3.76	2.66	3.76

(b)Test of exclusion

r	dgf	5% .v.	EU		EE		LV		LT	
3	3	7.815	25.028	(0.000)	15.724	(0.001)	14.675	(0.002)	16.879	(0.001)

(c)Test of stationarity

r	dgf	5%c.v.	EU		EE		LV		LT	
3	1	3.841	48.340	(0.000)	16.366	(0.000)	15.437	(0.000)	15.960	(0.000)

(d) Test of weak exogeneity

r	dgf	5%c.v.	EU		EE		LV		LT	
3	3	7.815	10.770	(0.001)	11.328	(0.001)	10.153	(0.001)	11.432	(0.001)

Note:(a) The critical values for the trace test and maximun eigenvalue statistics are from Pesaran et al. (2000); (b) p-value in round brackets.

Table 9 – Robustness analysis: VEC model estimated by 3SLS (quarterly data)

	Δeu		Δee		Δlv		Δlt	
Intercept	.		-0.006	(0.006)	.		.	
$\varepsilon_{1,t-1}$	.		0.303	(0.064)	0.261	(0.038)	.	
$\varepsilon_{2,t-1}$	.		0.093	(0.061)	0.123	(0.029)	.	
$\varepsilon_{3,t-1}$	.		-0.186	(0.062)	-0.117	(0.041)	0.101	(0.018)
Δeu_{t-1}	.		-0.555	(0.215)	.		.	
Δee_{t-1}	.		0.513	(0.201)	.		.	
Δlv_{t-1}	.		.		.		.	
Δlt_{t-1}	.		.		.		.	
Δeu_{t-2}	.		0.247	(0.235)	0.478	(0.136)	0.194	(0.101)
Δee_{t-2}	.		-0.133	(0.211)	-0.306	(0.132)	.	
Δlv_{t-2}	0.121	(0.114)	.		-0.351	(0.091)	-0.425	(0.151)
Δlt_{t-2}	.		-0.166	(0.055)	.		0.260	(0.057)

