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Real Exchange Rate Dynamics in Developing Countries: Three Empirical Essays

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SUMMARY

The research activities carried out to prepare this thesis are mainly based on the study of the real exchange rate dynamics in developing countries. Although several studies in this empirical literature focus on the developed countries, only few studies investigate the sources of real exchange rate movements in developing countries. Thus, in this research I analyze the main sources of disturbances in two groups of countries: transition economies of Central and Eastern Europe and Latin America countries.

The thesis consists of three articles. In the first Chapter, I study the sources of real exchange rate fluctuations for a sample of transition countries and, in particular, I use permanent-transitory decomposition with the aim to measure the influence real and nominal disturbances to real exchange rates and to compute the shock synchronization between each transition country and the Euro Area. The second Chapter focuses on macroeconomic interdependencies between the Euro area and three selected transition economies (the Baltic countries). I develop a modified version of Generalised Purchasing Power Parity theory to study the degree of real exchange rate convergence in these countries and then to establish whether they are ready to adopt the Euro. Finally, in the third Chapter, I propose an theoretical and empirical framework in order to analyze whether fiscal disturbances have a role in explaining real exchange rate fluctuations in Latin American countries.

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ON THE SOURCES OF THE REAL EXCHANGE FLUCTUACTIONS: EVIDENCE FOR TRANSITION ECONOMIES

Abstract

The present study investigates the sources of real and nominal exchange rate fluctuations in selected transition countries. We adopt permanent-transitory decomposition and Kalman filter analysis in order to measure the "time-varying" correlations between shocks (real and nominal) to real exchange rates occurred in each transition country and in the Euro Area. The main empirical findings are that permanent (real) shocks are major driving force of real exchange rate fluctuations. Moreover, higher trade intensity goes along with higher nominal shock synchronization in the countries where the nominal shocks play a small role in explaining real exchange rate movements. On the other hand, when nominal shock has a (relative) bigger role, the relationship between economic integration and synchronization of shock is negligible.

Keywords: *Transition economies, Real exchange rate, Permanent transitory decomposition.*

1. Introduction

A large part of empirical literature has investigated the sources of disturbances both to real and nominal exchange rates. Usually, the methodology developed by Blanchard and Quah (1989) is used in order to identify and quantify the effects of several shocks on the real exchange rate movements. The majority of analysis in this literature focus on industrialised economies, while there are a few studies investigating developing countries.

Focusing on the decomposition to the real and nominal US Dollar exchange rate of a group of industrialized countries, Lastrapes (1992) finds that both the permanent real and nominal appreciation is mainly caused by real shocks, while nominal shocks cause only nominal appreciation. Evans and Lothian (1993) and Enders and Lee (1997), using two quite different samples of real exchange rates of industrialized countries, show that transitory disturbances explain only a small part of the real exchange rate movements. However, they also argue that there are circumstances in which temporary shocks may have a more substantial contribution and the magnitude of permanent-transitory decomposition may vary over time. Clarida and Gali (1994) identify three different type of shocks to the four major real US Dollar exchange rates estimating their contribution during the post Bretton-Woods period and finding that nominal-monetary shocks account for a great part of the variance of the real exchange rate.

As for developing countries, the evidence provided by this literature is quite clear in indicating that real shocks mainly drive the real exchange rates movements. Chen and Wu (1997), using a sample of four Pacific Rim countries, document that permanent shocks are the key determinant of the variability of real exchange rates. Moreover, Hoffmaister and Roldos (2001) provide evidence of the small role played by temporary shocks in explaining real exchange rate movements.

As argued by Lastrapes (1993), measuring the relative importance of permanent and transitory disturbances on real exchange rate is a very important issue for economic policy, especially when considering the transition economies. After the European Union (EU) enlargement of 2004, the Central and Eastern Europe (CEE) countries have indicated their willingness to join ERM II, sparking a debate for the optimal timing of entry.[1] The role of exchange rates, and of exchange rate regimes, in buffering exogenous shocks is a key issue in this debate. The economics of exchange rate suggests basically two different sources of real exchange rate fluctuations. The first one, as described by disequilibrium models (for instance,

[1] Some countries have already entered ERM II.

Dornbusch 1976), is mainly driven by nominal disturbances as monetary and financial market shocks. The second one is associated with the adjustment processes which follow shocks to real macroeconomic variables.

Thus as suggested by economic theory, transition countries may choose early entry in ERM II so as to shield their economies from a variety of asymmetric shocks and speculative attacks (Kontolemis, 2003). In fact, Borghijs and Kuijs (2004), Pelkman, Gros, and Ferrer (2000) find that flexible exchange rates in transition countries are poor buffers against external and nominal shocks. On the other hand, more exchange rate flexibility can be preferred when both transition process and macroeconomic convergence are slower.

Early attempts aimed at understanding real exchange rate movements in transition countries centred on decomposing real exchange rate changes into those real (or permanent), and nominal (or temporary) shocks. As noted by Dibooglu and Kutan (2002), this identification scheme can be used to evaluate the effectiveness of monetary and fiscal policy in transition economies. For example, if the temporary disturbances have a big role in explaining real exchange rate movements, it may mean that there is a high degree of nominal price inertia in the economic system, and then that nominal components of exchange rates can cause changes in the real exchange rates and therefore in the competitiveness levels of these economies.

Mundell (1961), Obstfeld (2002) have shown that the processes of shock absorption of exchange rates may vary with the different nature of disturbances which hit the economic system. For example, in presence of real asymmetric shocks it needs an adjustment in relative prices in order to prevent a increase of inflation and/or output losses. On the other hand, negative nominal shocks can cause an increase of the interest rate and also an appreciation of exchange rate.

The present study investigates the sources of real and nominal exchange rate fluctuations in selected transition countries. Following Lastrapes (1992) and Evans and Lothian (1993), we adopt a structural identification framework that allows us to distinguish between permanent (real) and temporary (nominal) disturbances and to compute their relative importance in explaining real exchange rate dynamics. Furthermore, we adopt Kalman filter analysis in order to measure the "time-varying" correlations between these shocks to real exchange rates occurred in each transition countries and in the Euro Area. Finally in a sub-sample of countries, we provide some new evidence of the positive correlation between shocks symmetry and trade integration.

The paper is organised as follows. In Section 2 the empirical framework and preliminary results of the permanent-transitory decomposition used are presented. Section 3 shows the measures of the synchronization of shocks between each transition countries and the Euro Area, while in Section 4 we report some robustness analysis based on a different identification scheme. Some final remarks follow.

2. Permanent and temporary shocks to exchange rate

2.1 Econometric framework

The empirical strategy consists in the decomposition of the real and nominal exchange rate into shocks and responses to shocks. In particular, we apply the structural VAR method proposed by Blanchard and Quah (1989), in order to identify a sequence of permanent and temporary disturbances for a bi-variate VAR model that included real and nominal exchange rate.

We consider the following infinite-order moving average representation:

$$Z_t = \Phi(L)\varepsilon_t \tag{1}$$

$$\begin{bmatrix} \Delta r_t \\ \Delta s_t \end{bmatrix} = \sum_{i=0}^{\infty} L^i \begin{bmatrix} a_{11i} & a_{12i} \\ a_{21i} & a_{22i} \end{bmatrix} \begin{bmatrix} \varepsilon_{Tt} \\ \varepsilon_{Pt} \end{bmatrix} \tag{2}$$

where Δ is the first difference operator, r_t and s_t are the real and nominal exchange rate, respectively. The parameter ε_{Tt} (ε_{Pt}) denotes the transitory (permanent) disturbances occurred at time *t*.

The innovations ε_{Tt} and ε_{Pt} are assumed to be with zero mean and constant variance and independent between them. Furthermore, it is possible to assume that temporary shocks, ε_{Tt}, do not affect the real exchange rate in the long-run. This assumption can be implemented in equation (1) defining the cumulated effect of ε_{Tt} shock on the changes in the real exchange rate series as:

$$\sum_{i=0}^{\infty} a_{11i} = 0 \quad \text{or} \quad \upsilon'\Phi(1)\upsilon = 0 \tag{3}$$

where $\upsilon = (1,0)'$. This implies that the sum of effects of temporary shocks on the real exchange rate must equal to zero.

If Δr_t and Δs_t are stationary and the level of real and nominal exchange rate are not co-integrated, there is an autoregressive representation of the form:

$$\Phi^{-1}(L)Z_t = \varepsilon_t \tag{4}$$

Now it is possible to estimate a finite order VAR for system (1) that produces a vector of innovations ε_t with a matrix of variance–covariance equals to Σ:

$$e_t = [1 - \Theta(L)]Z_t \tag{5}$$

where $\Theta(L)$ is a finite-order matrix polynomial in the lag operator. Now VAR residuals can be expressed as linear combinations of the temporary and permanent shocks:

$$\varepsilon_t = Ce_t \tag{6}$$

where C is a 2x2 matrix. Since in general we do not know all four elements in matrix C in order to identify ε_t from the VAR residuals e_t we need four restrictions: three are obtained by normalizing the pure innovations of the above moving average representation to unity and assuming orthogonality. These restrictions can be written as follows:

$$CC' = \Sigma \tag{7}$$

Comparing Equations 3 and 4 and using Equation 5 results in:

$$D = \Phi(1) \tag{8}$$

where $D = \left[1 - \Theta^{-1}(1)\right]C^{-1}$.

The final restriction can be derived using Equations 3 and 8:

$$\upsilon' D \upsilon = 0 \tag{9}$$

Using these restrictions, we now able to identify the permanent and temporary disturbances from equation 6, and compute their importance in explaining the nominal and real exchange rate dynamics.

2.2 Empirical results

Data

Monthly data for real and nominal exchange rates of seven transition countries (Albania, Bulgaria, Czech Republic, Hungary, Poland, Romania and Slovak Republic) vis-à-vis the Euro are used to estimate model (1) over the period 1993:1-2007:12. Both nominal exchange rates and price data are taken from the IMF's International Financial Statistics (IFS), while those for the Euro area from the Eurostat Newcronos database. The real exchange rate is defined as the product of the nominal exchange rate (national currency per Euro) and the ratio between Euro 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).

[**TABLE 1**]

Table 1 summarizes average and standard deviation of both the real and nominal exchange rates for each transition country vis-à-vis the Euro Area. For Albania, Bulgaria, Hungary and Romania, the volatility of real exchange rate is greater than the volatility of nominal exchange rate. For the other countries, Czech Republic, Poland and Slovakia, the former is smaller than the latter. Real exchange rates series follow a real appreciation pattern even if their variability seem to be different across countries, as Figure 1 shows. Heterogeneity in variability could reflect different transition processes and in particular different strategies with respect to the exchange rate regimes. As stressed by Kocenda (2001a), the choice of a particular exchange rate regime is one of the major policy decisions that transition countries had to make. From 1992 to 1999, all these economies defined their regimes. In particular, The Czech Republic, Slovakia, Albania and Romania favoured the managed float regime (semi-fixed), Hungary and Poland moved to a pre-announced crawling band, and finally only Bulgaria adopted a currency board regime.

[**FIGURE 1**]

Preliminary analysis

As a preliminary step, ADF (Dickey and Fuller, 1979) and Philips and Perron (1998) tests are performed on both real and nominal exchange rates series (Table 2). The deterministic component includes an intercept and, when statistically significant, a linear trend.[2]

[**TABLE 2**]

Consistently with many studies of the PPP hypothesis in transition countries (see, among others, Sarno and Taylor 2001), we are unable to reject the null of unit root at conventional nominal significance levels for both tests for all countries we consider. Given the evidence of $I(1)$-ness for all individual real and nominal exchange rate series and in order to apply the identification strategy described above, now we need to test for the presence to cointegration among them.

Estimation Results

We use the general-to-specific procedure in order to specify the seven bivariate models, starting to a maximum order of 12. The results, shown in the second column of Table 3, suggest 12 lags for Albania, 10 for Bulgaria, 9 for Hungary, 7 for both Czech Republic and Poland, 4 for Romania and finally 2 for Slovak Republic. In the rest of Table 3 we report both the main univariate and multivariate diagnostic tests.

[2] The number of lags is chosen such that no residual autocorrelation is evident in the auxiliary regressions.

There is no evidence of serial autocorrelation in the distribution of residuals as well as problem of heteroscedasticity, both at the single equation and the system level. Some departures from normality are instead detected, even if Gonzalo (1994) and Rahbek et al. (2002) have shown that the estimates of a VAR model are usually robust to these forms of misspecification.

[**TABLE 3**]

The long-term component of the model is identified on the basis of the trace and the maximum eigenvalue test (Johansen, 1995). In most of the cases both tests statistics suggest the presence of no cointegration in the models at the 5 percent significance level. Only in one model, Czech Republic, the maximum eigenvalue test indicates the presence of one cointegration vector at the 5% level of significance, while the trace statistics suggests no cointegration.

It is possible to conclude that in all models the real and nominal exchange rate are not cointegrated, accordingly we can use the identification strategy discussed above by estimating the bivariate models (2) in first differences.

[**TABLE 4**]

Forecast error variance decomposition

As discussed in Section 2.1, our econometric framework allows us to decompose the real and nominal exchange rates into a sequence of permanent and temporary shocks. The interpretation of these disturbances is crucial in order to investigate the sources of real exchange rate movements.

Usually, it is possible to distinguish between nominal and real disturbances. Nominal shocks, such as unanticipated changes in the money supply, often can lead to a nominal depreciation and the effects on the real exchange rate can be considered only temporary; in fact as stressed by Rogoff, (1996), as long as initial nominal rigidities dissipate over time real exchange rate will only be influenced in the short-run (Rogoff, 1996).[3] On the other hand, real shocks are often identified as technology shocks. The so-called Harrod–Balassa–Samuelson (HBS) effect (Harrod, 1933; Balassa, 1964; Samuelson, 1964) represents the most important theoretical explanation for the long run potential movements of real exchange rates in the long run. The basic idea is that real appreciation is caused by the relative difference between productivity growth in tradable and nontradables sectors. In particular, we can have a real appreciation as long as the former exceeds the latter. Agenor (1998) and Chinn (2000) have

[3] As shown by Chari et al. (2002) if the presence of nominal rigidities is particularly strong, the real exchange rate can be affected permanently by monetary shocks.

shown that in different emerging market economies HBS effects can be favoured by strong structural reforms and market liberalization, as those observed during the transition processes of CEE countries.

Given the above discussion and the way the literature has treated temporary and permanent disturbances, our study interprets temporary shocks as primarily nominal disturbances (e.g. money supply) and permanent shocks as primarily real disturbances (e.g. technology).[4]

In order to provide information on the relative importance of each shock on the movements of real exchange rates in transition countries, we carry out forecast error variance decomposition (FEVD). Table 5 shows the percentage of the variance of each real exchange rate explained by permanent-real and temporary-nominal shocks over the simulation period (60 months).

[**TABLE 5**]

As can be seen, permanent (real) shocks are found to mainly cause real exchange rate movements in most of the countries (with the exception of Bulgaria), even if their relative importance varies across entities of reference[5]. It is possible to identify three different groups of countries: *i*) in Slovak Republic, Poland and Hungary permanent shocks explain more than 75% of the real exchange rate variability; *ii*) in Romania, Czech Republic and Albania this kind of shock explains more than 55%; *iii*) for the case of Bulgaria, the average contribution of real disturbances is weaker and nominal shocks are the main driving force of real exchange rate variability. These findings are broadly consistent with the empirical evidence for transition economies shown by Coricelli and Jazbec (2004), Dibooglu and Kutan (2001) and Kontolemis and Rose (2005), as well as those detected for industrialized countries by Lastrapes (1992) and Evans and Lothian (1993).

Cross-country heterogeneity in the relative weights of real and nominal shocks on real exchange rates may reflect different transition processes of these economies. In particular, the first two groups of countries (with the exception of Albania) reflect the institutional aspects of transition reforms with respect to the international trade arrangement among the CEE countries. Such an arrangement was institutionalized in March 1993 in the form of the Central

[4] This interpretation is consistent with previous studies where the validity of long-run money neutrality is assumed (Lastrapes, 1992; Chen and Wu, 1997; Enders and Lee, 1997; Dibooglu and Kutan, 2001).

[5] In table 5, standard errors obtained by using bootstrapping methods indicate that all the FEVD are significantly different from zero.

European Free Trade Agreement (CEFTA).[6] As stressed by Kocenda (2001b), these countries display similar and relatively high degrees of convergence in nominal variables (e.g. monetary aggregates and price levels), while more heterogeneous behaviours in term of real variables (e.g. labour productivity and unemployment dynamics). The CEFTA framework has indeed allowed them to develop a coordinated policy framework especially in terms of monetary reforms and prospective policies needed to satisfy a set of EU pre-accession criteria. On the other hand, the weight of productivity shocks in explaining long-run departures from the equilibrium relationships, expressed by the PPP condition, can reflect different starting conditions at the beginning of transition process, and in particular different privatization schemes implemented. Moreover, as argued by Babetskii et al. (2004), if the potential catch-up in term of GDP per capita between transition economies and EU is very high (as in the case of these countries), real disturbances to exchange rates mainly consist in productivity shocks in transition countries. Thus, the importance of real shocks on the real exchange rate movements remains particularly strong as long as GDP per capita and productivity levels diverge.

3. Measuring the synchronization of shocks

Having found that real disturbances dominate over nominal shock in explaining real exchange rate volatility in transition economies, this section is devoted to assess the degree of asymmetry of shocks between CEE economies and the Euro Area over time. In fact, according to classical OCA criteria higher symmetry of shocks (in particular real ones) between countries implies a lower cost to sharing a common monetary policy. In the following paragraphs, we try to provide a measure of this synchronization in order to find some new evidence of the actual situation of the accession countries and the other transition economies with respect the potential decision to join the Euro Area.

A number of studies focus on measuring the degree of shock asymmetry across economies, especially between European Union members and accession countries.[7] The Blanchard and Quah (1989) framework is used in order to distinguish the nature of the shocks (for instance, supply and demand or permanent and temporary) and then to separate these from responses. Following this way it is possible to compute correlation between two countries' series of shocks and interpret these as measure of acceptability of a optimal

[6] The original CEFTA group was: the Czech Republic, Slovakia, Hungary, Poland, and Slovenia. Romania joined the CEFTA only in 1996.
[7] See, among others, Frenkel, Nickel and Schmidt (1999).

currency area. For instance, Bayoumi and Eichengreen (1996) identify the 'core' European countries, for which the cost of a common monetary policy appears to be low, because of the high symmetry of shocks.

However, since the correlation is a static index, it is difficult to verify whether or not shocks become more symmetric or not. Furthermore, since the degree of economic integration changes over time, it is not reasonably to imagine that shock asymmetry remains constant.

A useful way to solve this problem is to use the Kalman filter in order to compute time-varying estimates of shock synchronization. Adopting a similar approach, Babetskii et al. (2004) measure time-varying correlations of demand and supply shocks between Euro Area and accessing countries and they find that the first ones tend to converge over time, while the second ones diverge. The different behaviour of supply shocks is interpreted by the authors as a signal of a transition process manly based on mechanisms due to HBS effects.

In order to compute the magnitude of symmetry between shocks from each transition country and the Euro Area, we need to introduce two new disturbances (permanent and temporary). To do this, we apply the same methodology described as before to the real and nominal exchange rate of Euro *vis-à-vis* the US Dollar.

Blanchard and Quah (1989) decomposition of exchange rates allows to identify synthetic measures of what drives the dynamics of exchange rates. As in Lastrapes (1992) these can be interpreted as real and nominal disturbances only with respect to the dynamics of real and nominal exchange rates. Thus, the interpretation of shock synchronization measures cannot be directly comparable to which obtained in where the driving forces of business cycles are identified in a closed economy set-up.

In order to link the results of permanent-transitory decomposition in Section 2 to the OCA theory, we need to better clarify the relationships between structural innovation obtained in our model and the more general business cycle shocks usually used in this analysis. Accordingly, we adapt the analysis of Bayoumi and Eichengreen (1996) using a general open-economy framework in the spirit of traditional international business cycle models (Ahemd et al. 1993 and Alexius 2005, among others). In particular, we consider a bi-variate VAR for output and price expressed in relative terms with respect to the foreign economy.[8,9] Then we apply the Blanchard and Quah (1989) decomposition imposing that supply shocks can be

[8] Monthly industrial production series are used as proxies for output. Since monthly data are not available for Albania and Bulgaria, only for these countries quarterly observations of GDP have been interpolated using Di Fonzo's (1990) method to obtain monthly series. Data for the period 2003:1-2007:12 are taken from Datastream.
[9] In practice, we take the (log) difference of the two variables (output and price) between each transition countries and the Euro Area.

have permanent effects on output while both supply and demand disturbances having permanent and temporary effects on the level of prices. In order to verify how much our measures are related to those shown in Bayoumi and Eichengreen (1996) framework, we next compute simple correlations between these shocks and those identified previously using the decomposition of exchange rates. We focus on the correlation between the two measures of nominal shocks which will be particularly relevant in the rest of the paper where we provide evidence of the links between the symmetry of nominal disturbances and the international trade intensities of transition countries and the EU economy (table 6).

[TABLE 6]

In most of the cases there is a high, positive and significant correlation between the two different specifications of nominal shocks. This is particularly evident in the case of EU, where we find a correlation index larger than 0.8. Only in the models for Albania and Bulgaria we detect a weak correlation between shocks, even if with the expected sign[10]. These findings allows us to interpret nominal shocks obtained by the permanent-transitory exchange rate decomposition and their respective synchronization measures as disturbances also related to more general dynamics of business cycle movements.[11]

3.1 Methodology

Following Boone (1997), we use the Kalman filter to compute the 'time-varying correlation coefficient' between transitory (and permanent) shocks to real exchange rate occurred in countries *i* and *EU* (Euro Area) given by b_t^{iEU} :

$$(\varepsilon_{St}^{i}) = a_t^{iEU} + b_t^{iEU}(\varepsilon_{St}^{EU}) + \mu_t \tag{10}$$

$$a_t^{iEU} = a_{t-1}^{iEU} + \nu_t^a \text{ } and \text{ } b_t^{iEU} = b_{t-1}^{iEU} + \nu_t^b \tag{11}$$

where ε_{St}^{i} are the shocks (with $S = transitory, permanet$), error terms μ_t and ν_t are white noise disturbances, *i* indexes a transition country, EU stands for Euro Area. Coefficients a_t^{iEU} and b_t^{iEU} are allowed to vary over time according to (11), which are called transition or state equations, while equation (10) is called the measurement or observation equation[12].

[10] These results can be influenced by the quality of the data for these two countries (see footnote 8).

[11] As pointed out by the referee, it is possible to interpret structural innovations to nominal exchange rates only as a part of business cycle innovations. In a theoretical framework based on the monetary approach to exchange rates determination, they represent indeed the monetary part of the more general business cycle shocks.

[12] Technically, the Kalman filter represents a recursive algorithm for computing the optimal estimator of unknown parameters. This is done by maximizing a likelihood function given the information available at time t. The estimator is optimal in the sense that it minimizes the mean square error (MSE). Furthermore, if all

The intuition behind this specification is simple. For example, in the presence of positive correlation of shocks between countries *i* and *EU*, coefficients b_t^{iEU} will be positive. The main advantage of the method in hand is thus to give optimal estimations of the time-varying coefficients in the presence of structural changes, which are likely to occur in the case of transition countries.

3.2 Empirical Results

Before presenting the results from Kalman filter estimates, we report moving window correlations between both transitory and permanent shocks in each transition country and the Euro Area (figures 2 and 3).[13] Regarding to transitory shocks, we find in the case of Albania a positive correlation up to 1999. After this year the sign of the correlation is not uniquely defined and its magnitude is often small. A similar patterns emerges for Bulgaria. For the case of the Czech Republic, the correlation is substantially negative even though it presents long-lasting swings. A positive correlation is detected in Romania, even if there are periods in which it is the opposite sign, while in Poland and Slovak Republic a negative correlation between nominal shocks seems to appear. Finally, only in the case of Hungary we find a positive index for the entire sample span.

[FIGURE 2]

As for permanent shocks, we do not find a clear evidence of positive correlation between shocks in each transition countries and the Euro Area. In most of the cases the index is negative, as in the case of Poland and Slovak Republic. Only in Bulgaria we detect a strong positive correlation between shocks, albeit declining in the most recent years.

[FIGURE 3]

Moving window correlations shown in figures 2 and 3 point out heterogeneous patterns between countries and shocks, so that the Kalman filter approach seems to be a useful tool in order to better analyze the degree of shocks symmetry in transition economies.

We estimate the coefficients of equation (10) after computed the permanent-transitory decomposition for the real exchange rate of the Euro Area with respect the US Dollar[14].

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disturbances are normal, the Kalman filter provides the maximum likelihood estimator (MLE) of coefficients in (10).

[13] Both correlations are computed on a moving-window of 36 monthly observations starting from 1994:1.

[14] The specification of the bi-variate VAR model for the real and nominal exchange rate of the Euro *vis-à-vis* the US Dollar followed the usual steps. Unit root tests statistics (ADF and PP) indicate the series are integrated of order 1. The general-to-specific procedure suggests an autoregression order of two and an unrestricted constant in the system. Both trace test and Max-eigenvalue test indicate no cointegration at the 0.05 level.

Figure 4 shows the estimated coefficients b_t^{iEU} when transitory shocks are included in equation (10), while Figure 5 reports the coefficients of the estimates with respect to permanent disturbances. Both Figures cover the period 1994:1-2007:12.

[**FIGURE 4**]

In the case of Albania, even if both estimated coefficients are positive, they show a decreasing pattern over time. The higher synchronization of both nominal and real shocks between Albania and Euro Area in the first years seems to disappear during the last period. In Bulgaria the symmetry of nominal shocks seems to strongly increase up to 1997; after then the estimated coefficient is slowly decreasing even if its value remains positive. This behaviour could reflect the process of economic and political stabilization caused by several reforms introduced in 1997, like the adoption of a currency board[15]. The synchronization of real shocks follows a similar path. In Hungary, a stable and slow growth of symmetry in nominal shocks starts from the end of 1995 and it coincides with the introduction of a crawling peg in order to sustain increased competitiveness (achieved by the significant initial adjustment of the exchange rate) and hence to calm speculation against the national currency (Kocenda, 2001a). The efforts to defend international competitiveness seem to have also favoured the symmetry in real shocks which has increased steadily from 1996. The Poland presents negative estimated coefficients for both nominal and real shocks.

The huge jump shown by the series of b_t^{iEU} in Romania with respect to temporary shocks can be partly explained by the important reforms introduced quickly between the end of 1996 and the first half of 1997. In particular, the definitive liberalization of the foreign exchange markets led to a dramatic nominal depreciation of the currency in the first months of 1997. After this period, the stabilization process seems to have not contributed to a higher symmetry in both nominal and real shocks mainly because of a persistent inflationary pressure. In the Slovak Republic case, negative coefficients are detected.[16] Nevertheless, the symmetry of nominal shocks seems to be slowly increasing after 1996. The real shock synchronization is decreasing up to the half of 2002, and after that it starts to show a positive trend. Finally, the Czech Republic presents the most complex series to interpret as both coefficients are estimated with negative signs.

In conclusion, with respect to the symmetry in nominal disturbances for most of the countries (with the exception of Albania and Czech Republic), we find an increasing or stable

[15] In the first period the currency was linked to DEM, while form 1999 it was fixed with respect to Euro.
[16] Slovakia was the first of this group of transition countries to join the Euro Area in the 2009.

dynamic after a first period of greater variability. This evidence is consistent with the results in Kocenda (2001b), according to which the transition processes of these economies can be distinguish into two different phases: the first stage is characterized by the main political and economic reforms (such as market liberalization, privatizations and the choice of the exchange rate regimes);[17] the second stage is the period in which the policy authorities have tried to manage the reform processes in order to achieve stability and credibility for the domestic economic system as well as to lower and stable inflation.

[**FIGURE 5**]

The patterns of the synchronization measures of the shocks seem to be more heterogeneous. According with the evidence shown in Babetskii et al. (2004), the presence of shock asymmetries (negative estimates of coefficient b_t^{iEU}) can be caused by the dominance of HBS effects for these economies, which during the transition process, have been displaying high productivity growth in the traded goods sector (see, among others, Coricelli and Jazbec 2004)[18]. Where these effects are not dominant, we find positive relationships of the real disturbances of each transition countries and the Euro Area. As argued by De Broeck e Sløk (2001), this can occur when the competitiveness levels of a transition economy are not eroded by a real exchange rate appreciation caused by productivity gains in the tradable sector because of a initial undervaluation of the national currencies and of the process of controlling inflation.

3.3. Extensions: trade integration and shock correlation

With their paper in 1998, Frankel and Rose have opened an huge economic debate regarding the *endogeneity* of OCA. Following the basic idea of the European Commission (1990), they argue that the presence of close trade links between countries can lead to a process of business cycle synchronization, through this channel, to an increase of the symmetry of disturbances (European Commission-view)[19]. The main alternative viewpoint it

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[17] This group reflects in part the institutional aspects of transition reforms with respect to the international trade arrangement among the CEE countries In 1993 the Slovak Republic, Hungary, Poland (plus chez Republic and Slovenia) founded the Central European Free Trade Agreement (CEFTA). Romania joined the CEFTA only in 1996 while Bulgaria became a member in 1998.

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

[19] Frankel and Rose (1998) stress the necessity of further analysis of the role of international trade by distinguishing between inter-industry and intra-industry trade. Inter-industry trade (trade which involves exports and imports of different goods, for example, when one country exports cotton and imports wines) reflects specialization, thus potentially causing asymmetries. On the other hand, intra-industry trade (when a country simultaneously exports and imports products of the same category, e.g., cars) should lead to business cycle co-movements. There is on-going theoretical work in this direction.

is represented by Krugman (1993), who suggests that international trade increases specialization and therefore the presence of asymmetric shocks. Thus, theoretically there is not a general consensus on the overall impact of trade integration on shock symmetry. In our framework we can test if higher (trade) integration follows through to a higher synchronization of the nominal and real shocks that affected the international competitiveness of the transition countries.

In order to develop this analysis, we firstly define three measures of trade intensity following Frankel and Rose (1998) and Babetskii (2005):

$$T_{iEUt}^{IM} = IM_{iEUt}/(IM_{it} + IM_{EUt}) \qquad (12)$$

$$T_{iEUt}^{EX} = EX_{iEUt}/(EX_{it} + EX_{EUt}) \qquad (13)$$

$$T_{iEUt}^{T} = (EX_{iEUt} + IM_{iEUt})/(EX_{it} + EX_{EUt} + IM_{it} + IM_{EUt}) \qquad (14)$$

where IM_{iEUt} are imports of country *i* from country *EU* (Euro Area), while IM_{it} and IM_{EUt} are total imports of country *i* and *EU*, respectively. Analogously, EX_{iEUt} are the exports from country *i* to Euro Area, EX_{it} and EX_{EUt} are the total exports of the two countries.

Secondly, we estimate the correlation between the time-varying coefficients of transitory shocks synchronization in (10), b_t^{iEU}, and these measures of trade integration (or trade intensity):

$$\rho(i, EU) = corr(b_t^{iEU}, T_{iEUt}^{K}) \qquad (15)$$

$$K = IM, EX, T$$

where *i* is the transition country while *EU* is the Euro Area. A positive correlation indicates that higher trade intensity goes along with higher synchronization of shocks, while negative correlation suggests the opposite.

Finally, assuming that trade intensity is exogenous to our measure of shock synchronization,[20] we can measure the relationships between these variables using standard regression techniques that allow us to exploit the time-dimension of our data:

[20] As argued by Babetskii (2005) the direction of causality between shocks (supply and demand shocks in his case) and trade intensity is not uniquely determinate. However, since that all transition countries show in the last decades a robust increase in the trade integration with the Euro Area, it is possible to think that this trend it is caused mainly by structural factors rather than transitory shocks.

$$b_t^{iEU} = a_1 + a_2\, T^{T}_{iEUt} + \varepsilon_t \qquad (16)$$

Monthly data for imports and exports over the period 1994:1-2007:12 are from OECD Stat. The data availability forces us to perform the analysis only on a sub-sample of transition countries: Czech Republic, Hungary, Poland and Slovak Republic. Descriptive statistics of the three measures of trade intensities are reported in table 7.

[**TABLE 7**]

In all countries the trade intensities show an increasing behaviour over time, even if the changes are small in magnitude. Their variability remains constant in all the sub-period considered. The trade openness of Poland with respect the Euro Area seems to be stronger than those obtained for the other countries, while the Slovak Republic shown the smallest trade intensities

The estimate results of equation 16 are reported in Table 8.

[**TABLE 8**]

In the case of Poland, there is a strong a positive correlation between all the measures of trade integration and the index of synchronization of the temporary (nominal) shock (0.801, 0.710 and 788); this relationship is confirmed also by the significant and positive slope of the *OLS* regression. Similar considerations can be made both for Hungary and Slovak Republic, even if the intensity of the links between the variables are weaker. In the case of Czech Republic, the results are different. We find that the correlation between b_t^{iEU} and T^{IM}_{iEUt} (the measure of trade intensity based on the imports) is negative, while the correlation with the other two measures is positive. However, in all three cases the absolute values of the indexes are quite small. The slope in the regression is positive even if it is not statistically significant.

As for permanent (real) shocks, we find robust results in favour of positive correlations between these disturbances and trade intensity. The magnitude of these correlations varies across countries. However the countries (Slovak Republic and Poland) in which these are smaller are also the countries in which the permanent shocks has a key and determinant role in explaining real exchange rate dynamics, as shown in Table 4.

Previous estimates are replicated using also quarterly data. To do this, we average out the monthly measures of shocks synchronization in order to obtain quarterly observations. The results shown in table 9 confirm all previous findings.

[**TABLE 9**]

Overall, it is possible to argue that where the nominal shocks play a small role in explaining real exchange rate movements (Hungary, Poland and Slovak Republic, see Table

4), higher trade intensity goes along with higher synchronization of these shocks between each transition country and the Euro Area. On the other hand, when nominal shock has a (relative) bigger role in the determination of the real exchange rate movements then the relationship between economic integration and synchronization of shock is negligible, and in some case can also be negative.

As pointed out in the last part of section 2, the countries where the impact of nominal shocks on real exchange rate is weaker are also the countries where the CEFTA agreement had led to a bigger convergence in terms of nominal variables. The higher synchronization of nominal shocks detected for these countries reflects the efforts performed during the transition process in order to facilitate international and policy cooperation.

Then, the results shown in this section are mixed. For some countries (Hungary, Poland and Slovak Republic) the paradigm described by the so called "European Commission-view"[21] can be apply to the nominal shocks that hit the international competitiveness index of a transition economy. Nevertheless, in the presence of nominal shocks explaining a relatively huge part of real exchange rate dynamics, the economic integration does not link with an higher symmetry of shocks.

Finally, we provide an estimate of a single measure of sensitivity of shock synchronization to trade intensity for all transition countries for which we have data of international trade data.[22] Quarterly information are used in order to avoid possible problems due to the high volatility of monthly data for trade intensity measures. The results of the country fixed effects estimates are reported in table 10.

[TABLE 10]

We document a positive and significant impact of all measures of trade intensities on the synchronization indexes of both transitory (nominal) and permanent (real) shocks, with a magnitude that seems to be greater for the latter. These findings give further support to the idea that the trade integration increases the symmetry of shock.

4. Robustness analysis: a different identification strategy for real and nominal shocks to real exchange rate

When HBS effects are particularly relevant in explaining the dynamics of real exchange rates, as in the case of transition countries, the real-nominal decomposition of exchange rates may be unsatisfactory since with the obtained structural shocks we cannot distinguish the

[21] European Commission (1990).
[22] Namely, Czech Republic, Hungary, Poland and Slovak Republic.

influence of real disturbances in traded sector from the real disturbances occurred in the nontraded sector. Thus, we augment the baseline VAR framework detailed in Section 2.1 above with other real variables in an effort to identify other relevant shocks related to the case of transition economies. To do this, we need further hypothesis in order to account for the presence of a traded good sector and of a nontraded good sector[23]. Following the De Arcangelis (1993), we re-formulate the system (2) as:

$$\begin{bmatrix} \Delta y_t \\ \Delta r_t \\ \Delta s_t \end{bmatrix} = \sum_{i=0}^{\infty} L^i \begin{bmatrix} a_{11i} & a_{12i} & a_{23i} \\ a_{21i} & a_{22i} & a_{23i} \\ a_{31i} & a_{32i} & a_{33i} \end{bmatrix} \begin{bmatrix} \varepsilon_{Tt} \\ \varepsilon_{TRt} \\ \varepsilon_{NTRt} \end{bmatrix} \tag{17}$$

where y_t is the adjusted difference between the log of real output per worker in each transition country and the Euro Area, which is used as proxy of observable relative labour cost[24]. In system (17), we consider three different structural disturbances: ε_{Tt}, is the nominal (transitory) shock, while ε_{TRt} and ε_{NTRt} represent the real shocks in the traded and nontraded sectors, respectively [25]. The main difference with respect the baseline empirical model described previously is that now we can account for two different real shocks to real exchange rates. As in the De Arcangelis (1993), we assume that relative labour cost, as well as its proxy y_t, is only affected by the real shock in traded sector ε_{TRt}. This implies that the real shock in nontraded sector of the economy, ε_{NTRt}, does not affect long-run differences in labour costs, since the shares of labour of the different sectors in the different countries remain constant when there are productivity changes in the nontraded sector[26]. Furthermore, we assume that real exchange rates can be affected permanently by both real disturbances, but not by the nominal one[27]. These assumptions, together with the standard normalization procedure and the orthogonally conditions imposed on the pure innovations of the moving average representation of system, lead to a just-identify structure for the structural shocks in (17).

[23] In particular, we assume that technologies are different between the two countries (each transition economy and the Euro Area) and between the two sectors.

[24] The variable y_t is defined as: $y_t = x_t - x_{EU,t} + r_t$, where x_t and x_{EUt} are the (log of) output per worker in each transition country and Euro Area, respectively.

[25] In particular, we can interpret ε_{TRt} and ε_{NTRt} as the technology disturbances that affect permanently the productivity differential of the two countries in the production of traded and nontraded goods, respectively.

[26] This is possible when the productivity changes in nontraded sector are compensated both by a reduction in price and an increase in the demand of nontraded goods so that the amount of labour input in this sector remains unchanged. Thus real wages increase (in fact the nominal wages do not change while the prices are decreasing) and the real exchange rate needs to increase in order to compensate the gains in the competitiveness of nontraded sector. The final result is that the relative labour cost between countries remains constant.

[27] This is because, as in the baseline model discussed in section 2, the neutrality of money holds in the long-run.

Since monthly data for output per worker are not available for our transition economies, we used quarterly observations over the period 1994:1-2007:4[28]. Table 11 presents the results of FEVD analysis for the augmented specification[29].

[**TABLE 11**]

In all models the importance of nominal structural innovation is weaker than that detected in the baseline specification[30]. Thus, the influence of real shocks seems still to dominate real exchange rate variability for Czech Republic, Hungary, Poland, Romania and Slovak Republic. For these countries, within real shocks, those occurred in the traded sector have a bigger role in explaining real exchange rate dynamics than those occurred in nontraded sector. On the contrary, in the case of Bulgaria we find that nominal disturbance is the main driving force of real exchange rate movements (in a way consistent with the evidence of Table 5) in the short run, while real shocks in nontraded sector are dominant in the long run[31]. In summary, the main results seem to be in support of the presence HBS effects for these economies (Coricelli and Jazbec, 2004) and of the hypothesis that the influence of productivity innovations in traded sector can cause long run deviations of the PPP condition (Sarno and Taylor 2001)[32].

5. Conclusions

The present study has investigated the sources of real and nominal exchange rate fluctuations in selected transition countries, using a structural identification framework that allows us to distinguish between permanent (real) and temporary (nominal) disturbances. Forecast error variance decomposition analysis is implemented in order to compute the relative importance of these shocks in explaining real exchange rate variability.

Time-varying correlations between disturbances (real and nominal) occurred in each transition country and in Euro Area have been computed through Kalman filter techniques in order to find a measure of the shock symmetry. Then, we have used these measures to verify

[28] Data for output per worker are taken from OECD, while those for both nominal exchange rates and price indexes from IMF's IFS database

[29] The empirical strategy adopted follows the steps used in Section 2.2. Due to the lack of data availability, no evidence for the augmented model for Albania is presented. Unit root and cointegration tests are not reported to save space, but it is possible to conclude that in all the models the variables (difference in output per worker, real and nominal exchange rate) are not cointegrated. Complete estimation results are available on request.

[30] As in the baseline model, all the FEVD are significantly different from zero.

[31] This last result is consistent with the evidence reported in section 3.2 where we argued that positive relationships of the real disturbances can be detected where HBS effects are not dominant, as in the case of Bulgaria for which we find a smaller role played by real shocks in traded sector and positive correlations between real disturbances with respect to Euro Area (see Figure 3).

[32] It is important to note that PPP does not hold in our framework because we do not find stationary behavior for real exchange rates in all countries (see Table 2).

if higher trade integration leads to increase the synchronization of shocks as suggested by the so-called "European Commission-view".

Our empirical findings are broadly consistent with previous evidence on transition economies (Dibooglu and Kutan 2001, Kontolemis and Rose 2005): permanent (real) shocks are found to mainly cause real exchange rate movements in all countries, even if their relative importance varies across economies.

In most of transition countries analyzed, the time-varying correlations between nominal disturbances with respect to Euro Area seem to follow an increasing path over time, albeit in some of countries negative co-movements emerge. As in Babetskii et al. (2004), we find that the behaviours of the synchronization measures of the real shocks seem to be more heterogeneous. One possible explanation is the strong presence in some countries of HBS effects. The importance played by the HBS effects is confirmed when an augmented identification strategy is used to investigate the role of productivity shocks in traded and nontraded sectors.

Finally, we find that higher trade intensity goes along with higher synchronization of nominal disturbances in the countries where the nominal shocks play a small role in explaining real exchange rate movements. On the other hand, when nominal shocks have a (relative) bigger role the relationship between economic integration and synchronization of shock is negligible, and in some cases it is negative. Overall, when we consider a single measure of sensitivity of shock synchronization to trade intensity using an unique panel all transition countries, our findings give further support to the idea that the trade integration increases the symmetry of shocks.

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Tables and Figures

Table 1 – Descriptive statistics

	real exchange rate		*nominal exchange rate*	
	average	st. dev.	average	st. dev.
Albania	0.141	0.254	0.004	0.120
Bulgaria	0.095	0.284	-0.859	1.472
Czech Republic	0.009	0.183	-0.072	0.090
Hungary	-0.035	0.149	-0.191	0.274
Poland	0.101	0.156	-0.093	0.194
Romania	0.103	0.247	-0.480	1.208
Slovak Republic	-0.012	0.226	-0.077	0.074

Note Mean and standard deviation of both real and nominal exchange rates of each country *vis-à-vis* the Euro Area over the period 2003:1-2007:12. All the variables are expressed in logarithms and constant prices (base year 2000=1).

Table 2 – Unit Root Tests

		(a) ADF test			(b) PP test	
		Det. part	statistics		Det. part	statistics
Albania	r	c	-2.52	r	c	-2.57
	Δr	-	-11.42	Δr	-	-12.95
	s	c	-1.87	s	c	-1.92
	Δs	-	-13.11	Δs	-	-13.12
Bulgaria	r	c	-1.28	r	c	-2.00
	Δr	-	-9.91	Δr	-	-15.32
	s	c	-2.49	s	c	-2.41
	Δs	-	-8.81	Δs	-	-8.94
Czech Republic	r	c	-0.96	r	c	-1.18
	Δr	-	-15.55	Δr	-	-15.65
	s	c	-0.15	s	c	-0.18
	Δs	-	-16.09	Δs	-	-16.28
Hungary	r	c	-0.303	r	c	-0.246
	Δr	-	-13-033	Δr	-	-13.031
	s	c	-1.954	s	c	-1.963
	Δs	-	-11.985	Δs	-	-11.997
Poland	r	c	-1.385	r	c	-1.342
	Δr	-	-11.296	Δr	-	-11.158
	s	c,t	-2.358	s	c,t	-2.329
	Δs	c	-9.538	Δs	c	-11.011
Romania	r	c	-1.252	r	c	-1.893
	Δr	-	-7.309	Δr	-	-11.946
	s	c,t	-1.460	s	C,t	-1.690
	Δs	c	-8.707	Δs	c	-8.727
Slovak Republic	r	c	0.314	r	c	0.293
	Δr	-	-12.270	Δr	-	-12.317
	s	c	-0.728	s	c	-0.861
	Δs	-	-12.825	Δs	-	-12.850

Note. (a) Augmented Dickey–Fuller test statistics for the null of a unit root process; 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 statistics are Phillips-Perron test statistics for the null hypothesis of unit root. The critical values at the 1% ,5% and 10% level of significance are equal to -4.01, -3.44 and -3.14, respectively if a constant term and a trend are included in the regression. They are equal respectively to -3.47, -2.88 and -2.58 if there is only a constant. Finally, they are equal to -2.58, -1.94 and -162 if no deterministic part are in the regression (MacKinnon, 1996) .