$\chi^2(28) = 31.199$

$\chi^2(28)$ 10% cv = 37.92; 5% cv = 41.34

Log-likelihood = 4.817033e+02

*Notes. Standard errors in round brackets***.**

Table 10 – Forecasting accuracy analysis

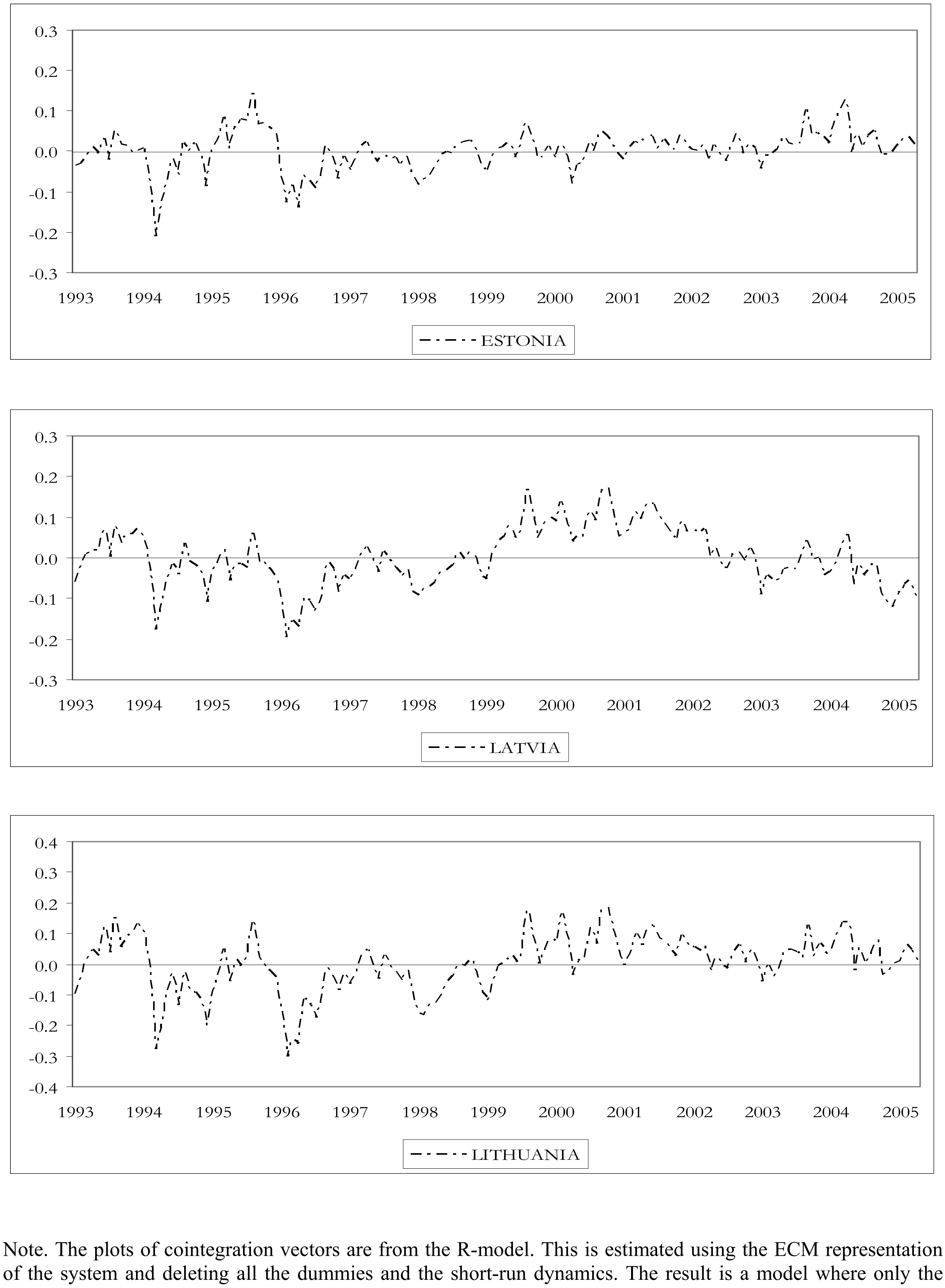
(a) Estonia-Euro	VECM (restr.)	VECM (unr.)		VAR (unr)		RW	
Mean-square error (MSE)							
1-month horizon	1.29E-04	1.158	(0.019)	0.637	(0.993)	0.650	(0.991)
6-month horizon	1.91E-03	1.173	(0.000)	0.707	(1.000)	0.754	(1.000)
12-month horizon	2.84E-03	0.891	(1.000)	0.764	(1.000)	0.792	(1.000)
24-month horizon	4.03E-03	0.854	(1.000)	0.715	(1.000)	0.747	(1.000)
Mean-absolute error (MAE)							
1-month horizon	9.57E-03	1.080	(0.032)	0.817	(0.993)	0.801	(0.997)
6-month horizon	3.97E-02	1.068	(0.000)	0.824	(1.000)	0.854	(1.000)
12-month horizon	4.82E-02	0.935	(1.000)	0.850	(1.000)	1.000	(1.000)
24-month horizon	5.55E-02	0.918	(1.000)	0.826	(1.000)	0.852	(1.000)