Table 3 – Misspecification tests

	# lags		(a) Univariate misspecification tests			b) Multivariate misspecification tests		
			$AR_{(1-4)}$	Normality	Heterosk.	$AR_{(1-4)}$	Normality	Heterosk.
Albania	12	*r*	0.189 [0.947]	42.160 [0.000]	1.025 [0.449]	0.995 [0.462]	68.535 [0.000]	0.930 [0.685]
		s	0.437 [0.782]	51.307 [0.000]	1.320 [0.126]			
Bulgaria	10	*r*	0.720 [0.579]	140.330 [0.000]	1.842 [0.007]	1.530 [0.088]	566.870 [0.000]	0.939 [0.645]
		s	0.701 [0.592]	138.570 [0.000]	1.307 [0.152]			
Czech Republic	7	*r*	0.503 [0.737]	50.299 [0.000]	0.894 [0.621]	0.738 [0.754]	36.476 [0.000]	0.881 [0.803]
		s	0.299 [0.878]	59.415 [0.000]	0.911 [0.598]			
Hungary	9	*r*	1.436 [0.225]	26.436 [0.000]	0.944 [0.565]	1.233 [0.299]	36.671 [0.000]	1.225 [0.089]
		s	2.451 [0.050]	36.919 [0.000]	1.363 [0.111]			
Poland	7	*r*	0.723 [0.577]	10.661 [0.005]	0.769 [0.856]	0.958 [0.503]	5.867 [0.209]	1.240 [0.092]
		s	1.645 [0.166]	10.329 [0.006]	0.923 [0.633]			
Romania	4	*r*	0.807 [0.583]	227.620 [0.000]	1.089 [0.357]	1.249 [0.185]	137.070 [0.000]	6.864 [0.0000]
		s	0.578 [0.773]	98.171 [0.000]	1.514 [0.049]			
Slovak Republic	1	*r*	0.291 [0.884]	17.108 [0.000]	0.912 [0.458]	0.642 [0.849]	128.670 [0.000]	1.329 [0.199]
		s	0.435 [0.783]	12.781 [0.002]	1.437 [0.224]			

Note: p-values in square brackets

Table 4 – The trace and maximun eigenvalue test for cointegration

	Model	Rank	Eigenvalue	Trace test		Max eigenvalue test	
Albania	2	0	0.0506	16.774	(0.141)	8.727	(0.464)
		1	0.0468	8.047	(0.081)	8.047	(0.081)
Bulgaria	2	0	0.0615	17.897	(0.103)	10.799	(0.267)
		1	0.0409	7.098	(0.121)	7.098	(0.121)
Czech Republic	2	0	0.0888	19.506	(0.063)	16.004	(0.048)
		1	0.0202	3.503	(0.491)	3.503	(0.491)
Hungary	2	0	0.0564	10.531	(0.098)	9.919	(0.084)
		1	0.0036	0.612	(0.495)	0.612	(0.495)
Poland	4	0	0.1032	26.897	(0.037)	18.189	(0.074)
		1	0.0508	8.709	(0.199)	8.709	(0.199)
Romania	4	0	0.0789	25.261	(0.059)	14.459	(0.225)
		1	0.0595	10.802	(0.095)	10.802	(0.095)
Slovak Republic	2	0	0.0660	16.660	(0.146)	11.403	(0.223)
		1	0.0310	5.257	(0.256)	5.257	(0.256)

Note:The critical values for trace test and maximun eigenvalue statistics are from Pesaran et al. (2000); p-values in round brackets; in the second column we report the usual classification of VEC models with respect to the deterministic component included. In particular, when there no constant and trend are in the data we have model 1; when only a costant term is present in the cointegration space we have the model 2; with a trend in the VAR we have the model 3, while with a costant and a trend in the cointegration space we have model 4. Finally, when we allow for a quadratic deterministic trend in the data we have model 5 (Juselius, 2006).

Table 5 – Forecast error variance decompositions of Real exchange rate

Months ahead	*Permanent shock*	*Temporary shock*	*Standard error*
		Albania	
1	49.006	50.994	0.033
3	51.763	48.237	0.035
6	54.957	45.043	0.037
12	55.900	44.100	0.039
24	55.922	44.078	0.039
60	55.889	44.111	0.039
		Bulgaria	
1	0.954	99.046	0.073
3	12.747	87.253	0.080
6	12.563	87.437	0.086
12	12.781	87.219	0.088
24	12.737	87.263	0.089
60	12.729	87.271	0.089
		Czech Republic	
1	58.798	41.202	0.021
3	57.959	42.041	0.022
6	57.556	42.444	0.022
12	59.015	40.985	0.023
24	59.072	40.928	0.023
60	59.039	40.961	0.023
		Hungary	
1	89.567	10.433	0.019
3	88.675	11.325	0.019
6	86.110	13.890	0.019
12	84.315	15.685	0.020
24	83.962	16.038	0.020
60	83.923	16.077	0.020
		Poland	
1	79.207	20.793	0.026
3	78.751	21.249	0.027
6	76.742	23.258	0.027
12	75.853	24.147	0.027
24	75.830	24.170	0.027
60	75.827	24.173	0.027
		Romania	
1	63.029	36.971	0.046
3	66.530	33.470	0.049
6	66.415	33.585	0.049
12	66.581	33.419	0.050
24	66.594	33.406	0.050
60	66.594	33.406	0.050
		Slovak Republic	
1	98.626	1.374	0.018
3	97.925	2.075	0.019
6	97.921	2.079	0.019
12	97.921	2.079	0.019
24	97.921	2.079	0.019
60	97.921	2.079	0.019

Note. The percentage of the variance of real exchange rate explained by permanent and temporary shocks. Standard errors are obtained using bootstrapping methods. The simulation horizon is equals to 60 months.

Table 6 –Correlations between different specification of nominal shocks

	Albania	*Bulgaria*	*Czech Republic*	*Hungary*	*Poland*	*Romania*	*Slovak Republic*	*EU*
correlation	0.366	0.320	0.451	0.688	0.481	0.628	0.752	0.829
p-value	(0.000)	(0.005)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)

Note. In the upper (lower) part of the table we report the cross-correlations between transitory (permanent) shocks in each transition country and Euro Area.

Table 7 –Descriptive statistics of trade intensities

		T1		T2		T3	
		average	*st.dev*	*average*	*st.dev*	*average*	*st.dev*
Czech Republic	*Total sample*	0.010	0.003	0.009	0.002	0.014	0.003
	1993-95	0.006	0.001	0.005	0.000	0.006	0.001
	1996-98	0.009	0.001	0.008	0.001	0.008	0.001
	1999-01	0.010	0.001	0.010	0.001	0.010	0.001
	2002-04	0.012	0.001	0.013	0.001	0.013	0.001
	2005-07	0.014	0.001	0.016	0.001	0.015	0.001
Hungary	*Total sample*	0.004	0.001	0.010	0.004	0.009	0.003
	1993-95	0.006	0.001	0.004	0.000	0.005	0.000
	1996-98	0.008	0.001	0.007	0.001	0.007	0.001
	1999-01	0.010	0.001	0.010	0.001	0.010	0.001
	2002-04	0.011	0.001	0.011	0.001	0.011	0.001
	2005-07	0.012	0.001	0.012	0.001	0.012	0.001
Poland	*Total sample*	0.011	0.003	0.005	0.001	0.010	0.003
	1993-95	0.009	0.001	0.008	0.001	0.008	0.001
	1996-98	0.014	0.002	0.008	0.001	0.011	0.001
	1999-01	0.014	0.001	0.010	0.001	0.012	0.001
	2002-04	0.016	0.001	0.013	0.001	0.014	0.001
	2005-07	0.018	0.001	0.016	0.001	0.017	0.001
Slovak Republic	*Total sample*	0.009	0.003	0.013	0.003	0.004	0.001
	1993-95	.	.	.	.	.	.
	1996-98	0.003	0.000	0.003	0.000	0.003	0.000
	1999-01	0.003	0.000	0.004	0.000	0.003	0.000
	2002-04	0.005	0.001	0.005	0.001	0.005	0.001
	2005-07	0.005	0.000	0.006	0.001	0.006	0.001
	2005-07	0.005	0.000	0.006	0.001	0.006	0.001

Note Mean and standard deviation of the three measures of trade intensities expressed by equations (12), (13) and (14) over the period 2003:1-2007:12.

Table 8 –Trade integration and shoks correlation

transitory shock		*correlation*		*OLS estimates*	
	T_{iEUt}^{IM}	T_{iEUt}^{EX}	T_{iEUt}^{T}	a_2 *coefficient*	*Std. error*[2]
Czech Republic	-0.010(0.894)	0.117(0.131)	0.067(0.389)	0.064	0.078
Hungary	0.459(0.000)	0.570(0.000)	0.528(0.000)	0.529	0.066
Poland	0.801(0.000)	0.710(0.000)	0.788(0.000)	0.787	0.048
Slovak Republic	0.191(0.029)	0.393(0.000)	0.327(0.000)	0.182	0.047
permanent shock		*correlation*		*OLS estimates*	
	T_{iEUt}^{IM}	T_{iEUt}^{EX}	T_{iEUt}^{T}	a_2 *coefficient*	*Std. error.*
Czech Republic	0.675(0.000)	0.724(0.000)	0.713(0.000)	0.713	0.054
Hungary	0.906(0.000)	0.882(0.000)	0.906(0.000)	0.905	0.033
Poland	0.178(0.021)	0.301(0.000)	0.257(0.001)	0.257	0.075
Slovak Republic	0.410(0.000)	0.610(0.000)	0.549(0.000)	0.300	0.040

Note: In the second, third and fouth column the correllations between the measure of shoks syncronization and the indexes of trade intensity are reported (*p-value* in round brackets). In the last two columns are shown the estimates of the coefficent a_2 in the equation (16) and their standard errors.

Table 9 –Trade integration and shoks correlation (quarterly observations)

transitory shock		*correlation*		*OLS estimates*	
	T_{iEUt}^{IM}	T_{iEUt}^{EX}	T_{iEUt}^{T}	a_2 *coefficient*	*Std. erorr*
Czech Republic	0.004(0.987)	0.112 (0.411)	0.070(0.607)	0.070	0.607
Hungary	0.477(0.000)	0.606(0.000)	0.558(0.000)	0.558	0.112
Poland	0.813(0.000)	0.714(0.000)	0.796(0.000)	0.796	0.082
Slovak Republic	0.194(0.211)	0.406(0.006)	0.333(0.006)	0.186	0.082
permanent shock		*correlation*		*OLS estimates*	
	T_{iEUt}^{IM}	T_{iEUt}^{EX}	T_{iEUt}^{T}	a_2 *coefficient*	*Std. error.*
Czech Republic	0.712(0.000)	0.723(0.000)	0.724(0.000)	0.677	0.100
Hungary	0.909(0.000)	0.870(0.000)	0.901(0.000)	0.900	0.059
Poland	0.182(0.179)	0.313(0.024)	0.260(0.053)	0.259	0.131
Slovak Republic	0.446(0.002)	0.607(0.000)	0.558(0.000)	0.297	0.068

Note: In the second, third and fouth column the correllations between the mesaure of shoks syncronization and the indexes of trade intensity are reported (*p-value* in round brackets). In the last two columns are shown the estimates of the coefficent a_2 in the equation (16) and their standard errors.

Table 10 –Trade integration and shoks correlation in between transition countries and EU

Dependent variable	[1]	[2]	[3]	[4]	[5]	[6]
	Synchronization measure of transitory shocks			*Synchronization measure of permanent shocks*		
constant	0.078	0.078	0.078	-0.081	-0.081	-0.081
	(0.059)	(0.059)	(0.059)	(0.052)	(0.050)	(0.051)
T_{iEUt}^{IM}	0.364			0.526		
	(0.059)			(0.053)		
T_{iEUt}^{EX}		0.426			0.568	
		(0.057)			(0.051)	
T_{iEUt}^{T}			0.415			0.560
			(0.058)			(0.051)
F-test	2.27[0.081]	2.44[0.066]	2.40[0.069]	3.07 [0.029]	3.34[0.020]	3.28[0.022]
R-sq- within	0.155	0.211	0.201	0.325	0.379	0.368
R-sq -between	0.355	0.014	0.027	0.355	0.014	0.027
R-sq -overall	0.150	0.206	0.196	0.315	0.368	0.357

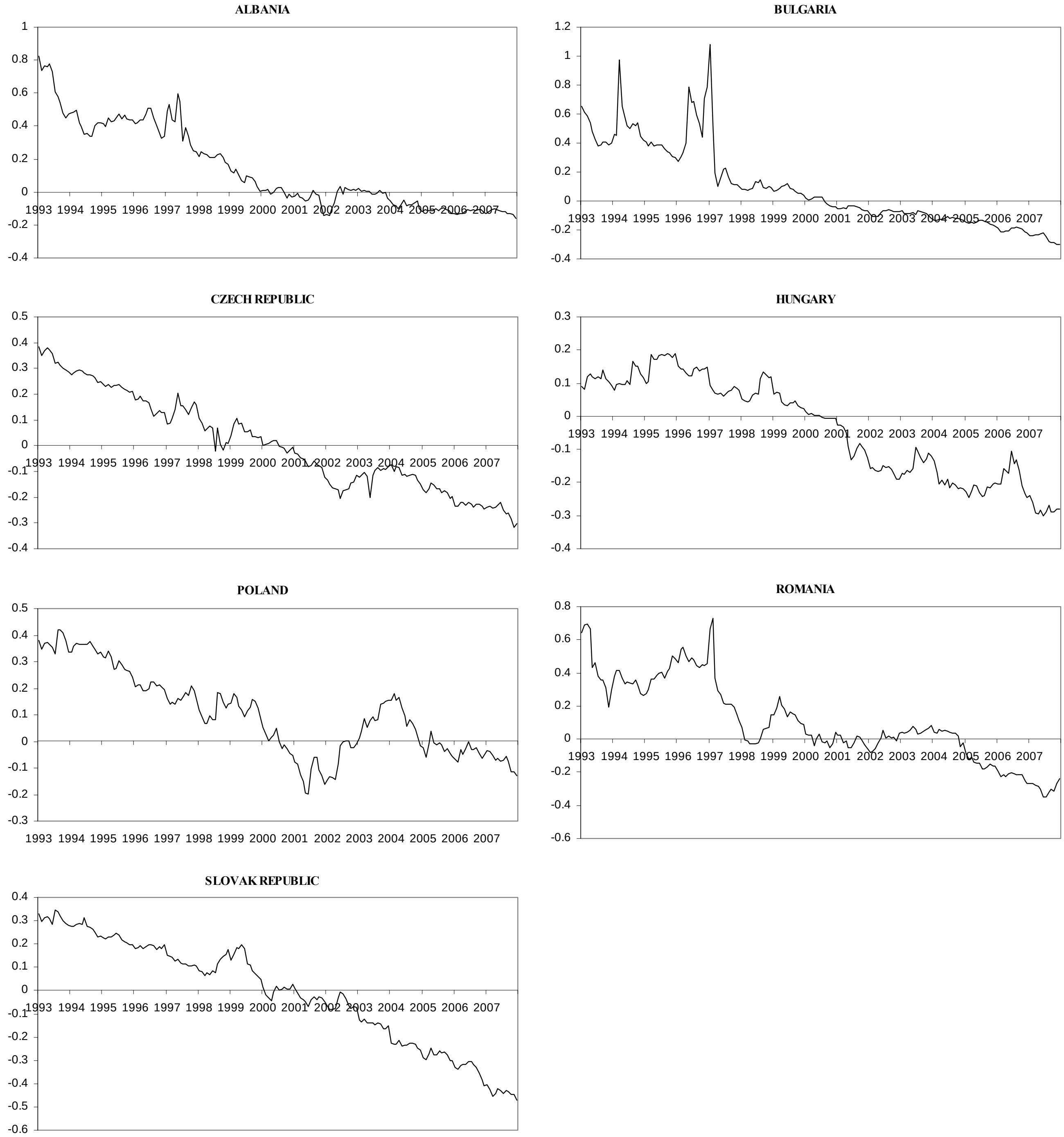
Note: All models refer to panel with fixed effects using quarterly observations from 1994:1 to 2007:4. Standard errors in parentheses. *P-values* for F-test for joint significance of fixed effects in square brackets.

Table 11 –Robusteness analysis: Forecast error variance decompositions of Real exchange rate using quarterly data

Quarters ahead	*Real shock_NTR*	*Real shock_TR*	*Nominal shock*	*Standard error*
		Bulgaria		
1	1.882	28.951	69.167	0.087
2	9.301	18.427	72.272	0.111
4	37.598	11.399	51.003	0.142
8	57.406	8.169	34.425	0.174
20	50.896	21.941	27.163	0.207
		Czech Republic		
1	21.772	77.447	0.781	0.032
2	25.603	73.451	0.947	0.033
4	25.752	73.157	1.091	0.033
8	25.754	73.154	1.092	0.033
20	25.754	73.154	1.092	0.033
		Hungary		
1	12.554	87.240	0.206	0.039
2	12.196	86.309	1.495	0.040
4	13.819	83.968	2.214	0.041
8	13.844	83.794	2.363	0.041
20	13.844	83.792	2.364	0.041
		Poland		
1	0.979	88.836	10.185	0.049
2	4.018	80.855	15.128	0.052
4	10.156	75.434	14.410	0.055
8	13.930	68.236	17.835	0.058
20	20.854	61.148	17.998	0.062
		Romania		
1	0.683	98.654	0.664	0.065
2	32.538	66.298	1.164	0.079
4	32.588	66.240	1.172	0.079
8	32.588	66.239	1.172	0.079
20	32.588	66.239	1.172	0.079
		Slovak Republic		
1	19.381	79.415	1.203	0.035
2	19.358	79.231	1.410	0.035
4	19.090	79.278	1.632	0.036
8	19.265	79.069	1.666	0.036
20	19.323	79.003	1.674	0.036

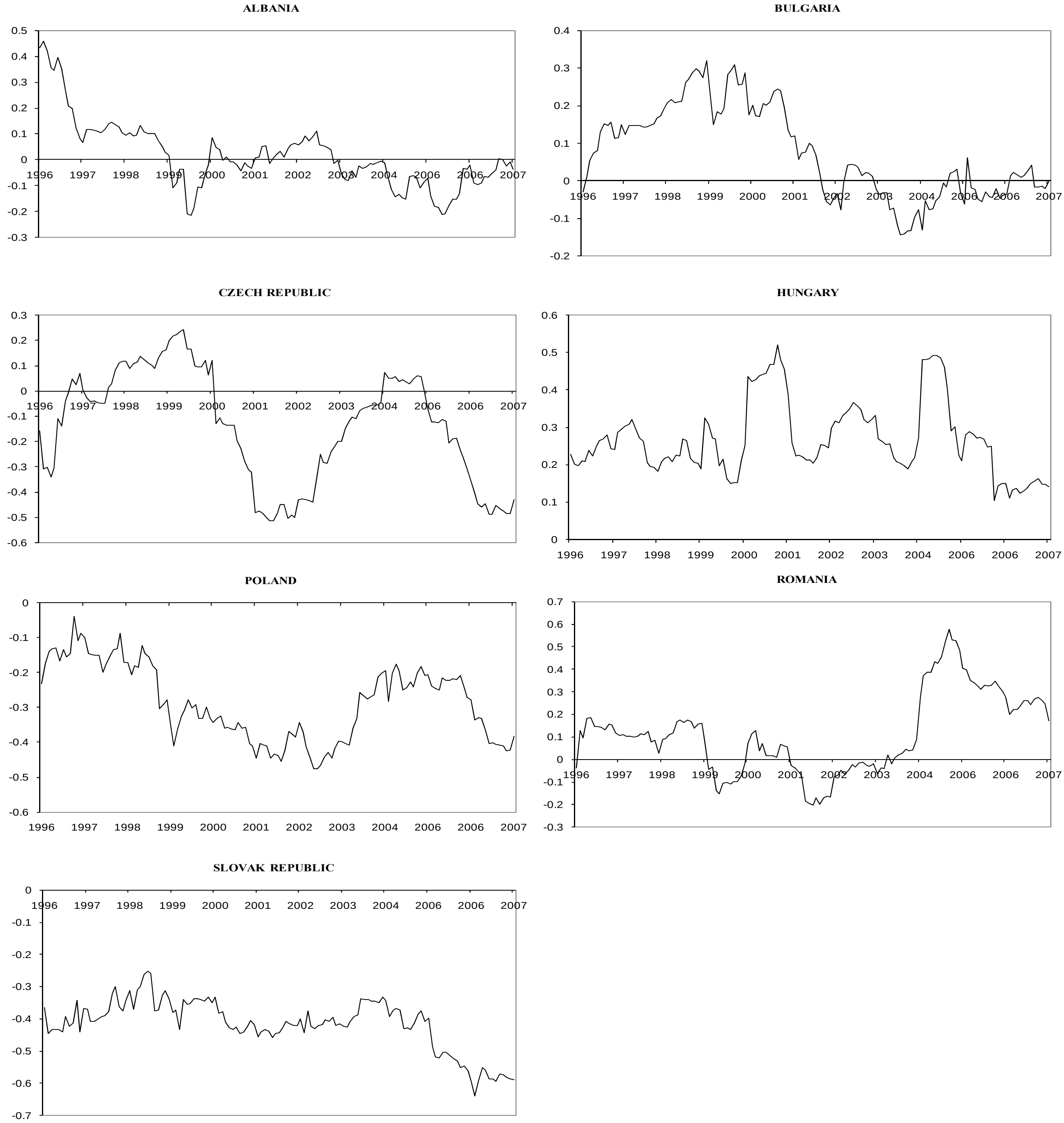
Note. The percentage of the variance of real exchange rate explained by real shocks in nontraded sector (second column), real shocks in traded sector (third column) and nominal shocks. Standard errors are obtained using bootstrapping methods.

Figure 1 – Real exchange rates in transition countries



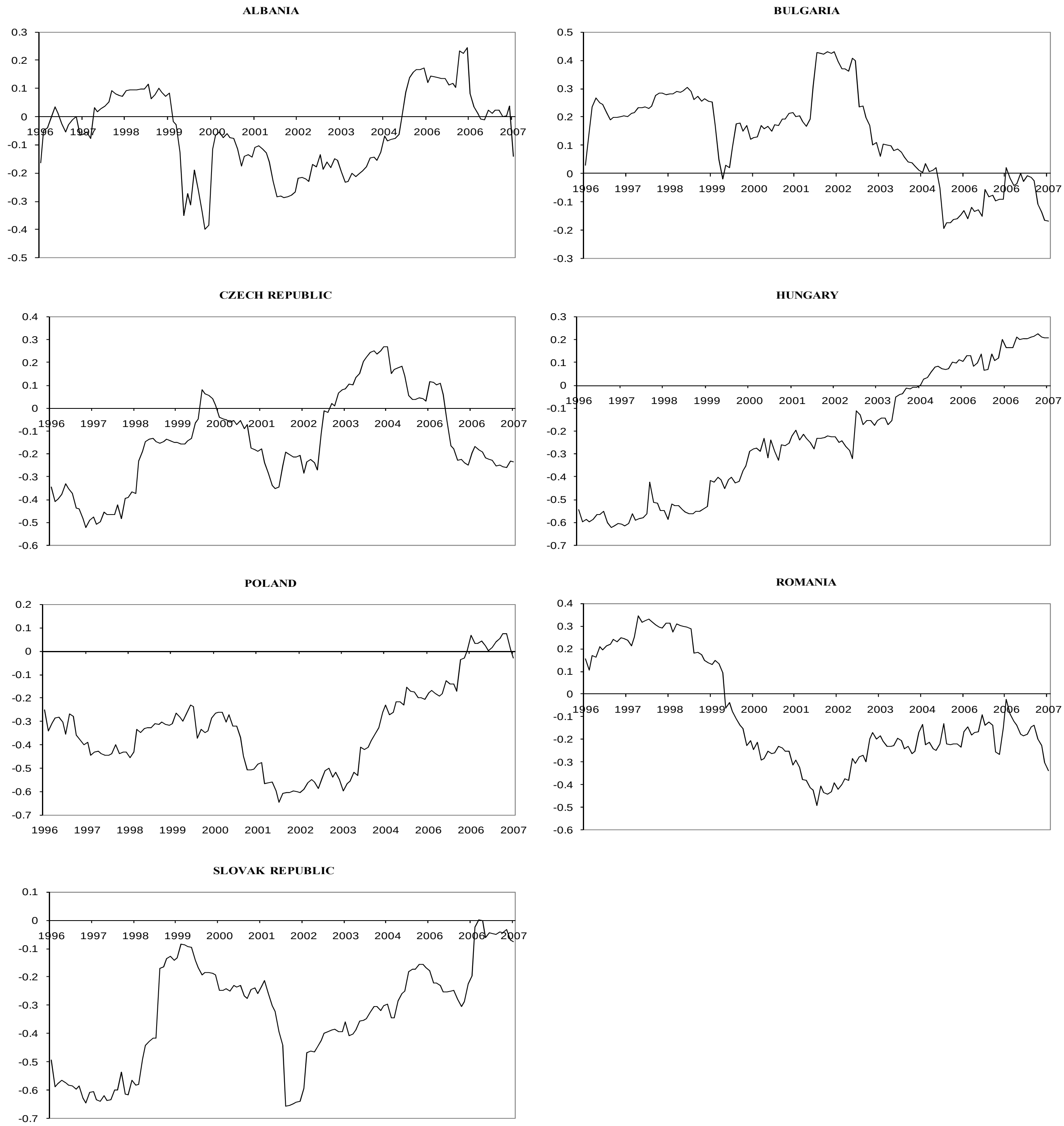
Note: The real exchange rate is defined as the product of the nominal exchange rate (national currency per Euro) and the ratio between Euro and domestic prices. All the variables are expressed in natural logarithms and constant prices (base year 2000=1).

Figure 2 – Moving window correlations between transitory shocks



Note: Correlations computed on a moving window of 36 monthly observations starting from 1994:1.

Figure 3 – Moving window correlations between permanent shocks



Note: Correlations computed on a moving window of 36 monthly observations starting from 1994:1.

Figure 4 – Time-varying correlations between transitory shocks

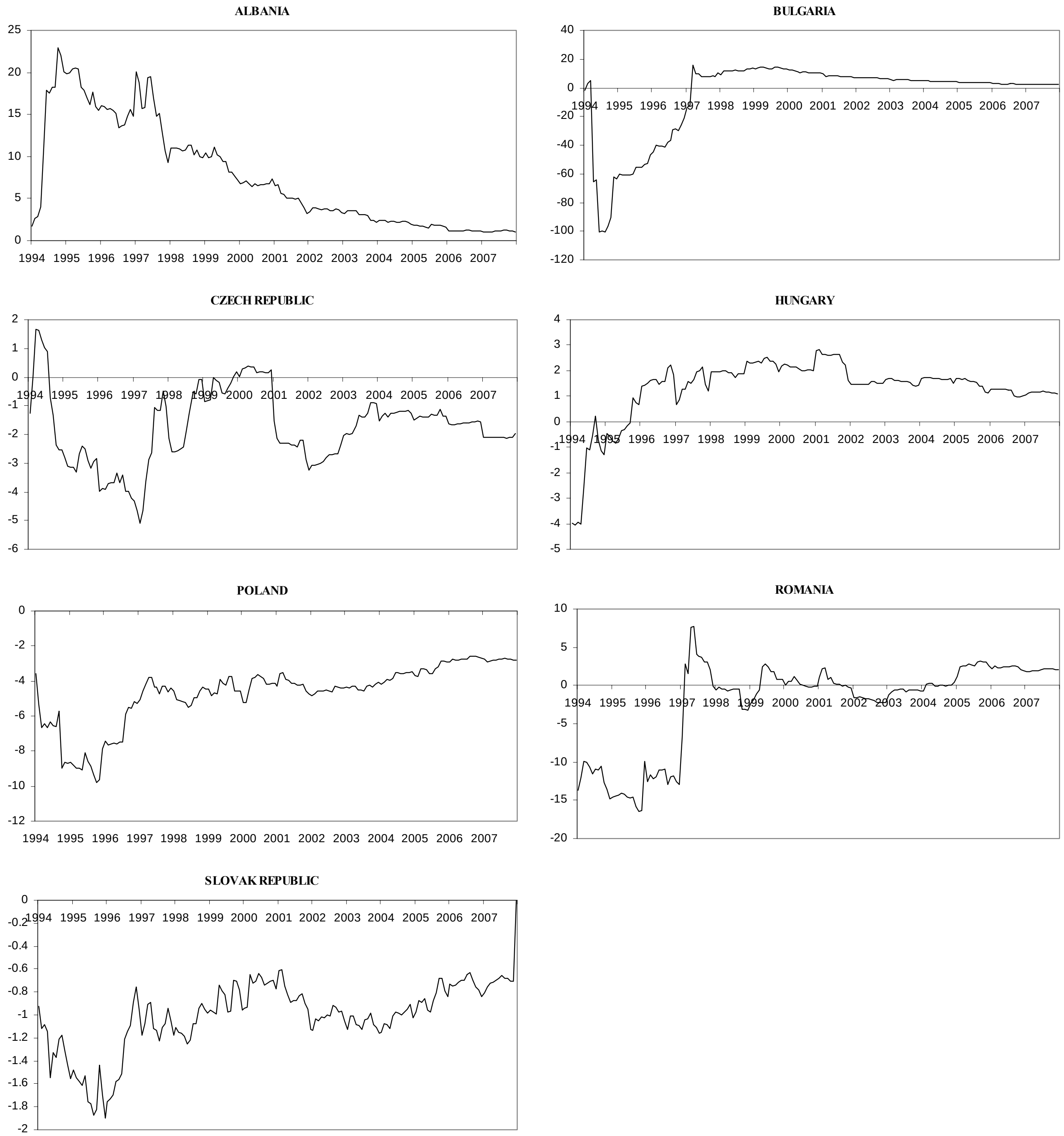
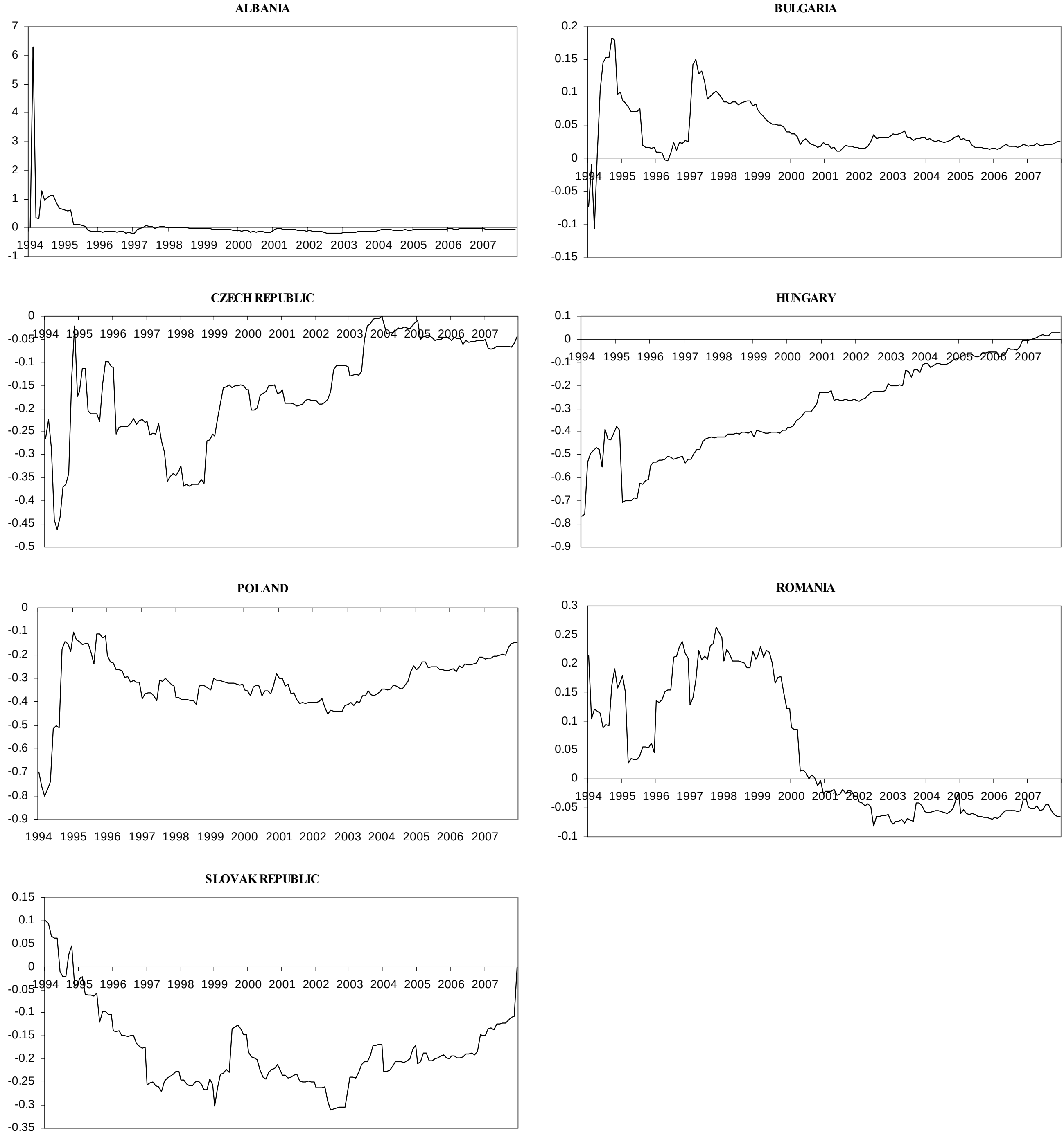


Figure 5 – Time-varying correlations between permanent shocks



REAL EXCHANGE RATE DYNAMICS IN BALTIC COUNTRIES

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 underpinning is based on the Generalised Purchasing Power Parity (GPPP) hypothesis, which is empirically tested within a Vector Error Correction framework. Using both monthly and quarterly data over the period 1993-2005, our results suggest that GPPP holds for the real exchange rate of each Baltic country vis-à-vis the Euro area, reflecting a degree of real convergence consistent with Optimum Currency Area criteria. Further, our 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, purchasing power parity, vector error correction models*

1. Introduction

After the European Union (EU) enlargement of 2004, the expected step ahead in the European integration process involves 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. Eastern European transition economies' (Balkan, Baltic, Central Europe, and Commonwealth of Independent States, CIS, [33] countries) willingness to shore the EMU is motivated by the expectation of enjoying the benefits of a monetary union, in particular policy discipline, deeper economic and financial integration with incumbent countries, a reduction in transaction costs, and lower interest rates (Lättemäe and Randveer 2004).

Among 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), 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 Baltic economies, once they achieved their independence at the end of 1991. 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 question how ready the accession countries are should be answered 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. These reflect policy choices being made during the transition process to achieve exchange rate stabilisation as well as the control of the inflation.

From a more theoretical perspective, price convergence represents a necessary condition in order to stabilize both the nominal (explicit policy target) and the real exchange rate

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

(implicit policy target), allowing to safeguard member countries' intra-regional competitiveness and to avoid the incentive to implement "beggar thy neighbours" policies. However, Ahn et al. (2002) stress that a Purchasing Power Parity (PPP)-based approach may be partial in capturing the major changes in economic policies and the significant restructuring processes 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 the non-stationarity of their long-run macroeconomic determinants. In practice, the GPPP hypothesis between the EMU and 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. Thus, 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 the preliminary assessment of the stationarity of each real exchange rate. The econometric methodology is based on Vector Error Correction (VEC) models. More in details, this paper aims at assessing in sequence: a) whether the GPPP hypothesis hold 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 in driving the international competitiveness of those economies; d) the properties of our multivariate time-series approach in out-of-sample forecasting exercises with respect to a number of alternative models.

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. The empirical framework

The desirability for a given country of joining a common currency 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 evaluate whether the timing and modalities of implementation of a currency area can be considered optimal (Eichengreen, 1990). Recent contributions on the issue at stake shift the emphasis of the debate away from the role of exchange rates as output stabilizer and instead focus on public finance argumentations. Under this perspective, whether domestic money may be managed by the government as a tool for budgetary finance throughout 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, 2000; Canzoneri and Rogers, 1990; Aizenman, 1992). Most researchers point out that these causative factors can hardly be identified when theory is put into empirical testing (Baldwin, 1991; Buiter, 2000). Moreover, there is disagreement within the academic circle on the economic effects of monetary integration with respect to income correlation among member countries and intra-area trade flows.[34]

Along the lines of the approach popularized by Rose (2000), a number of works have analysed 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 throughout gravity models (De Nardis and Vicarelli, 2003, among others). Here, instead, we focus mainly on the role of cross-country income correlations as most influential theoretical argumentation in order to establish whether Baltic countries are ready to join the Euro area. In many academic and policy circles (Artis, 2003), instead, such a canonical criterion is still considered a first useful framework to consult when deciding upon the adoption of common currency.

More in details, 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 hypothesis) between the Baltic countries and the Euro area. According to the GPPP theory, bilateral real exchange rates individually non-stationary may 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

[34] 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). In this perspective, the economic systems of each member country of an OCA would 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

functions. If it is possible to identify (at least) one stationary linear combination of otherwise integrated 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 regional model for the Baltic economies

As already mentioned, the Baltic countries have some distinguishing features with respect to the other Eastern Europe transition economies. After their separation from the former Soviet Union most of their trade was oriented towards the neighbouring Western European countries. Moreover, the main explicit policy objectives were defined taking into account the possibility of joining the 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.[35]

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), 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:

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

where y_{it} stands for domestic output; $y_{it}^{*} = \sum_{j=0,\, j\neq i}^{4} \xi_{ij} \cdot y_{it}$ indicates for foreign output, obtained as 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 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:

[35] 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).

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

or equivalently:

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

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

2.2. Empirical strategy

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:

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

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

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

where $\boldsymbol{\Sigma}_\varepsilon$ 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:

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

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 $\boldsymbol{\Pi}^q \cdot \mathbf{C} = \mathbf{C} \cdot \boldsymbol{\Pi}^y$.

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 matrix $\boldsymbol{\alpha}$ contains the feedback coefficients (loadings) and 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 discriminate 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 of stochastic properties of the joint dynamics of real exchange rates (our observational variables) closely mirroring the one for domestic output time-series (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 unit root tests

Monthly data for the real exchange rates of the Euro area, 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 rate 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 exercise, 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)[36]. In keeping with many studies of the PPP hypothesis in transition countries (see, among others, Taylor and Taylor 2001 and Égert 2004), in each case, we are unable to reject the unit root-null hypothesis at conventional nominal levels of significance, even when controlling for breaks in the series. On the other hand, 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]

3.2 Model specification

Estimating (5) requires taking two steps. First, the lag length p is chosen so estimated residuals resemble the multi-normal distribution as closely as possible, this being an essential requirement for a correct statistical inference. Second, the long-term component of the model is identified on the basis of the trace test 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 estimated residuals match the multi-normal distribution in a satisfactory way both at single equation and system level [37]. Furthermore, Chow tests (available on request) indicate the presence of no residual instability in the model.

[TABLE 3]

Trace and maximum eigenvalue test statistics suggest the presence of three cointegration relationships in the system at the 5 percent significance level[38].

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[36] Critical values for these tests are provided by MacKinnon (1996). A constant term is included in each regression, while the number of lags is chosen such that no residual autocorrelation is evident in the auxiliary regressions. We have also carried out alternative unit root tests (Philips and Perron and KPSS) to check for robustness. The results are qualitatively similar (available from the authors upon request).

[37] 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 of misspecification.

[38] 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 must be constant If $\lambda_i \to 0$. The first *r* trace statistics should grow linearly, while the other ones must be constant over the time. The graph shows that the first three statistics grow in fact linearly as expected, while the fourth one is less clearly increasing (graphs for recursive trace test statistics are available from the authors upon request).

In the rest of the Table exclusion, stationary and weakly exogeneity tests are reported. Testing separately the null hypothesis of each coefficient being equal to zero against the alternative suggests the results show that all variables are statistically different form zero [Panel (b)]. Further, none of variable is stationary by itself in the cointegration space, in a way consistent with the univariate unit root and stationarity tests [Panel (c)]. Finally, except for the Euro/US dollar equation, there is no clear evidence of weakly exogeneity [Panel (d)].

[TABLE 4]

3.3 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 χ^2-distributed LR ratio test with 3 degrees of freedom, the test statistics (4.63), calculated using the Bartlett small-sample correction (with 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).[39]

This is one of the most important findings of the 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).[40] To model deviations from PPP, many authors

[39] Notice that these conclusions are robust even controlling for possible parameter instability. When, we perform recursive estimates of likelihood ratio test the results remain stable. In practice, we perform the estimation for the sub-sample 1993:1 to 1997:9, imposing the restrictions on the three vectors and computed the value of the LR test. Then we add recursively one observation until the end of the estimation sample (2005:12) and in each step we re-calculated the value of the test. Plotting these values, we note that the restrictions are favored by the data over the recursive sample, since the statistics are always below the value of the χ^2 with 3 degree of freedom. (The graph of recursive LR test statistics is not reported and is available on request).

[40] 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 an 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

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).[41] Nevertheless, in the case of Baltic economies, De Broeck e Sløk (2001) argue that even if productivity gains in the tradable 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 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 common currency union.

Figure 1 presents the cointegration relationships from the R-model.[42] There appears to be a clear cointegrating relationship in the case of Estonia, especially after the 1996. By contrast the years 1993-1995 are characterized by a slightly higher variability, consistently with the first phase of important reforms implemented from the beginning of 1992. Similar considerations can be done for the Lithuanian case, even if the period of high variability turns to be longer (until the end of 1999) because of a transition process slower than the Estonian one. The graph of the cointegrating relationship for Latvia is qualitatively similar to that of Lithuania, where the higher variability in the first years are likely to be related to the initial currency peg to the Russian ruble and, subsequently, to the IMF's SDRs so as to prevent the hyperinflation experienced by Russia. Only in the last period has the lat been definitely pegged to the Euro.

[FIGURE 1]

3.4 The adjustment process 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

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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.

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

[42] 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, Johanesen, 1995).

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 a cointegration vector analysed 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 Euro-Area, over a simulation horizon of 60 months (5 years).[43]

[FIGURE 2]

In all cases, the convergence towards the steady-state follows a decreasing trajectory, with the adjustments from disequilibrium that come to an end within the fifth year of 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).[44]

Persistence profiles may also represent 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 performed by means a standard two-sided t-test.[45] 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 leads to 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.[46] This confirms that Estonia and Lithuania follow a similar dynamic path, while the absorption process for Latvia is somewhat slower relative to the one of the two other Baltic countries.

3.5 Modelling the short-run: the structure of the α matrix

The short-run dynamics of model (5) is modelled using a parsimonious (subset) VEC model, obtained dropping those parameters of the matrices $\boldsymbol{\alpha}$ and $\mathbf{P}_l^q$ with p-values lower

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

[44] 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.

[45] 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 $f_{k^*} \approx 0$.

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

than a threshold,[47] 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 (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 long run deviation from the long-run equilibrium condition of the two other Baltic countries, while Estonian reform policies seem to have guaranteed a quicker transition process.

4. Dynamic simulation

This Section we move from a reduced-form to a structural representation of the multivariate time-series model so as to ascertain the role of global and idiosyncratic shocks hitting the Baltic countries.

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.

[47] The AIC criterion with $t = 1.60$ is used as a significance threshold level for 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.

In keeping with one of the central messages of 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 which extent global, regional and domestic economic conditions affect domestic real exchange rate variability, considered as proxies of output fluctuations according to our theoretical underpinnings.