(b) Latvia-Euro	VECM (restr.)	VECM (unr.)		VAR (unr)		RW	
Mean-square error (MSE)							
1-month horizon	4.64E-04	0.774	(0.991)	0.731	(0.951)	0.795	(0.998)
6-month horizon	3.95E-03	0.775	(1.000)	7.848	(1.000)	0.714	(1.000)
12-month horizon	3.83E-03	0.715	(1.000)	1.086	(0.980)	0.935	(1.000)
24-month horizon	4.04E-03	0.943	(1.000)	1.530	(0.000)	1.302	(1.000)
Mean-absolute error (MAE)							
1-month horizon	1.74E-02	0.897	(0.953)	0.846	(0.964)	0.878	(0.996)
6-month horizon	5.46E-02	0.883	(1.000)	0.907	(1.000)	0.859	(1.000)
12-month horizon	5.32E-02	0.815	(1.000)	0.985	(1.000)	0.910	(1.000)
24-month horizon	5.27E-02	0.950	(1.000)	1.250	(0.168)	1.143	(1.000)

(c) Lithuania-Euro	VECM (restr.)	VECM (unr.)		VAR (unr)		RW	
Mean-square error (MSE)							
1-month horizon	1.68E-04	0.746	(0.933)	0.934	(0.279)	1.126	(0.949)
6-month horizon	8.70E-04	0.509	(1.000)	1.195	(0.190)	1.084	(0.068)
12-month horizon	6.63E-04	0.624	(1.000)	1.000	(0.021)	1.256	(0.000)
24-month horizon	1.24E-03	0.903	(1.000)	1.468	(0.001)	1.355	(0.000)
Mean-absolute error (MAE)							
1-month horizon	1.00E-02	0.901	(0.818)	1.044	(0.152)	1.125	(0.912)
6-month horizon	2.60E-02	0.737	(1.000)	1.102	(0.218)	1.044	(0.048)
12-month horizon	2.44E-02	0.779	(1.000)	1.114	(0.153)	1.061	(0.000)
24-month horizon	3.15E-02	0.923	(1.000)	1.142	(0.039)	1.088	(0.000)

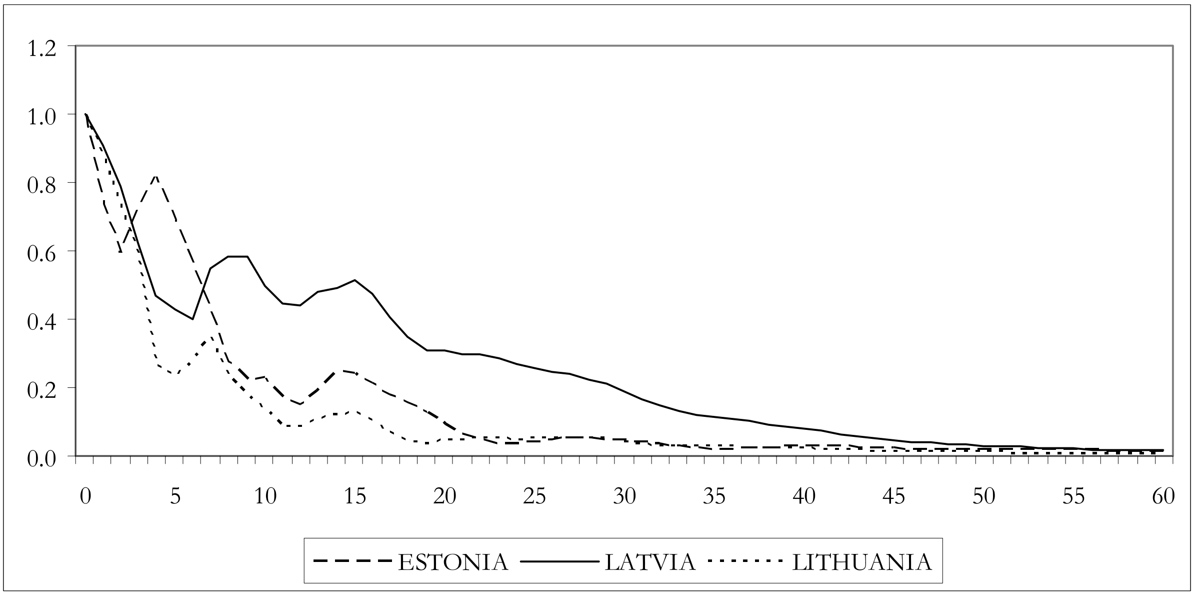
Note: The forecast period is 2006:1 to 2007:12. For the restricted VECM the MSE and MAE are expressed in levels, while for the alternative models they are expressed as the ratio between the criterion value of each model to that of the VECM. Thus, a value less than one indicates superior forecasting performance of the VECM. The *p-value* for the DM (Diebold-Mariano) test is reported in round brackets.