To do this, we resort to the forecast error variance decomposition (FEVD) tool, 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 v's from u's implies the unique determination of the $k^2 = 16$ elements in $\mathbf{B}$. In our identification scheme, a first set of 10 constraints arises by assuming that structural shocks are orthonormal. Choosing the cointegration produces $r(k-r) = 3$ 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 in the matrix of the transitory shocks in which the causal order of the variables is chosen following 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 percentage of regional impulses. The last column (mean) presents the average contribution of the shocks over the entire simulation span (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 percent of the forecast variance, compared to 88.4 percent in Latvia and 74.6 percent in Lithuania. In general, this shock accounts for a considerable percentage of the variance of the whole system (86.9 percent).

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 percent), while it is bigger in Latvia (11.5 percent) and even more so in Lithuania (25.25 percent). Moreover, for the latter country the percentage of the regional shock due to the idiosyncratic component is equal to 91 percent. 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 turn out to be realizations of integrated processes.[48] On the basis of the general-to-specific procedure, the specification of our quarterly model includes three lags (in the levels), in a way consistent with the pre-estimation period used in the monthly version (nine lags in the levels). Table 8 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 8]

Table 9 (part a) suggests that both the trace and the maximum eigenvalue statistics indicate the presence of three cointegration relationships [Panel (a)], that all variable are jointly significant in the cointegrating space, and no variable could be considered weakly exogenous and stationary by itself. [Panel (b), Panel (c), and Panel (d), respectively].

[TABLE 9]

The over-identifying restrictions discussed in Section 3 above are still valid for quarterly data: the Bartlett-corrected (with estimated correction factor of 4.53) χ^2-distributed LR test

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

statistics is equal to 2.97 with a p-value of 0.40.[49] This finding provides further evidence on the stationarity of the bilateral real exchange rate, and thus on the joint validity of the PPP condition, between each Baltic country vis-à-vis the Euro area.[50]

The persistence profiles plotted in Figure 4 show a roughly similar in the case of monthly estimates. For all the equation, 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 Latvian and Lithuania long-run equilibrium relationships, the half lives are close to 3 and 1 quarters respectively, a very similar results to those found before. 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 can depend on the first period of simulation where the behaviours of the Estonia's absorption process seems to be characterized by an higher variability with respect both the other two countries and the simulation computed using monthly observations.

[FIGURE 3]

Table 10 reports the coefficients of the subset VECM 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 strongly not-rejected by the data, being the p-value of χ^2-distributed the LR test statistics 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.

[49] The three cointegration vectors are plotted in Figure 3. As in the previous model with monthly observations, the Estonia's vector seems to be stationary, especially after the first period of transition, while the others two show a higher variability of the long run equilibrium conditions, as consequence of a slower and less mature transition processes.

[50] As done previously (see footnote 8), we carried out the analysis using three bivariate VEC models in which we can study separately each real exchange rate of Baltic country and the real exchange rate of the Euro vis-à-vis the US Dollar. The autoregression order for Estonia-Euro, Latvia-Euro and Lithuania-Euro model suggested by usual general-to-specific procedure is three, two and two, respectively. We also include an unrestricted constant for all models. The trace test statistic indicates the presence of an 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 in three different vectors are not rejected by the data 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). These evidences are further confirmation of the results consistency shown in the main text, both for the monthly and quarterly estimates.

[TABLE 10]

In summary, robustness analysis confirms the main results obtained from estimating the base model with monthly observations. In particular, the previous specification of the cointegration space is not rejected by the data, and the adjustment coefficients are also robust, confirming that Estonia exhibits the highest speed of adjustment, while the Lithuania's equation the less affected behaviour. Similarly, the presence of spillover effects between long run equilibrium conditions also appears to be a robust finding.

5.2 Forecasting exercises

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[51]. 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[52] are estimated to 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.

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[51] 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.

[52] 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, 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.

Table 6 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 11]

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 the measure of forecasting accuracy indicates that the unrestricted VEC model is preferable.

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[53] 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.

[53] 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.

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, these results 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 among the Euro area and a group of three transition countries: Estonia, Lithuania and Latvia. Its aim of the study is to test, using a theoretical framework built on the GPPP hypothesis, if the transition process undergone by these economies in recent years make 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 the global and regional structural on real exchange rate of Baltic countries

The results suggest that the GPPP hypothesis hold for each of the real exchange rates of the Baltic countries over the period 1993:1-2005:12, confirming that, as stressed De Broeck e Sløk (2001) the 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 Estonian kroon -Euro real exchange rate, with the presence of *spillover* effects between the equilibrium conditions, especially in the case of Estonia and Latvia.

Forecast error variance decomposition analysis shows that symmetric shocks are the main driving force of real exchange rates in the Baltic countries, with asymmetric shocks playing a more important role in explaining the variability of the real exchange rates in the case of Lithuania and Latvia (less so in Estonia). Quarterly data produce qualitatively similar results, confirming the robustness of all our main findings, and out-of-sample forecasting analysis generally shows that the selected subset VECM specification outperforms rival models of the real exchange rates of the Baltic countries.

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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:

$$\underset{5\times 5}{\boldsymbol{\Xi}} \cdot \mathbf{y}_t = \underset{5\times 4}{\boldsymbol{\Gamma}} \cdot \mathbf{q}_{0t} + \underset{5\times 1}{\boldsymbol{\gamma}} \cdot i_{0t} \tag{A1}$$

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

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

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

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

Recasting system (A1) as:

$$\underset{5\times 5}{\boldsymbol{\Xi}} \cdot \mathbf{y}_t = \begin{bmatrix} \underset{5\times 4}{\boldsymbol{\Gamma}} & \vdots & \underset{5\times 1}{\boldsymbol{\gamma}} \end{bmatrix} \cdot \begin{bmatrix} \mathbf{q}_{0t} \\ -- \\ i_{0t} \end{bmatrix}$$

we obtain:

$$\begin{bmatrix} \mathbf{q}_{0t} \\ -- \\ i_{0t} \end{bmatrix} = \begin{bmatrix} \underset{5\times 4}{\boldsymbol{\Gamma}} & \vdots & \underset{5\times 1}{\boldsymbol{\gamma}} \end{bmatrix}^{-1} \cdot \underset{5\times 5}{\boldsymbol{\Xi}} \cdot \mathbf{y}_t = \underset{5\times 5}{\mathbf{C}^*} \cdot \mathbf{y}_t . \tag{A2}$$

Tables and figures

Table 1 – Hypothesis testing

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

Table 2 – Unit Root Tests

(a) ADF tets	*eu*	*ee*	*lv*	*lt*
Deterministic part	c	c	c	c
Test statistics	-1.33	-3.21	-2.10	-1.92
	Δeu	Δee	Δlv	Δlt
Deterministic part	-	-	-	-
Test statistics	-11.23	-4.77	-4.37	-5.74
(b) UR tests with unknown breaks	*eu*	*ee*	*lv*	*lt*
Deterministic part	c	c	c	C
Test statistics	-1.17	-3.34	-5.20	-0.81
Detected break	2000:12	1994:3	1993:11	1993:12

Note. (a) Statistics are 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 –3.57 to two decimal places if there is a constant (*c*) in the regression, and -2.61 if no deterministic components are included in the regression, while at the 5% level of significance these values are -2.92 and -1.95, 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.* 2001). 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 – VECM 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

	Δ*ee*	Δ*lv*	Δ*lt*	Mean
Global shock	97.60	88.40	74.65	86.88
Regional shock	2.42	11.50	25.25	13.06
Idiosyncratic shock	*72*	*78*	*91*	

Note. The permanent shock is associated to the common trend of the system (that is the bilateral real exchange rate between the US and the Euro area) and represents the global-external shocks that hit in a symmetric way all Baltic countries. Individual temporary shocks identify idiosyncratic disturbances. Idiosyncratic shocks are then aggregated so as to quantify the overall relevance of regional factors in explaining real exchange rate fluctuations. The figures represent the percentage of the variance of each variable of the system explained by global, regional and idiosyncratic shocks, where the latter (in italics) are expressed as a percentage of regional disturbances. The last column (mean) presents the average contribution of the shocks over the entire simulation period (60 months).

Table 7 – Robustness analysis: Unit Root Tests (quarterly data)

	eu	*ee*	*lv*	*lt*
Deterministic part	C	C	c	c
Test statistics	-1.35	-2.44	-2.02	-1.98
	Δ*eu*	Δ*ee*	Δ*lv*	Δ*lt*
Deterministic part	-	-	-	-
Test statistics	-7.16	-2.88	-4.67	-3.75
(b) UR tests with unknown breaks	*eu*	*ee*	*lv*	*lt*
Deterministic part	C	C	c	C
Test statistics	-1.11	-2.97	-1.52	-0.44
Detected break	2002:2	2002:2	1996:4	2004:4

Note. (a) Statistics are 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 –3.57 to two decimal places if there is a constant (c) in the regression, and -2.61 if no deterministics components are included in the regression, while at the 5% level of significance these values are -2.92 and -1.95, 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. 2001). The last row shows the detected break modelled as a shift dummy.

Table 8 – Robustness analysis: misspesification test (quarterly data)

(a) Univariate misspecification tests

	EU	EE	LV	LT
$AR_{(1\text{-}7)}$	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]
$ARCH_{(1\text{-}7)}$	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

$AR_{(1\text{-}7)}$	1.7537 [0.0117]
Normality	12.360 [0.1358]
Heteroscedasticity	0.2488 [1.0000]

Note: p-values in square brackets

Table 9 – 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 trace test and maximun eigenvalue statistics are from Pesaran et al. (2000); (b) p-value in round brackets.

Table 10 – Robusteness analysis: VECM 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

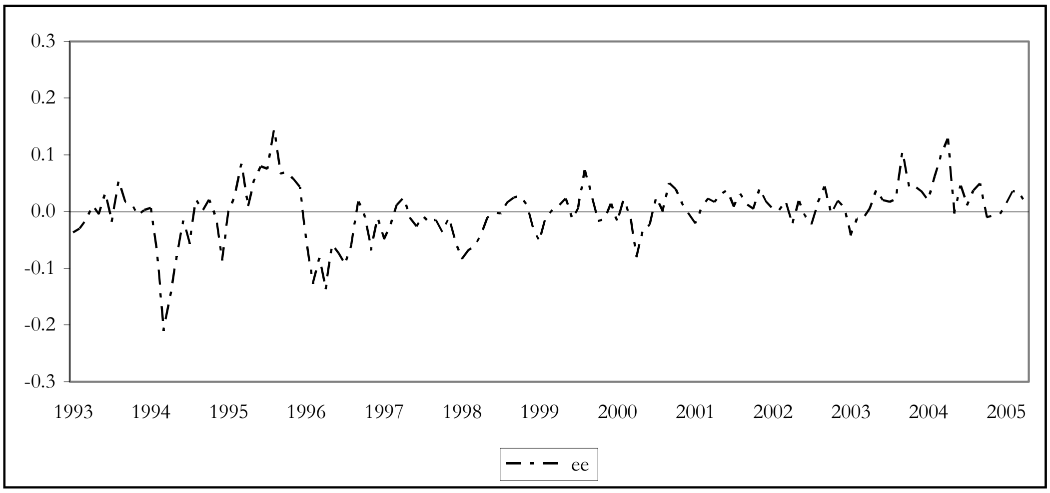
Notes. Standard errors in round brackets.

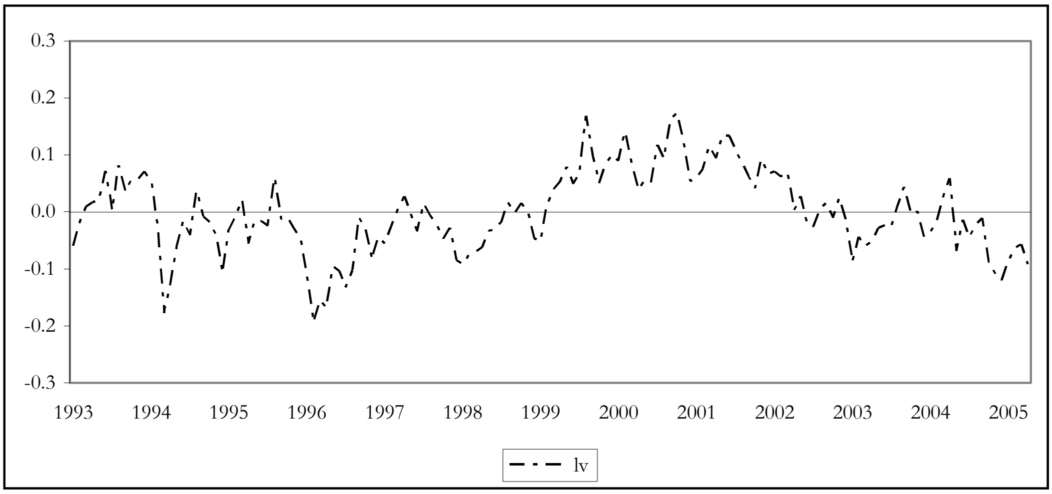
Table 11 – Forecasting accuracy analysis

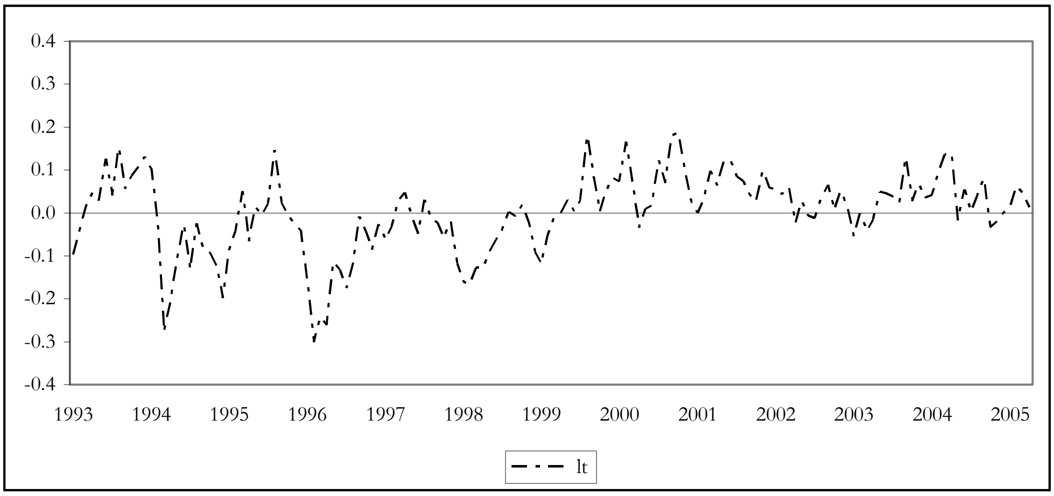
(a) Estonia-Euro	VECM (ristr.)	VECM (unr.)	VAR (rist.)	VAR (unr)	RW
Mean-square error (MSE)					
1-month horizon	7.0e-05	0.9161	0.5207	0.5566	0.5229
3-month horizon	2.4e-04	1.1137	0.5759	0.6047	0.7786
6-month horizon	7.9e-04	1.0138	0.5504	0.6085	0.5670
9-month horizon	9.4e-04	0.8381	0.5974	0.6730	0.8445
10-month horizon	0.0011	0.7770	0.5739	0.6378	0.8131
Mean-absoluta error (MAE)					
1-month horizon	0.0067	0.9510	0.7832	0.7841	0.7572
3-month horizon	0.0144	1.0497	0.7413	0.7555	0.8284
6-month horizon	0.0260	0.9945	0.7476	0.7947	0.7333
9-month horizon	0.0282	0.9202	0.7711	0.8276	0.8650
12-month horizon	0.0263	0.1053	0.8548	0.9126	0.9559
(b) Latvia-Euro	VECM (ristr.)	VECM (unr.)	VAR (rist.)	VAR (unr)	RW
Mean-square error (MSE)					
1-month horizon	3.4e-04	0.8994	0.6230	0.6703	0.4742
3-month horizon	7.1e-04	0.7460	0.6775	0.6171	0.3271
6-month horizon	8.4e-04	0.9569	1.1071	0.9369	0.3020
9-month horizon	7.0e-04	0.7418	1.2946	1.0157	0.6585
12-month horizon	6.6e-04	0.7354	1.5104	1.1602	0.7893
Mean-absoluta error (MAE)					
1-month horizon	0.0137	1.0146	0.7809	0.8281	0.7778
3-month horizon	0.0213	0.8734	0.8681	0.7810	0.6463
6-month horizon	0.0284	0.8862	0.9890	0.9104	0.4524
9-month horizon	0.0238	0.8744	1.1660	1.1440	0.7818
12-month horizon	0.0229	0.8793	1.2696	1.1224	0.8532
(c) Lithuania-Euro	VECM (ristr.)	VECM (unr.)	VAR (rist.)	VAR (unr)	RW
Mean-square error (MSE)					
1-month horizon	1. 3e-04	0.8758	0.7367	0.8258	0.7396
3-month horizon	3.7e-04	0.6247	0.8722	0.6031	1.3670
6-month horizon	4.3e-04	0.6404	0.9738	0.9292	3.6322
9-month horizon	3.8e-04	0.7853	0.9032	0.7722	7.6757
12-month horizon	4.4e-04	0.7074	0.8123	0.7086	7.8547
Mean-absoluta error (MAE)					
1-month horizon	0.0078	1.0613	1.0479	1.0861	1.0679
3-month horizon	0.0163	0.7375	0.9648	0.8070	1.2293
6-month horizon	0.0189	0.8218	0.9924	0.9253	1.9156
9-month horizon	0.0185	0.8731	0.9010	0.8261	2.6545
12-month horizon	0.0194	0.8743	0.8917	0.8229	2.7812

Note: The forecast period is 2006:1 to 2006:6. For 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 the round brackets.

Figure 1 – The cointegration vectors from the R-model

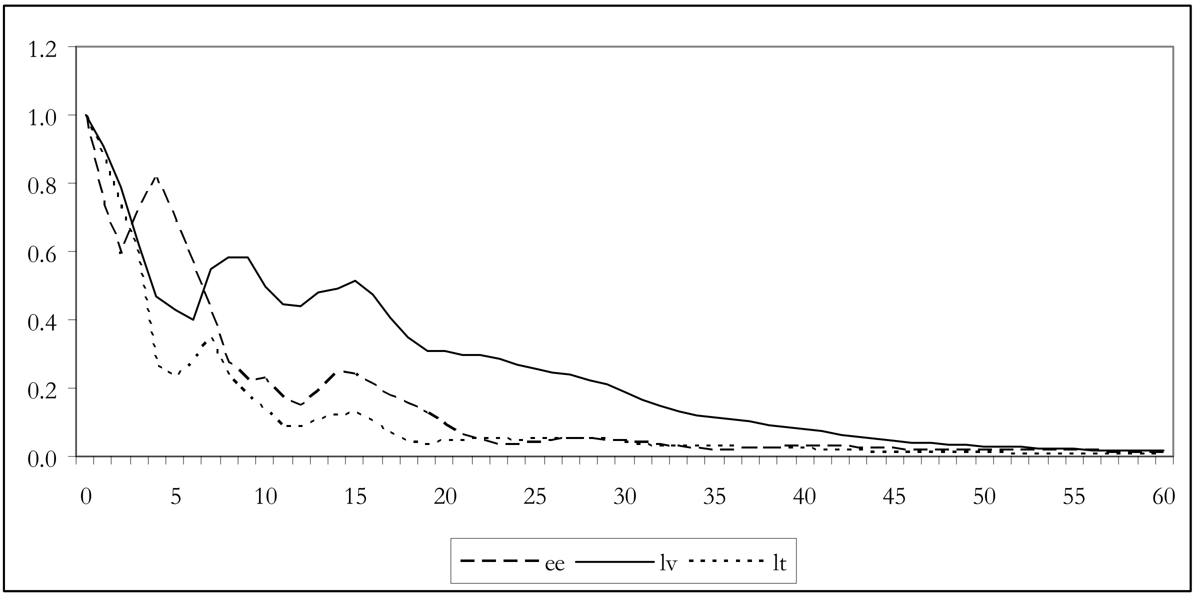






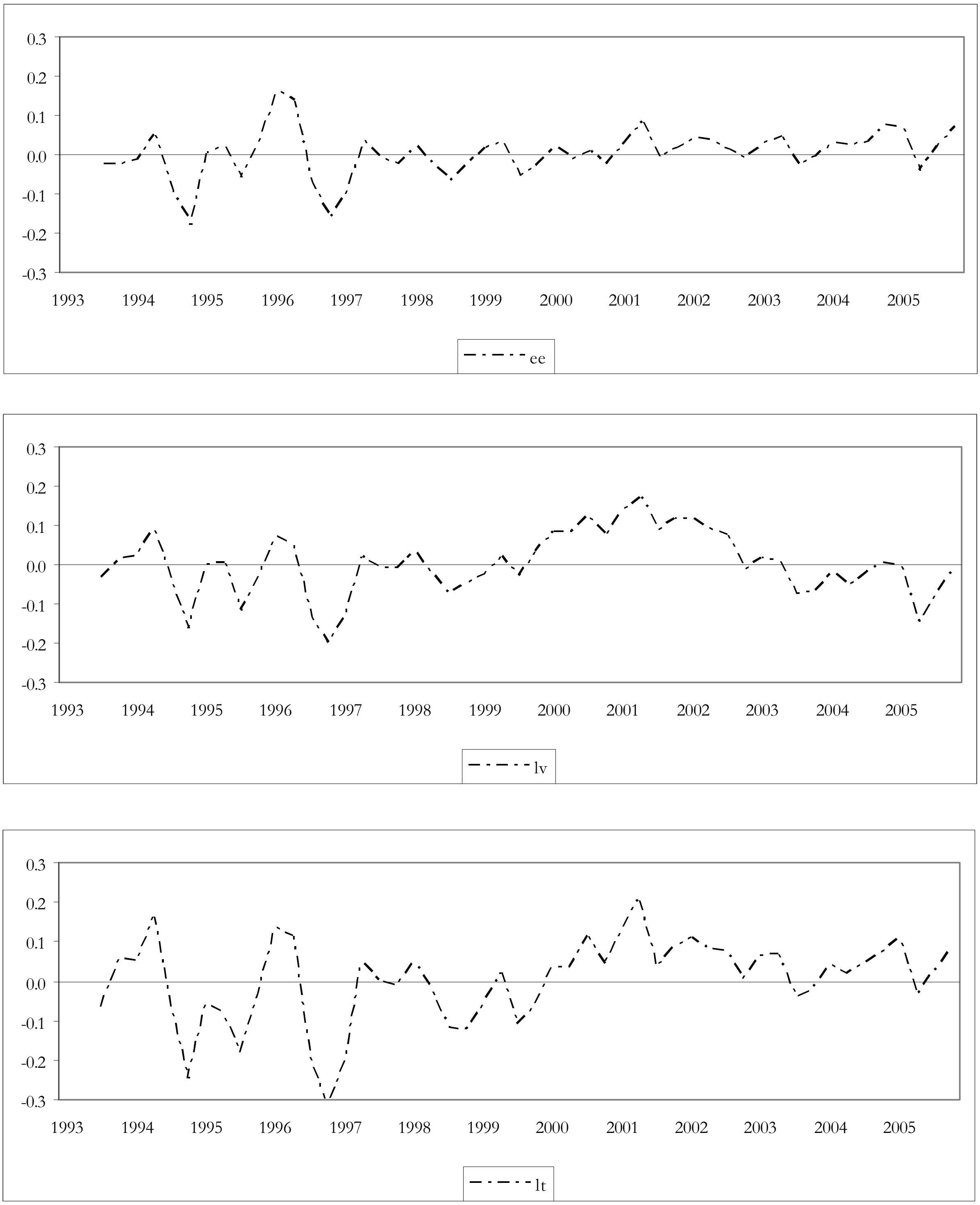
Note. The plots of cointegration vectors are from the R-model. It is computed estimating the ECM representation of the system deleting all dummies and the short-run dynamics. The result is a model where only the long-run properties of the data are isolated.

Figure 2 – Persistence profile of cointegration vectors



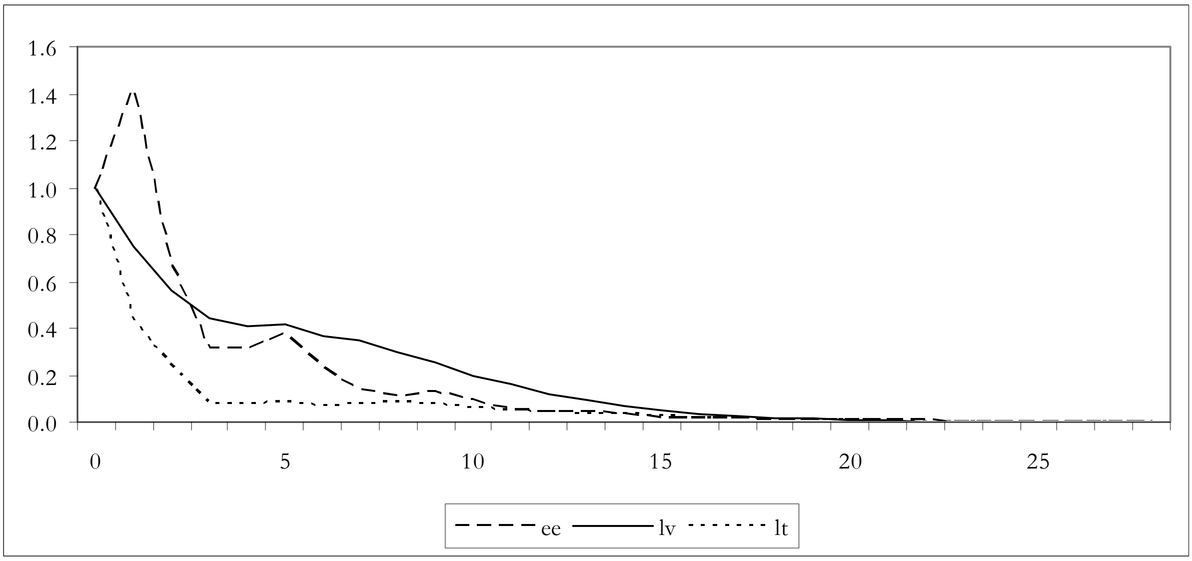
Note. The vertical axis indicates the magnitude of the deviation (normalized to unity on impact) from the steady-state level. The horizontal axis measures the number of months after the shock.

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



Note. The plots of cointegration vectors are from the R-model. It is computed estimating the ECM 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.

Figure 4 – Robustness analysis: persistence profile of 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 measures the number of quarters after the shock.

THE INFLUENCE OF FISCAL SHOCKS ON REAL EXCHANGE RATE IN LATIN AMERICA COUNTRIES

Abstract

This paper analyses the effects of fiscal shocks using a two-country macroeconomic model for output, labour input, government spending and relative prices which provides the orthogonality restrictions for obtaining the structural shocks. Dynamic simulation techniques are then applied, in particular to shed light on the possible effects of fiscal imbalances on the real exchange rate in the case of six Latin American countries. Using quarterly data over the period 1980-2006, we find that in a majority of cases fiscal shocks are the main driving force of real exchange rate fluctuations.

Keywords: *Fiscal shocks, real exchange rate, Latin American countries*.

1. Introduction

In recent years, the macro-economic effects of fiscal shocks have been extensively analysed in the empirical literature (Hemming et al., 2002). Some studies have applied Vector AutoRegression (VAR) methods previously used to analyse the effects of unanticipated monetary impulses. Fiscal shocks have usually been modelled by imposing sign restrictions on the impulse responses (Mountford and Uhlig, 2005), by relying on a Choleski ordering with the fiscal variable appearing first (Favero, 2002), or by exploiting decision lags in fiscal policy and institutional information about the elasticity of fiscal variables to economic activity within a structural VAR framework (Blanchard and Perotti, 2002). Based on the 'narrative approach' of Romer and Romer (1989), another strand of literature has pursued an alternative route where specific exogenous fiscal episodes are isolated through dummies (Burnside et al. 2004, and Christiano, et al. 1999, among others).

Despite this growing empirical literature, there is still no consensus on the size or even the sign of the effects of fiscal shocks on output or the real exchange rate. A possible explanation are the difficulties involved in their identification. As pointed out in Mountford and Uhlig (2005), standard problems encountered in the application of the VAR methodology to assess the effects of fiscal shocks are the following: *i*) how to distinguish movements in fiscal variables which are caused by fiscal policy shocks from those in response to other shocks; *ii*) how properly to define fiscal shocks; *iii*) the existence of a temporal lag between the announcement and the implementation of fiscal policy. The narrative approach makes it possible to circumvent potentially controversial identifying assumptions (typical of the VAR approach), but it has the drawbacks that these episodes could be in part anticipated or that substantial fiscal shocks, of different type or sign, could have occurred around the same time (Perotti, 2002).

In general, all these studies have mainly focused on developed countries and within a closed-economy setup, while the literature on the international transmission of fiscal policy has been almost exclusively theoretical (Baxter, 1992; Bianconi and Turnovsky, 1997; Obstfeld and Rogoff, 1995). An exception is a recent empirical paper by Arin and Koray (2008), who investigate how US fiscal shocks affects the US economy and how they are transmitted to Canada. In the present paper, we focus instead on the effects of fiscal shocks on international competitiveness (the bilateral real exchange rate vis-à-vis the US dollar) in six Latin American (LA) countries (namely, Argentina, Bolivia, Brazil, Chile, Mexico and Peru), over the period 1980-2006.

Since the seminal contribution of Krugman (1979), it is well known among international economists that most of the LA countries suffered speculative attacks on their currencies from international investors mainly because of inconsistencies between domestic macroeconomic policies and the adopted exchange rate regime. In turn, real exchange rate misalignments have often led to macroeconomic disequilibria, and hence the correction of external imbalances might require both demand management policies and real exchange rate devaluations (see, among others, Edwards, 1988). As a result, equilibrium real exchange rates have changed over time, periods of large appreciations being followed by severe depreciations or periods of stability. Furthermore, real exchange rate variability in the LA countries over the eighties was greater than almost anywhere else in the world (Edwards, 1989), owing to debt crises that resulted in a real depreciation of the domestic currency, with frequent devaluations and inflationary episodes.

To the best of our knowledge, the effects of fiscal shocks on the real exchange rates in the LA countries are yet to be investigated in the literature. Previous analyses of the sources of real exchange rate fluctuations have typically focused on the role of real demand (Enders and Hurn, 1994), monetary (Clarida and Gali, 1994; Weber, 1997) or productivity (Alexius, 2005) shocks, and have overlooked the possible effects of fiscal unbalances on countries' international competitiveness. Notable exceptions are the studies of Obstfeld (1993) and Asea and Mendoza (1994), where, in contrast to more traditional monetary approaches, the focus is on the role of fiscal policy and other real variables (such as productivity shocks) in real exchange rate models. Further, only a few studies (Chowdhury, 2004; Hoffmaister and Roldós, 2001) have investigated the sources of real exchange rates fluctuations in developing economies, mainly relying on the approach proposed by Blanchard and Quah (1989) to assess the relative contribution of temporary and permanent disturbances, while the recent paper by Rodríguez and Romero (2007) explicitly analyses the permanent/transitory decomposition of real exchange rates in four LA countries.

We adopt a framework which allows for a wide range of (structural) shocks potentially affecting the real exchange rate. Specifically, we employ a two-country macroeconomic model for output, labour input, government spending and relative prices, along the lines of the studies by Ahmed et al. (1993) and Hoffmaister and Roldós (2001), where the modelling approach to macroeconomic fluctuations developed by Blanchard and Quah (1989) is extended to an open-economy setting allowing for the possible existence of cointegration relationships among the variables of the system. As pointed out by Agénor et al. (2000), macroeconomic fluctuations in developing countries are related to those in industrial economies, and these linkages may have important policy implications for stabilisation and adjustment programmes (Agénor and Montiel, 1996). Applying the same theoretical framework to a relatively homogeneous sample of countries of the same area (namely the LA region), and including the US economy in the analysis as the most appropriate proxy for foreign factors, enables one to establish whether there are empirical regularities across this set of countries, despite their historically different experiences (Ahmed, 2003).

The theoretical model consists of four blocks linked to each other according to a quasi-recursive scheme, and provides the orthogonality restrictions to be imposed to achieve the identification of the structural shocks. These disturbances are identified as supply-side (relative productivity and relative labour inputs) and demand-side (relative fiscal and relative preference) shocks. Their dynamic effects on the real exchange rate are then examined within a structural Vector Error Correction (VEC) framework by means of dynamic simulation (such as forecast variance error decomposition and impulse response analysis) and historical decomposition techniques.

Using quarterly data over the period 1980-2006, we provide clear evidence that fiscal shocks are a key determinant of real exchange rate dynamics for most of the LA countries we consider. However, the sign and the size of the effects of unanticipated fiscal impulses on the level of the real exchange rate vary, reflecting different degrees of productivity of government expenditure. Further, using alternative econometric specifications, we show that the contribution of demand shocks to explaining real exchange rate fluctuations increases when shorter cyclical fluctuations are taken into account. Finally, omitting the cointegration relationships, which we show exist, is found to lead to overestimating the role of demand shocks and underestimating the contribution of fiscal disturbances.

The layout of the paper is as follows. Section 2 describes the econometric model. Section 3 presents the empirical strategy adopted while the Section 4 shows the preliminary empirical results. In Section 4, dynamic simulations based on forecast error variance and historical decompositions are discussed, while robustness analysis is reviewed in Section 5. Some final remarks follow in Section 6.

2. The model

The framework adopted in this Section enables us to study the sources of macroeconomic fluctuations in a bipolar world for a group of six LA economies (namely, Argentina, Bolivia, Brazil, Chile, Mexico and Peru). In line with previous empirical papers on these countries, we assume that the US economy is the relevant foreign country (Berg et al., 2002; Ahmed, 2003).

The chosen model aims to provide a theoretical structure to analyse the role of supply and demand shocks (with particular emphasis on fiscal disturbances) in explaining the fluctuations of the real exchange rate, one of the most common indicators of international competitiveness. This allows us to go beyond the dichotomy between permanent/supply and transitory/demand shocks previously explored in the literature on real exchange rate determination in the developing countries (Chowdhury, 2004; Rodríguez and Romero, 2007).

2.1 The theoretical framework

We analyse the dynamic interactions between the domestic economy and the US using a stylised long-run model, which consists of four blocks linked to each other according to a quasi-recursive scheme. Following Ahmed et al. (1993) and Hoffmaister and Roldós (2001), we rely exclusively on long-run restrictions, which can be directly derived from macroeconomic theory (unlike short-run ones, which can therefore be controversial). In what follows, the suffix i (j) indicates domestic (US) variables, while t indexes time. All variables are expressed in logarithms.

We start by defining a standard production function in the spirit of Binder and Pesaran (1999) and Garratt et al. (2003):

$$y_{it} = \alpha_{1i} + n_{it} + \alpha_{2i}\theta^{y}_{it} \ , \ y_{jt} = \alpha_{1j} + n_{jt} + \alpha_{j}\theta^{y}_{jt} \qquad (1)$$

where y's, α_1's, n's indicate total real output, a generic deterministic component (e.g., an intercept and a linear trend) and labour inputs, respectively, while the θ^{y}'s represent technology shocks driving real output over time. From the equations in (1), relative labour productivity can be expressed as:

$$\pi_t \equiv (y_{it} - n_{it}) - (y_{jt} - n_{jt}) = \alpha + \phi_t \qquad (2)$$

where $\alpha = (\alpha_{1i} - \alpha_{1j})$ and $\phi_t = (\alpha_{2i}\theta^{y}_{it} - \alpha_{2j}\theta^{y}_{jt})$ represents the relative technology shock.

In the long run labour inputs are expected to respond to country-specific exogenous shocks originating in the labour market and/or from permanent changes in government supply policies. Accordingly, we can write down the following functional form for both labour input levels:

$$n_{it} = \beta_{1i} + \beta_{2i}\theta^{n}_{it} \ , \ n_{jt} = \beta_{1j} + \beta_{2j}\theta^{n}_{jt}$$

where the β_1's indicate deterministic components, and the θ^{n}'s represent idiosyncratic labour-supply disturbances. Hence, the relative employment level, n_t, can be expressed as:

$$n_t \equiv n_{it} - n_{jt} = \beta + \upsilon_t \tag{3}$$

where $\beta = (\beta_{1i} - \beta_{1j})$ and $\upsilon_t = (\beta_{2i}\theta^n_{it} - \beta_{2j}\theta^n_{jt})$ is the relative labour-supply shock.

Having defined the stochastic disturbances driving relative labour productivity and relative labour inputs, we move on to modelling the public sector of the two economies. Let $\tilde{g}$ be government size (defined as the ratio of government purchases of goods and services to output); taking the (log of) private output (the difference between total output and government spending) in the two economies, y^P, and using the approximation $\ln(1-x) \cong x$ we obtain the following relationships:

$$y_{it}{}^P = y_{it} - \tilde{g}_{it} \ , \ y_{jt}{}^P = y_{jt} - \tilde{g}_{jt} \tag{4}$$

As in Ahmed et al. (1993), the size of domestic (foreign) government depends both on domestic and foreign permanent fiscal policy shocks, the θ^{y^P} parameters, through a feedback reaction function governed by the γ_2's which measure the response to an exogenous change in the foreign (domestic) government size:

$$\tilde{g}_{it} = \gamma_{1i} + \theta^{y^P}_{it} + \gamma_{2i}\theta^{y^P}_{jt} \ , \ \tilde{g}_{jt} = \gamma_{1j} + \theta^{y^P}_{jt} + \gamma_{2j}\theta^{y^P}_{it} \tag{5}$$

where the γ_1's are constant quantities. Using equations (5) to substitute into (4) we then obtain:

$$y_{it}{}^P = \gamma_{1i} + y_{it} + \theta^{y^P}_{it} + \gamma_{2i}\theta^{y^P}_{jt} \ , \ y_{jt}{}^P = \gamma_{1j} + y_{jt} + \theta^{y^P}_{jt} + \gamma_{2j}\theta^{y^P}_{it}$$

or, in relative terms:

$$z_t \equiv y_{it}{}^P - y_{jt}{}^P = (y_{it} - y_{jt}) + (\varphi_t - \gamma)$$

where $\gamma = \gamma_{1i} - \gamma_{1j}$ and $\varphi_t = [(1-\gamma_{2i})\theta^{y^P}_{jt} - (1-\gamma_{2j})\theta^{y^P}_{it}]$ represents the relative fiscal shock. Using conditions (2) and (3), we can express relative private output as a linear function of the structural shocks:

$$z_t = \alpha + \beta - \gamma + \phi_t + \upsilon_t + \varphi_t \tag{6}$$

Finally, consumers in both economies are assumed to make their consumption decisions to maximise their utility. Adopting a log-linear specification with identical preferences in the two countries, the closed-form solution is such that the (log of) relative prices, q_t, equals the marginal rate of substitution (Ahmed et al., 1993). In turn, the balanced-growth path implies that the ratio of world consumption of each good to total private output of that good is constant (d), ensuring that the following condition holds:

$$q_t = \delta + (\theta^q_{jt} - \theta^q_{it}) + (y^P_{it} - y^P_{jt}) \tag{7}$$

where the θ^q's are time-varying preference shocks entering the agents' utility function. Let $\eta_t = (\theta^q_{jt} - \theta^q_{it})$ be the relative preference shock. Combining (6) and (7), we can express the real exchange rate as:

$$q_t = \alpha + \beta - \gamma + \delta + \phi_t + \upsilon_t + \varphi_t + \eta_t \qquad (8)$$

Equation (8) expresses real exchange rate dynamics as a combination of the underlying disturbances, which are left unrestricted to encompass a large number of competing theories of real exchange rate determination. Choosing a theory rather than another is thus an empirical issue to be determined by the data. Suppose, for instance, that supply-side shocks dominate the dynamics of the q_t variable. This would support empirically the Harrod-Balassa-Samuelson (HBS) view of real exchange rate determination.[54] Consider, instead, the case where φ_t turns out to be the most relevant source of real exchange rate fluctuations. This would give empirical support to the model of Roldós (1995), within which public spending shocks can lead to permanent shifts in the real exchange rate. Next suppose that preference shocks are the main driving factor of q_t. This would be consistent with a general equilibrium, two-country models with a representative utility-maximising agent in the presence of cash-in-advance constraints (Stockman, 1980; Lucas, 1982). Clearly, any of the above-mentioned theoretical hypotheses could be a plausible explanation for the behaviour of the real exchange rate in the LA countries. However, were q_t to depend only on constant terms, this would put into question the empirical validity of the purchasing power parity (PPP) hypothesis, and would be more difficult to rationalise. Recent surveys covering this issue are Sarno and Taylor (2002) and Taylor (2006).

2.2 The long run model

Equations (2)-(3)-(6)-(8) constitute the building-blocks to study the interactions between domestic and foreign economies in the *long run*. As in Ahmed and al. (1993) and Hoffmaister and Roldós (2001), the theoretical relations shown in the previous section describe the long run equilibrium paths of the variables included in the model, and thus they are to rely only on long run restrictions to identify the fundamental disturbances and the dynamics of the variables. From each equation it s possible to extract the set of restrictions needed in order to

[54] See Sarno and Taylor (2002) and Alexius (2005) for the empirical content of this paradigm for developing and industrialised economies.

achieve identification of structural shocks of the system. To this aim, we jointly consider them as forming a tetra-variate system:

$$\begin{bmatrix} \pi_t \\ n_t \\ z_t \\ q_t \end{bmatrix} = \begin{bmatrix} \alpha \\ \beta \\ \alpha+\beta-\gamma \\ \alpha+\beta-\gamma+\delta \end{bmatrix} + \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 1 & 1 & 1 & 0 \\ 1 & 1 & 1 & 1 \end{bmatrix} \cdot \begin{bmatrix} \phi_t \\ \upsilon_t \\ \varphi_t \\ \eta_t \end{bmatrix} \tag{9}$$

Model (9) can be retrieved by inverting an estimated vector autoregression (VAR) in labour productivity, labour input, private output and the real exchange rate. However, it can be done only if system (9) is covariance stationary. In turn, it implies proper transformations of the variables involved in order to achieve stationarity. Ideally, there can be three different cases. If all variables are stationary; then the VAR can be estimated and inverted using variables in the levels. As a polar case, if all elements in the RHS of model (9) are integrated but not cointegrated; it implies first differencing all variables before obtaining the moving average representation. Finally, if there exits one (or more) cointegration relationship(s) among the variable in (9), then we are able to rewrite the system into its common trend representation (Stock and Watson 1988, 1993) where the stationary component(s) (i.e. the cointegration vectors) identify the temporary shock(s) affecting the system, while the integrated components are the stochastic trends.

3. Empirical strategy and identification of structural shocks

3.1 Hypothesis testing hypotheses: a steady-state of the model

As discussed above, the model (9) allows for different empirical estimations according to the long-run behaviour of the variables included in the system. However, as in Alexius (2005), our *a priori* assumption is that there is a *long run*-relationship between real exchange rate and the fundamental variables in the theoretical model. The presence of cointegration[55] in the system allow to identify some transitory disturbances and to quantify how much of the deviations around long-run paths (especially in the case of real exchange rate) is caused by innovations to permanent components and how much by innovations to transitory ones.

Accordingly, We assume that the four variables (relative productivity, relative labour input, relative private output and real exchange rate) are driven by three common stochastic trends (ϕ_t, υ_t and φ_t) in the long-run. These trends evolve over time according to the following laws of motion:

[55] As discussed below, the data are broadly consistent with this assumption and the empirical specification outlined in this section.

$$\phi_t = \phi_{t-1} + \varepsilon_t^{\phi} = \phi_0 + \sum_{i=1}^{t} \varepsilon_i^{\phi} \,,\; \upsilon_t = \upsilon_{t-1} + \varepsilon_t^{\upsilon} = \upsilon_0 + \sum_{i=1}^{t} \varepsilon_i^{\upsilon} \,,\; \varphi_t = \varphi_{t-1} + \varepsilon_t^{\varphi} = \varphi_0 + \sum_{i=1}^{t} \varepsilon_i^{\varphi}$$

where ϕ_0, υ_0 and φ_0 denote initial conditions and the ε's are uncorrelated white-noise processes such that $E(\varepsilon_t^l) = 0$, $E(\varepsilon_t^l)^2 = \sigma_{\varepsilon^l}^2$, $E(\varepsilon_t^l \varepsilon_s^l) = 0$ for $s \neq t$, with $l = \phi, \upsilon, \varphi$.