Figure 1. The cointegration vectors from the R-model



Note. The plots of cointegration vectors are from the R-model. This is estimated using the ECM representation of the system and deleting all the dummies and the short-run dynamics. The result is a model where only the long-run properties of the data are isolated.

Figure 2. Persistence profile of the cointegration vectors

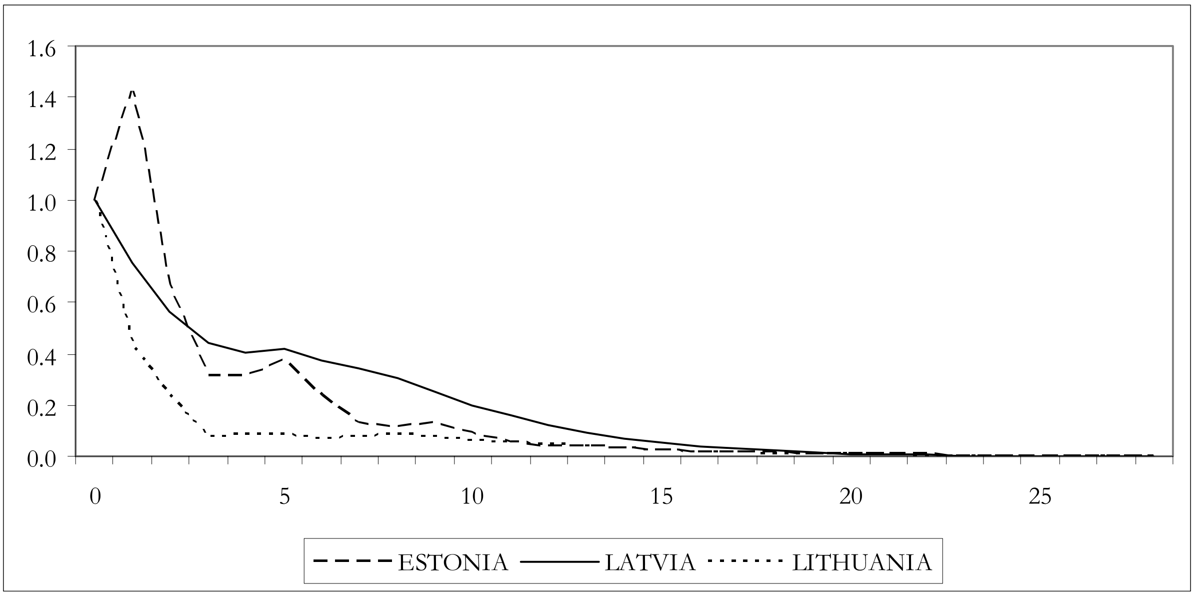


Note. The vertical axis indicates the magnitude of the deviation (normalised to unity on impact) from the steady-state level. The horizontal axis reports the number of months after the shock.

Figure 3. Robustness analysis: the cointegration vectors from the R-model (quarterly observations)

Note. The plots of cointegration vectors are from the R-model. This is estimated using the ECM representation of the system and deleting all the dummies and the short-run dynamics. The result is a model where only the long-run properties of the data are isolated.

Figure 4. Robustness analysis: persistence profile of the cointegration vectors (quarterly observations)



Note. The vertical axis indicates the magnitude of the deviation (normalized to unity on impact) from the steady-state level. The horizontal axis reports the number of quarters after the shock.