The model also contains the transitory stochastic component η_t, which is assumed to be orthogonal with respect to ε_t^{ϕ}, ε_t^{υ} and ε_t^{φ} and obeys the following law of motion:

$$\eta_t = \rho\eta_{t-1} + \varepsilon_t^{\eta} = \varepsilon_t^{\eta} / (1 - \rho L) \;,\; \rho < 1$$

where ε_t^{η} is an uncorrelated white noise process.

To find the steady state of the model, the initial values of all the permanent shocks (ϕ_0, υ_0 and φ_0) along with the deterministic component of all the variables of the theoretical model (α, β, γ, δ) are set equal to zero. Accordingly, the steady state can be represented as follows:

$$\begin{bmatrix} \pi_t \\ n_t \\ z_t \\ q_t \end{bmatrix} = \begin{bmatrix} 1 & 0 & 0 \\ 0 & 1 & 0 \\ 1 & 1 & 1 \\ 1 & 1 & 1 \end{bmatrix} \cdot \begin{bmatrix} \phi_t \\ \upsilon_t \\ \varphi_t \end{bmatrix} \tag{10}$$

The long-run structure (10) implies that not only shocks originating from the supply-side of the economy (productivity, ϕ_t , and labour supply shocks, υ_t) but also demand disturbances (namely, fiscal shocks, φ_t) can induce permanent shifts of the steady state of the system. By contrast, relative preference shocks, η_t, are assumed to have transient effects on the levels of the variables. This assumption can be rationalised in terms of the transitory nature of shocks driving demand for domestic and foreign (aggregate) goods. The set of restrictions in (9), as well as those present in the whole system (9), derives directly by theoretical assumptions of the Section 2 and the structural shocks are obtained imposing these restrictions on their long run effects. In the next section we will describe the empirical strategy that we need to achieve the identification of structural shocks mainly based of a common trend representation of the system (9); however it is important to note that our framework, as discussed above, allows for different representations without changes in the causality order of the variable, with the restrictive assumption of cointegration not being strictly required to achieve identification.[56] On the other hand, testing for cointegration is a relevant empirical issue in modelling real exchange rate fluctuations (Alexius, 2005).

3.2 The structural MA-representation

Equations (2)-(3)-(6)-(8) represent the building-blocks to study the interactions between domestic and foreign economies. Adopting the same notation as above, we focus on the following k -dimensional VAR model in error correction form:

$$\begin{bmatrix} \Delta\pi_t \\ \Delta n_t \\ \Delta z_t \\ \Delta q_t \end{bmatrix} = c + \Pi \begin{bmatrix} \pi_{t-1} \\ n_{t-1} \\ z_{t-1} \\ q_{t-1} \end{bmatrix} + \sum_{i=1}^{p-1} \Gamma_i \begin{bmatrix} \Delta\pi_{t-i} \\ \Delta n_{t-i} \\ \Delta z_{t-i} \\ \Delta q_{t-i} \end{bmatrix} + \begin{bmatrix} u_t^{\pi} \\ u_t^{n} \\ u_t^{z} \\ u_t^{q} \end{bmatrix}, \; u_t \sim N(0, \Sigma_u) \qquad (11)$$

where c is a vector of deterministic components, the Γ 's are matrices of autoregressive parameters, Δ is the first difference operator and the vector $u_t = [u_t^{\pi} \;\; u_t^{n} \;\; u_t^{z} \;\; u_t^{q}]'$ contains the estimated residuals. Given our theoretical assumptions, we expect the long-run matrix Π to have rank one, i.e. the presence of one cointegrating vector in model (9).

Structural identification is achieved following the common trends methodology (Warne 1993). Omitting the deterministic component, the reduced-form moving average (MA) representation of the model defines the data generating process (DGP) as a function of the initial conditions (set equal to zero for the sake of exposition) and of the reduced-form shocks u 's. This is given by:

$$x_t = C\sum_{i=1}^{t} u_i + C^*(L)u_t$$

where the matrix C measures the impact of cumulated shocks to the system, $C^*(L)$ is an infinite polynomial in the lag operator L, and $u_t = [\pi_t \;\; n_t \;\; z_t \;\; q_t]'$.

In practise, in the common trend representation, the number of cointegrating vectors *r* determines the number of independent stochastic trends *k* in the common trends model (with *k* = *n-r*). The loading matrix Φ determines how the variables in VAR model are affected by the stochastic trends in the long run (Stock and Watson, 1993). In order to identify the structural shocks, we need to impose *k(k-1)/2* restrictions on the t matrix A. The structural shocks are identified following a theoretical framework described by the equations (2)-(3)-(6)-(8) that imposes a recursive structure which requires a causal order of the type: world technological

[56] For instance, allowing for permanent shifts in demand between domestic and foreign goods would amount to introducing an additional stochastic trend into the system (Ahmed et al., 1993; Hoffmaister and Roldós, 2001). The model would then exhibit four common trends and no cointegration among the variables.

shock, relative labour supply shock, relative fiscal policy shock and relative preference shock The structural shocks are identified by imposing restrictions on their long-run effects. In particular, we are assuming the both relative private output and real exchange rate are affected by all the shocks. Furthermore, fiscal shocks are identified by the assumption that do not affect relative productivity and relative labour inputs, while productivity shocks by the assumption that do not hit the long run behaviours of relative labour inputs. Finally, the model (10) contains an *over-identifying* restriction, that is the labour supply shocks do not affect the relative productivity in the long run[57].

The reduced form and the structural residuals are linked through the relationship $u_t = B\varepsilon_t$, where B is a non-singular matrix (Warne, 1993). Hence, the structural MA representation is the following:

$$x_t = \Phi\sum_{i=1}^{t}\varepsilon_i + \Phi^*(L)\varepsilon_t \tag{12}$$

where the matrix $\Phi = CB$ represents the permanent component of the model, and the matrix polynomial $\Phi^*(L) = C^*(L)B$ the transitory or cyclical component.

Structural identification allows to decompose each of the four time series into the sum of distinct components driven by structural shocks. Focusing on the real exchange rate, q_t, we have $q_t = q_t^{\phi} + q_t^{\upsilon} + q_t^{\varphi} + q_t^{\eta}$ with:

$$\begin{aligned} q_t^{\phi} &= \Phi_{41}\sum_{i=1}^{t}\varepsilon_i^{\phi} + \sum_{i=1}^{t}\Phi_{i,41}^{*}\varepsilon_i^{\phi},\ q_t^{\upsilon} = \Phi_{42}\sum_{i=1}^{t}\varepsilon_i^{\upsilon} + \sum_{i=1}^{t}\Phi_{i,42}^{*}\varepsilon_i^{\upsilon}, \\ q_t^{\varphi} &= \Phi_{43}\sum_{i=1}^{t}\varepsilon_i^{\varphi} + \sum_{i=1}^{t}\Phi_{i,43}^{*}\varepsilon_i^{\varphi},\ q_t^{T} = \sum_{i=1}^{t}\Phi_{i,44}^{*}\varepsilon_i^{\eta}, \end{aligned} \tag{13}$$

respectively, where Φ_{jk} is the element in the j-th row and k-th column in Φ, and $\Phi_{i,jk}^{*}$ that in the j-th row and k-th column in the matrix Φ_i^{*} which forms the polynomial $\Phi^*(L)$ in (12).

The decomposition (13) makes it possible to assess to what extent each of the four stochastic elements included in the model contributes to explaining the evolution of the real exchange rate (and the other variables of the system) over time. Once the model has been identified, dynamic simulations (such as forecast error variance decomposition and impulse analysis) and historical decomposition can be performed.

[57] These last assumptions are consistent with other empirical studies which tried to identify productivity shocks starting form neo-classical scheme where productivity shocks are directly derived as Solow residuals (see, among others, Alexius and Carlsson, 2005)

4. Data and estimation results

4.1 Data sources and construction of variables

Quarterly observations over the period 1980q1-2006q4 are used. Data for the nominal exchange rate (E), defined as national currency per US dollar, consumer price index (P) and real GDP (Y) are from the IMF's International Financial Statistics (IFS) database (code AE…ZF, 64…ZF and 99BVP…RZF, respectively). For Argentina and Brazil these data were obtained from Datastream. Employment levels (N), measured in thousand of employees, are taken from Datastream for all countries. Finally, the shares of government expenditure in good and services (G)[58] are from Penn World Table 6.2. When quarterly observations are not available, annual data have been interpolated to create quarterly series using the Di Fonzo (1990) method. Finally, seasonal adjusting has been carried out using TRAMO/SEATS. Private output is obtained by multiplying the level of real GDP for the share of private output calculated as ($1-G$). The real exchange rate (Q) is defined as E times the ratio between US and domestic prices. Thus, an increase in Q means a real depreciation. All the variables are expressed in constant prices (base year 2000=1). Table 1 below provides further details.

[TABLE 1]

As a preliminary analysis, we performed standard ADF (Dickey and Fuller, 1979) unit root tests on (the log of) each variable. The deterministic component includes an intercept and, when statistically significant, a linear trend. The number of lags is chosen such that no residual autocorrelation is evident in the auxiliary regressions. In all cases we are unable to reject the unit root-null hypothesis at conventional nominal levels of significance. On the other hand, differencing the series appears to induce stationarity. The PP (Philips and Perron, 1988) unit root test and the KPSS (Kwiatkowski et al. 1992) stationarity test corroborate these results.[59]

[TABLE 2]

4.2 Baseline specification: VEC model estimates

The order of autoregression of the models is chosen according to the usual optimal lag length criteria (Akaike information criterion, AIC, and Bayesian information criterion, BIC), setting the maximum lag equal to eight. In the case of discordant results between the

[58] This methodology is consistent with those used, for example, in Blanchard and Perotti (2002) and Perotti (2002).

[59] Results from the PP and KPSS tests are not reported to save space, but are available from the authors upon request.

alternative criteria, we give preference to the AIC to allow for a richer system specification. The order of autoregression turns out to be two for Mexico, three for Chile, four for Argentina and Brazil, five for Peru and eight for Bolivia. System misspecification tests (not reported to save space) suggest no traces of heteroskedasticity and serial correlation.[60] Departures from normality are detected in all models. However, as pointed out by Lee and Tse (1996), the maximum likelihood approach to cointegration developed by Johansen (1995) produces testing procedures which are fairly robust to the presence of non-normality.

The number of cointegration vectors is determined on the basis of the trace test statistics of Johansen (1992). Their critical values are taken from Osterwald-Lenum (1992). Notice that the VAR specification considered here is model $H_1^*(r)$ in Johansen's notation, where a linear deterministic trend is implicitly allowed for, but this can be eliminated by the cointegrating relations so that the process contains no trend stationary components. Table 3 presents the results. The trace test suggests the presence of one cointegration relationship in all models at the 5 percent level of significance, except in the case of Bolivia where the test suggests choosing a rank of two, but a single long-run equilibrium condition at the 1 percent. These results are broadly consistent with our *a priori* theoretical assumptions about the existence of (at least) three common stochastic trends driving each system.[61]

[TABLE 3]

Structural residuals are then extracted from the reduced-form disturbances by imposing (at least) $k^2 = 16$ restrictions on the elements of matrix B. Following the identification strategy discussed in section 3, a first set of constraints is obtained by assuming that the structural shocks are orthonormal: this implies $k(k+1)/2 = 10$ (non-linear) restrictions. The choice of the cointegration rank allows to distinguish transitory shocks from permanent innovations and produces additional $r(k-r) = 3$ restrictions; in our case, there are *four* additional zero restrictions in the 4×3 matrix in (10), producing an over-identified structure, which can be tested by means of the usual χ^2-distributed likelihood ratio (LR) tests. The statistics for Argentina, Bolivia, Chile and Mexico turn out to be 1.28, 0.68, 0.37 and 1.60 respectively; by contrast, in the case of Brazil and Peru, their value is 263.20 and 140.81, respectively. By

[60] Only in the case of Bolivia are there symptoms of autocorrelation, mainly in the equations for relative productivity and relative labour services.

[61] The maximum eigenvalue test statistics indicate one cointegrating relationship only for three countries (Argentina, Bolivia and Peru), while in the other models (Brazil, Chile and Mexico) there is evidence of four common stochastic trends. In general, we favour the conclusions of the trace test in line with Johansen (1992), according to which the maximum eigenvalue test may produce a non-coherent testing strategy. Results are available on request.

comparing these test statistics to the critical values of a χ^2 distribution with one degree of freedom, we are unable to reject the null hypothesis of the validity of the over-identifying restriction only for the first four models. Accordingly, we impose the over-identified structure in the case of Argentina, Bolivia, Chile and Mexico, while for the Brazilian and Peruvian systems we employ a just-identified structure.

5. Empirical evidence

Once structural and data-consistent identification of the VEC models is achieved, dynamic simulations (forecast error variance decomposition and impulse response analysis) as well as historical decomposition exercises can be performed. We employ these techniques in order to address three main issues: *first*, we assess the role played by the underlying (structural) sources in explaining the fluctuations of the variables in each country model, also discriminating between supply and demand disturbances (Section 4.1); *second*, we study the sign and the magnitude of the response of the real exchange rate to an unanticipated fiscal shock (Section 4.2); *third*, we analyse the contribution of each structural shock in driving real exchange dynamics over the sample period under investigation (Section 4.3).

5.1 Sources of system-wide and variable fluctuations

Several studies have analysed the role of demand shocks (for instance, monetary and fiscal policies) and supply disturbances (productivity and labour supply shocks or structural restructuring policies, such as tariff and trades reforms) in a closed-economy context, both at the aggregate (Blanchard and Quah, 1989; Gali, 1999, among others) and, more recently, at the disaggregate level (Chang and Hong, 2005; Busato et al., 2005) for the US economy or other developed countries. We assess the relative contribution of the structural shocks in explaining macroeconomic fluctuations by means of forecast error variance decomposition analysis. Such a tool makes it possible to summarise the information contained in the structural MA representation (11) and provides a convenient measure of the relative importance of each shock to forecast error variance as a function of the simulation horizon. Table 4 presents the share of variability (in percentage terms) which can be attributed to each structural shock for the individual variables of the model as well as for the system as a whole (row labelled as "system") over a simulation horizon of 20 quarters. Aggregating the shocks, we consider supply shocks (ϕ and υ) and demand disturbances (φ and η).

[TABLE 4]

The results are interesting in a number of respects. *First*, supply shocks are the most relevant source of macroeconomic fluctuations in all systems. Their contribution ranges from more than 70 percent in Argentina, Brazil and Mexico to around 60 percent in Bolivia. This finding is broadly consistent with the empirical evidence for developed economies.[62] *Second*, a closer look at the contribution of structural disturbances to fluctuations of individual variables shows the existence of three distinct groups of countries. The results for Argentina, Bolivia and Mexico reveal that productivity shocks are the main driving forces of relative productivity and relative private output variability, while relative labour services and the real exchange rate fluctuations are mainly governed by labour input and fiscal shocks, respectively. By contrast, while fiscal shocks still represent the main driving forces of the variability of international competitiveness in the Chilean economy, relative preference (labour input) disturbances turn out to drive variability in the dynamics of relative private output (productivity and labour services). Finally, in just-identified structures (Brazil and Peru), we observe that relative productivity and relative labour services fluctuations originate from productivity shocks, with labour input and fiscal shocks dominating the variability of private output and real exchange rates changes. *Third*, focusing on the main variable of interest in our analysis (the real exchange rate) we find evidence of a difference in behaviour between over-identified and exactly identified systems: in the former class of models international competitiveness is driven by the demand-side of the economy, whilst in the latter group of countries the real exchange rate responds mostly to supply-side disturbances. Further, fiscal shocks are the main driving force of real exchange rate movements in the majority of cases (Argentina, Bolivia and Mexico), ranging from 60 to 90 percent, while they are less relevant for Chile and Peru, even though their effects are still sizeable (35 and 21 percent, respectively). Only in the case of Brazil is the contribution of this shock negligible.

5.2 Fiscal shocks and international competitiveness of the LA countries

The estimated models can be used to assess the effects of mutually orthogonal one-off structural shock on the dynamics of individual variables by calculating the impulse response functions (IRFs). In particular, we focus on the effects of an unanticipated relative fiscal shock on the international competitiveness of the LA countries. Tervala (2006) points out that, as government spending may exhibit different degrees of productivity, if this is low or

[62] Bergman (1996), for instance, using a bivariate VAR model for output and inflation, shows that more than one half of the macroeconomic fluctuations in the G7 countries are due to supply shocks at the typical business cycle frequency (the twenty-quarter horizon).

zero, a rise in government spending causes a fall in domestic consumption, leading to a relative consumption change, which decreases the relative demand for domestic money, and consequently the real exchange rate depreciates. These predictions are coherent with the conclusions in Obstfeld and Rogoff (1995). On the other hand, if the productivity of government spending is sufficiently high, domestic consumption rises more than foreign consumption. Hence, the nominal and real exchange rate appreciate because the relative consumption change increases the relative demand for domestic money in a way consistent with the implications of Keynesian-style two-country models of the world economy.[63]

Figure 1 shows the response of the real exchange rate (solid line) to a positive relative fiscal shock for each of the six LA countries. 95 percent confidence bounds (dashed lines), generated by Monte Carlo with 5000 replications, are also reported. The simulation horizon for IRFs is set equal to 20 quarters.

[FIGURE 1]

As can be seen, the Argentine real exchange rate appreciates after a relative fiscal shock. The loss in international competitiveness is consistent with the conclusions of Roldós (1995), according to whom public spending shocks can lead to real exchange rate appreciation. Since confidence bounds include the baseline path (the horizontal axis), deviations from the pre-shock level cannot be judged to be statistically significant at the chosen significance level in the fifth year of the simulation horizon. By contrast, in the models of Bolivia, Chile and Mexico we observe that fiscal shocks lead to real depreciation, albeit the deviations from the steady-state level for the Bolivian case appear to be statistically significant only in the first five quarters. Following the lines of reasoning of Tervala (2006), this finding may suggest low productivity of government spending policies in these two countries. The response of the Brazilian real exchange rate appears not to be statistically significant, consistently with the evidence discussed in the previous sub-Section. Finally, in the case of Peru we observe a statistically significant but short-lived depreciation of the real exchange rate.

In brief, the results from the IRFs suggest a close relationship between relevance of fiscal shocks as a driving source for real exchange rate fluctuations and effects of unanticipated fiscal shocks on the level of international competitiveness in the LA economies. The sign of the response of international competitiveness to this type of shock cannot as clearly be determined *ex-ante*. The evidence reported here indicates a substantial inelasticity of

[63] The international transmission of fiscal policy shocks in micro-founded general equilibrium models crucially depends on assumptions related to whether the fiscal shock is permanent or temporary, whether international asset markets are complete or not (Baxter, 1992), whether labor supply is fixed or variable, and how government purchases (Bianconi and Turnovsky, 1997) are financed. On this topic, see Arin and Koray (2008).

international competitiveness to shocks originating from public spending policies in countries where fiscal shocks have scant role in explaining real exchange variability. On the other hand, fiscal shocks may induce either real exchange rate appreciation (for Argentina) or depreciation (for Bolivia, Mexico and Chile) in the real exchange rate.

5.3 Explaining real exchange dynamics in LA countries over the years 1980-2006

This sub-Section describes how to assess the observed real exchange rates patterns for our six LA countries in the light of historical shifts in their fundamentals. Indeed, the existence of a stable long-run relationship among the variables of each model does not prevent the relative weight of those factors from changing over time in response to complex and interrelated reciprocal influences. Hence, it could be instructive to examine the hypothetical time path of international competitiveness if all disturbances had been associated to only one source of shock.

Table 5 summarises the relevance of each structural component in explaining international competitiveness variability over time. OLS estimates are obtained by regressing changes in the real exchange rate on its component driven by individual orthogonal shocks according to the decomposition (12). Since structural components are mutually orthogonal by construction, the total variation of the regressand (changes in international competitiveness) must be *fully* captured by the explanatory variables (supply shocks, ϕ and υ, and demand disturbances, φ and η).

[TABLE 5]

The results indicate that the in-sample variability of the real exchange rate is dominated by demand shocks in most of the models, with percentages ranging from 39 percent in the case of Chile to 92 percent in that of Bolivia. In particular, for five out of the six countries (Brazil being the only exception), fiscal shocks account for a considerable percentage of real exchange rate movements, ranging from one-fifth (for Peru) to four-fifth (for Mexico) of total variability. Also, notice that in most cases (Argentina, Bolivia, Mexico and Peru) the effects of fiscal impulses are stronger than those of productivity shocks. Finally, the relative importance of the temporary components (namely, preference shocks) varies across countries, being at its highest in Brazil, where it explains 43 percent of the historical variance (the effects of fiscal shocks being negligible), and in Peru, where the corresponding share is 34 percent, whilst is countries such as Mexico and Chile it is as low as 6.13 and 6.47 respectively.

In order to check for possible shifts in the relative explanatory contribution of shocks for real exchange rate changes over the sample span, we use a recursive method, i.e. we employ the estimated models to replicate the previous exercise over the window embracing the period from the first available observation to 1994*q*4 and then extending it by a datapoint at a time. Summary statistics (mean, standard error of the mean, minimum and maximum values) for each system are reported in Table 6.

[TABLE 6]

The results broadly confirm the previous evidence in a number of ways. *First*, fiscal shocks are the most relevant source of variation for real exchange rates in the over-identified models. *Second*, in all models, the mean values of each shock resulting from the recursiveing procedure are *quantitatively* very close to their full-sample counterparts and *qualitatively* similar to the results from the forecast error variance decomposition exercise. *Third*, the standard error of the mean, as well as the minimum-maximum range, suggest that the relative contribution of the four driving forces in explaining real exchange rate changes are almost constant over time.

The last piece of evidence concerns the relationship between structural shocks and the pattern over time of the level of international competitiveness. Figure 1 shows, for each country, the real exchange rate series purged of the deterministic part (solid line), and its component explained by the fiscal shocks (dashed line).

[FIGURE 2]

Visual inspection suggests the following. It seems that the effects of fiscal shocks in the period 1981-1986 are considerable for all the countries under examination. After this period, however, this is still the case only in the models of Bolivia and Mexico, while in Chile and Peru long swings in the real exchange rate are only partially caused by the fiscal components. Consistently with the previous results, fiscal shocks do not appear to have significant explanatory power for real exchange rate movements in Brazil.

Combining this evidence with the results of the IRF analysis, we can conclude that in the cases where the real exchange rate depreciates in response to a relative fiscal shock (namely, Bolivia, Chile, Mexico and Peru), the explanatory power of these structural shocks is more pronounced over the entire investigation period. In the case of Argentina, where the IRF analysis shows the opposite result (that is, a real appreciation in response to a relative fiscal shock), the time path of the real exchange rate is less influenced by the component caused by this shock. According to Tervala (2006), the different time paths followed by the component of real exchange rates driven by fiscal shocks can be explained by the different degree of productivity of government expenditure in each country.[64]

6. Robustness analysis

6.1 Alternative specifications of the baseline model

The results from structural VAR models relying on long-run restrictions may vary considerably depending on the exact specification of the empirical model. As argued by Faust and Leeper (1997), identification procedures, which involve restrictions on the long-run effects of structural shocks, may imply that type-II errors are more likely in confidence intervals because of the imprecision of the long-run parameter estimates. Therefore, in this section we study the robustness of the results discussed above with respect to changes in the empirical specification of the systems.

Three alternative empirical specifications are estimated in order to investigate how the relative weights of demand shocks (and in particular fiscal shocks) vary with the nature of the fluctuations. We filter the data by different methods, namely first differences, FD, the HP filter (Hodrick and Prescott, 1997), HP, and linear detrending, LD. In particular, FD series are used to isolate short cycle fluctuations, HP-filtered series for intermediate frequencies and LD series for low frequencies. We expect the role of demand shocks to decrease with the persistence of shocks.[65] Notice that all alternative specifications neglect the existence of possible cointegration relationships. Thus, our robustness checks can shed light on the consequences of ignoring the presence of long-run equilibrium relationships between the variables.

Table 7 presents the results from imposing the over-identifying long-run restriction in the three alternative empirical specifications. p-values are in square brackets.

[TABLE 7]

Overall, the long-run structure implied by our theoretical relationship of reference is not rejected by the data in nine (one at the 1 percent, one at the 5 percent and the remaining seven at the 10 percent level of significance) out of eighteen cases. In particular, the outcome from the FD specification is fully consistent with the baseline design, even though the test

[64] As shown by Rodriguez and Romero (2007), in Argentina the hyperinflation phenomenon which took place at the beginning of the 1990 and the following abolition of the currency board explain the dominant effect of the variability of transitory components on the behaviour of the real exchange rate, while in the case of Brazil (where the initially floating exchange rate was subsequently fixed) international competitiveness has mainly been driven by real shocks in the last decade.

[65] The FD specification is the baseline model (10) with $\Pi = 0$ and with four common trends. Such a specification is consistent with the conclusions of the maximum eigenvalue test for Brazil, Chile and Mexico.

statistics are slightly less supportive of our economic priors. In the present context, this conclusion is not surprising since the FD specification produces loss of relevant information, in the presence of documented cointegration relationships. Notice, further, that in the LD specification we observe the rejection of the null hypothesis in all models but one (the Chilean case).

Following the same criterion as in the previous Section, we perform a forecast error variance decomposition under the over-identified structure for the specifications where the over-identifying restriction holds, but employing the just-identified structure when the constraint imposed on the long-run matrix is rejected by the data. The simulation horizon is set equal to 20 quarters. Table 8 shows the contribution (in percentage terms) of aggregate demand shocks and fiscal shocks to the overall forecast error variance of the real exchange rate under the three alternative empirical specifications.

[TABLE 8]

As expected, in most cases the relative importance of demand shocks is stronger in the specification where the short-run cycle frequency, FD, is isolated, and decreases when a longer cyclical component is taken into account, that is when we move from the HP to the LD specification.

Comparing these results to those from the VEC models, we observe that the explanatory power of demand shocks under the alternative specifications is greater than their counterparts from the baseline specification with cointegration. Moreover, focusing on the individual demand shocks, there is evidence of a bigger role for relative preference shocks, these now becoming the most important source for real exchange rate fluctuations.

As shown by Alexius (2005), the lack of the long run-equilibrium conditions between fundamental variables and the real exchange rate eliminates the relationship between the latter and productivity disturbances. Thus, the relative system impact of supply disturbances tends to decrease. In addition, if the long-run properties of the system are not properly taken into account, the effects of fiscal shocks are underestimated, as the relationship between government size and the dynamics of the real exchange rate is overlooked.

6.2 Alternative specification of the private output and fiscal shocks

As discussed by Ramey (2008), there are two main approaches to the identification of fiscal shocks. The first one is mainly based on VAR technique, in which the fiscal shocks are identified through a structural representation where the causal order of the variables allows to determinate the effects of an unexpected movements in the government size (Blanchard and Perotti, 2002).

In contrast, Ramey and Shapiro (1998) use a narrative approach to identify shocks to government spending. They argue that many shocks identified from a VAR are simply anticipated changes in government spending, thus they focus only on few episodes caused by important political events unrelated to the state of the economy.

Our approach clearly refers to the first one, even if we do not consider explicitly (in equation 4 and 5) the permanent increase in taxes which is implied by increases in government size determined by fiscal shocks. Blanchard and Perotti (2002) using a VAR model with three variables: taxes, government expenditures and GDP, argue that the identification of structural shocks between taxes and GDP is particularly hard because of the presence of contemporaneous relationships between variables. They need to incorporate institutional information on taxes and spending in order to achieve a robust identification of fiscal shocks. Thus, as Kneller, Bleaney, Gemmell (1999) point out, empirical studies that include only government expenditure and no taxes may be mis-specified.

However, in order to partially overcome these problems, we implicitly assumed in our, as in Ahmed et al. (1993), that changes in taxation have no effects on total output in the long-run. This can happen if wealth and the substitution effects of higher taxes on labour supply cancel out, as in a Real Business Cycle (RBC) model with a Cobb-Douglas production function, constant relative risk aversion preferences and where the level of taxes is proportional to output.

In particular, our identification of fiscal shocks comes directly out by the definition of private output expressed in equation (4), where it is calculated as the difference between actual output and government spending[66]. Since this measure of private output could be misleading, in this section we try to specify in better way the fiscal shocks including in the private output equation the investments and net exports. Data for the ratio between capital investments over GDP, as well as net exports over GDP are from World Bank's *WDI indicators.* Annual data have been interpolated to create quarterly series using the Chow and Lin (1971). We re-estimate the six models using the same specification of the baseline scheme, allowing for one cointegration vector in all systems. Table 9 presents the share of variability (in percentage terms) which can be attributed to each structural shock for the real exchange rate of each Latin America country over a simulation horizon of 20 quarters.

[66] The definition of private output adopted in the model is consistent with other papers that study the international transmission of structural shocks. See Ahmed et al. (1993) and Alexius (2005), among others.

[TABLE 9]

As in the previous tables we aggregated the shocks, considering both supply shocks (ϕ and υ) and demand disturbances (φ and η). With the new specification of private output, fiscal shocks are the main driving force of real exchange rate variability not only for Argentina, Bolivia and Mexico (as in the baseline estimations), but also for Chile and Peru. Furthermore, it is possible to note that also in the case of Mexico the average contribution of fiscal shocks is bigger, even if supply shocks seem to be always the most important source of real exchange rate movements. Regarding the differences between supply and demand shock, the latter plays a bigger role in explaining the variability of exchange rate than in the previous results, according to the increasing impact of fiscal shocks detected in this last specification.

Also in this case, we perform robustness analysis using the same alternative empirical specifications: FD, the HP filter and linear detrending, LD. The results of the simulation exercises are presented in table 10.

[TABLE 10]

The results seem to confirm the fact that the relative importance of demand shocks in explaining real exchange rate variability is stronger in the FD specification (short-run cycle frequency), while it decreasing for HP and LD ones (where longer cyclical component are is isolated). However, in most of the cases, there is evidence of a small role for fiscal shocks, even smaller than that detected in the robustness analysis of the baseline model shown in table 8. Once again, this last result can be confirm the importance of using a model with long run equilibrium conditions for the fiscal shocks.

Overall, the new specification of private output seems to confirm all the main previous results, and in many cases these become stronger if you look for the magnitude of the detected effects.

6. Conclusions

This paper adopts a modelling approach aimed at assessing the role of a wide class of underlying (structural) disturbances in driving real exchange rates (defined relative to the US dollar) in six LA countries (Argentina, Bolivia, Brazil, Chile, Mexico and Peru), along the lines of the studies of Ahmed et al. (1993) and Hoffmaister and Roldós (2001). These disturbances are identified as relative productivity, labour, fiscal and preference shocks.

Using quarterly data over the period 1980-2006, we analyse the case of the LA economies, for which the effects of fiscal shocks on the real exchange rate had not previously been studied. Specifically, we show that fiscal shocks are a key determinant of changes in international competitiveness for most of the countries we consider. Our approach sheds new light on the driving forces of real exchange rate dynamics in developing economies. A simpler modelling strategy, relying exclusively on a standard permanent/transitory decomposition, would provide only partial evidence, as, by construction, it would allow for only two types of shocks, ignoring the possibility of a wider class of disturbances hitting the economy as a whole (and consequently the real exchange rate as well) that also need to be investigated.

Therefore, our contribution to the literature on fiscal shocks is two-fold. *First*, we identify fiscal shocks in a multicountry/multivariate time series context, allowing for the existence of possible cointegration relationship among the variables of the system. *Second*, we present some new empirical evidence for six Latin American countries. We find that the effects of unanticipated fiscal impulses on the level of the real exchange rate vary, reflecting different degrees of government expenditures productivity. Further, using alternative econometric specifications, we show how the importance of fiscal shocks (and more in general of demand shocks) on the variability of the international competitiveness varies with the frequency of cyclical fluctuations isolated in the models. The explanatory power of demand shocks increases when shorter cyclical fluctuations are taken into account. Moreover, neglecting the presence of cointegration, which in fact holds in our case, amounts to overlooking the linkage between productivity and government spending and the real exchange rate. As we show, this leads to overestimating the role of demand shocks and underestimating the contribution of fiscal disturbances, putting into question the reliability of earlier evidence for which this criticism is relevant (see, e.g. Ahmed et al., 1993; Chowdhury, 2004; Hoffmaister and Roldós, 2001; Rodríguez and Romero, 2007). The main results of the paper are robust when different specification of private output is used allowing for the presence of investments and net exports.

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Tables and figures

Table 1 – Construction of the variables

Variable	*Definition*
Relative productivity	$\pi_t^i = (\ln Y_t^i - \ln N_t^i) - (\ln Y_t^{US} - \ln N_t^{US})$
Relative employment	$n_t^i = \ln N_t^i - \ln N_t^{US}$
Relative private output	$z_t^i = \ln[Y_t^i(1-G_t^i)] - \ln[Y_t^{US}(1-G_t^{US})]$
Real exchange rate	$q_t^i = \ln E_t^i + (\ln P_t^{US} - \ln P_t^i)$

Note. For each variable the suffix *i* refers to each Latin America country in turn, while the suffix *US* refers to the base country (the US economy). The subscript *t* stands for time.

Table 2 – Unit root tests

	π				n				z				q			
	Levels		First differrences		Levels		First differrences		Levels		First differrences		Levels		First differrences	
	DP	TS	DP	TS	DP	TS	DP	TS	DP	TS	DP	TS	DP	TS	DP	TS
Argentina	c,t	-2.16	c	-9.92	c,t	-2.42	c	-4.35	c	-1.94	.	-4.13	c,t	-3.79	c	-3.44
Bolivia	c,t	-2.08	c	-6.45	c,t	-2.01	c	-4.74	c,t	-0.32	c	-1.87	c	-1.46	.	-4.96
Brazil	c	-1.66	.	-6.44	c,t	-2.18	c	-5.19	c	-1.73	.	-3.70	c	-1.24	.	-4.63
Chile	c,t	-2.05	c	-7.94	c,t	-1.76	c	-6.59	c,t	-2.86	c	-3.63	c,t	-1.16	c	-4.45
Mexico	c,t	-1.76	c	-5.73	c,t	-2.06	c	-4.22	c	-1.42	.	-3.95	c	-2.33	.	-11.7
Peru	c,t	-1.52	c	-6.64	c,t	-0.03	c	-5.43	c	-1.49	.	-2.70	c	-1.95	.	-4.74

Note. ADF test statistics for the null hypothesis of a unit root process for the variables in the levels and in first differences are reported in columns "TS". The critical value at the 1 percent level of significance is -4.05 if a constant and a linear trend (*c,t*) are included in the regression, -3.49 with only a constant term (*c)* and -2.59 if no deterministic parts (-) are included. At the 5 percent level of significance these values are -3.45, -2.89 and -1.94, respectively (MacKinnon, 1996). The specification of the deterministic component is presented in the column "DP". Definitions of the variables are provided in Table 1.

Table 3 – Cointegration rank

	Lags	Rank			
		0	1	2	3
Argentina	4	66.77	28.53	11.03	1.87
Bolivia	8	69.77	31.16	11.77	0.42
Brazil	4	47.31	22.6	8.3	0.47
Chile	3	48.89	26.05	11.11	2.66
Mexico	2	49.48	27.07	11.47	0.63
Peru	5	56.06	28.68	4.64	0.74

Note. Critical values for the trace test statistics at the 95 percent for rank 0, 1, 2, 3 and 4 are 47.21, 29.68, 15.41 and 3.76, respectively, while at the 99 percent are 54.46, 35.65, 20.04 and 6.65, respectively (Osterwald-Lenum, 1992). The column "Lag" reports the number of lags included in the VAR specification suggested by the AIC.

Table 4 – Forecast error variance decomposition

	Individual shocks				*Nature of shocks*	
			Argentina			
	ϕ	υ	φ	η	*Supply*	*Demand*
$\Delta\pi$	82.80	1.68	11.87	3.65	*84.48*	*15.52*
Δn	0.96	97.46	0.37	1.21	*98.42*	*1.58*
Δz	87.24	11.96	0.69	0.11	*99.20*	*0.80*
Δq	16.13	8.99	60.69	14.19	*25.12*	*74.88*
System	*46.78*	*30.02*	*18.41*	*4.79*	*76.81*	*23.19*
			Bolivia			
	ϕ	υ	φ	η	*Supply*	*Demand*
$\Delta\pi$	50.58	5.43	34.82	9.17	*56.01*	*43.99*
Δn	7.61	91.93	0.12	0.34	*99.54*	*0.46*
Δz	70.68	6.14	19.16	4.02	*76.82*	*23.18*
Δq	3.32	0.71	90.77	5.20	*4.03*	*95.97*
System	*33.05*	*26.05*	*36.22*	*4.68*	*59.10*	*40.90*
			Brazil			
	ϕ	υ	φ	η	*Supply*	*Demand*
$\Delta\pi$	90.77	4.67	0.73	3.83	*95.44*	*4.56*
Δn	67.84	26.20	2.92	3.04	*94.04*	*5.96*
Δz	6.90	17.83	69.03	6.24	*24.73*	*75.27*
Δq	22.73	61.64	0.48	15.15	*84.37*	*15.63*
System	*47.06*	*27.59*	*18.29*	*7.07*	*74.65*	*25.36*
			Chile			
	ϕ	υ	φ	η	*Supply*	*Demand*
$\Delta\pi$	20.06	47.70	6.02	26.22	*67.76*	*32.24*
Δn	11.56	82.91	0.28	5.25	*94.47*	*5.53*
Δz	43.83	6.77	1.69	47.71	*50.60*	*49.40*
Δq	23.76	27.85	34.82	13.57	*51.61*	*48.39*
System	*24.80*	*41.31*	*10.70*	*23.19*	*66.11*	*33.89*
			Mexico			
	ϕ	υ	φ	η	*Supply*	*Demand*
$\Delta\pi$	86.20	0.65	0.97	12.18	*86.85*	*13.15*
Δn	2.49	95.96	0.07	1.48	*98.45*	*1.55*
Δz	70.92	25.75	2.52	0.81	*96.67*	*3.33*
Δq	17.52	1.63	78.35	2.50	*19.15*	*80.85*
System	*44.28*	*31.00*	*20.48*	*4.24*	*75.28*	*24.72*
			Peru			
	ϕ	υ	φ	η	*Supply*	*Demand*
$\Delta\pi$	67.42	22.32	2.86	7.40	*89.74*	*10.26*
Δn	67.32	25.31	3.78	3.59	*92.63*	*7.37*
Δz	7.35	13.03	64.80	14.82	*20.38*	*79.62*
Δq	23.35	39.07	21.05	16.53	*62.42*	*37.58*
System	*41.36*	*24.93*	*23.12*	*10.59*	*66.29*	*33.71*

Note. Average percentage contribution of each structural shock in explaining variable fluctuations over a simulation horizon of 20 quarters. ϕ, υ, φ, η indicate relative productivity, relative labour, relative fiscal and relative preference shocks, respectively. The column "Supply" is the aggregate contribution of ϕ and υ disturbances. The column "Demand" is the aggregate contribution of φ and η disturbances. The row "System" indicates the average contribution of individual shocks and aggregate disturbances, disentangled according to their nature, for the whole system.

Table 5 – Historical decomposition

	Individual shocks				*Nature of shocks*	
	ϕ	υ	φ	η	*Supply*	*Demand*
Argentina	19.07	24.65	35.35	20.93	*43.72*	*56.28*
Bolivia	5.21	2.90	73.36	18.53	*8.11*	*91.89*
Brazil	15.51	39.87	1.25	43.37	*55.38*	*44.62*
Chile	39.11	21.89	32.53	6.47	*61.00*	*39.00*
Mexico	10.98	2.04	80.85	6.13	*13.02*	*86.98*
Peru	17.82	27.11	21.02	34.05	*44.93*	*55.07*

Note. Percentage contribution of each structural shock in explaining the historical variance of the real exchange rate quarterly changes. ϕ, υ, φ, η indicate relative productivity, relative labour, relative fiscal and relative preference shocks, respectively. The column "Supply" is the aggregate contribution of ϕ and υ disturbances. The column "Demand" is the aggregate contribution of φ and η disturbances.

Table 6 – Historical decomposition – Recursive Method

	ϕ	υ	φ	η
	Argentina			
	ϕ	υ	φ	η
Mean	19.07	22.24	38.93	19.76
Std. err. of mean	0.05	0.22	0.33	0.14
Minimum	18.23	19.65	34.92	17.80
Maximum	20.08	25.16	42.27	21.71
	Bolivia			
	ϕ	υ	φ	η
Mean	2.59	2.53	79.07	15.81
Std. err. of mean	0.11	0.03	0.36	0.24
Minimum	1.98	2.08	73.36	12.33
Maximum	5.21	2.93	83.57	18.88
	Brazil			
	ϕ	υ	φ	η
Mean	17.65	38.03	1.09	43.22
Std. err. of mean	0.16	0.17	0.01	0.17
Minimum	15.44	35.14	0.84	40.22
Maximum	20.18	41.18	1.25	46.73
	Chile			
	ϕ	υ	φ	η
Mean	39.98	22.03	31.05	6.94
Std. err. of mean	0.14	0.10	0.13	0.06
Minimum	38.04	20.77	28.73	6.10
Maximum	42.51	23.70	32.87	7.65
	Mexico			
	ϕ	υ	φ	η
Mean	10.32	1.87	82.70	5.11
Std. err. of mean	0.03	0.02	0.12	0.08
Minimum	9.87	1.66	80.85	3.93
Maximum	10.98	2.17	84.53	6.13
	Περυ			
	ϕ	υ	φ	η
Mean	22.29	30.68	14.37	32.66
Std. err. of mean	0.29	0.31	0.45	0.17
Minimum	17.82	26.24	11.20	30.49
Maximum	24.31	33.18	21.31	35.56

Note. Percentage contribution of each structural shock in explaining the historical variance of the real exchange rate quarterly changes. $\phi, \upsilon, \varphi, \eta$ indicate relative productivity, relative labour, relative fiscal and relative preference shocks, respectively. Summary statistics computed over simulation windows of increasing size, extended by a datapoint at a time, are reported by rows. All windows start with the first available observation, but they have different ending quarters. The smallest window covers the period up to 1994*q*4, while the largest window embraces the entire sample span.

Table 7 – Robustness Analysis - Model specification

	Model specification		
	FD	*HP*	*LD*
Argentina	[0.01]	[0.24]	[0.00]
Bolivia	[0.30]	[0.00]	[0.00]
Brazil	[0.00]	[0.00]	[0.00]
Chile	[0.09]	[0.54]	[0.45]
Mexico	[0.10]	[0.41]	[0.00]
Peru	[0.00]	[0.65]	[0.00]

Note. p-values from a χ^2-distributed LR over-identifying test with one degree of freedom are reported in squared brackets. FD, HP and LD indicates first differences, HP (Hodrick and Prescott, 1997) and linear detrending filters, respectively.

Table 8 – Robustness Analysis - Forecast error variance decompositions

	Model specification					
	FD		*HP*		*LD*	
	Demand shocks	φ	*Demand shocks*	φ	*Demand shocks*	φ
Argentina	*79.31*	2.20	*83.02*	6.62	*43.17*	14.22
Bolivia	*94.48*	21.24	*82.46*	12.26	*89.05*	7.64
Brazil	*85.26*	30.37	*42.64*	14.89	*60.52*	25.66
Chile	*79.85*	5.13	*82.16*	30.23	*45.01*	3.05
Mexico	*91.51*	7.61	*63.33*	4.18	*17.29*	2.06
Peru	*89.23*	5.52	*45.60*	0.40	*27.78*	2.06

Note. Average percentage contribution of demand and relative fiscal shocks (φ) in explaining real exchange rate fluctuations at different cyclical frequencies over a simulation horizon of 20 quarters. FD, HP and LD indicates first differences, HP (Hodrick and Prescott, 1997) and linear detrending filters, respectively.

Table 9 – Robustness Analysis(1) – FEVD using an alternative specification of private outuput

	Individual shocks				*Nature of shocks*	
	ϕ	υ	φ	η	*Supply*	*Demand*
Argentina	1.70	3.85	71.95	22.25	5.55	94.20
Bolivia	4.85	0.85	91.90	2.50	5.70	94.40
Brazil	25.40	29.50	35.25	9.85	54.90	45.10
Chile	17.75	0.60	79.30	2.30	18.35	81.60
Mexico	24.20	34.80	40.90	0.00	59.00	40.90
Peru	4.00	17.75	76.15	2.00	21.75	78.15

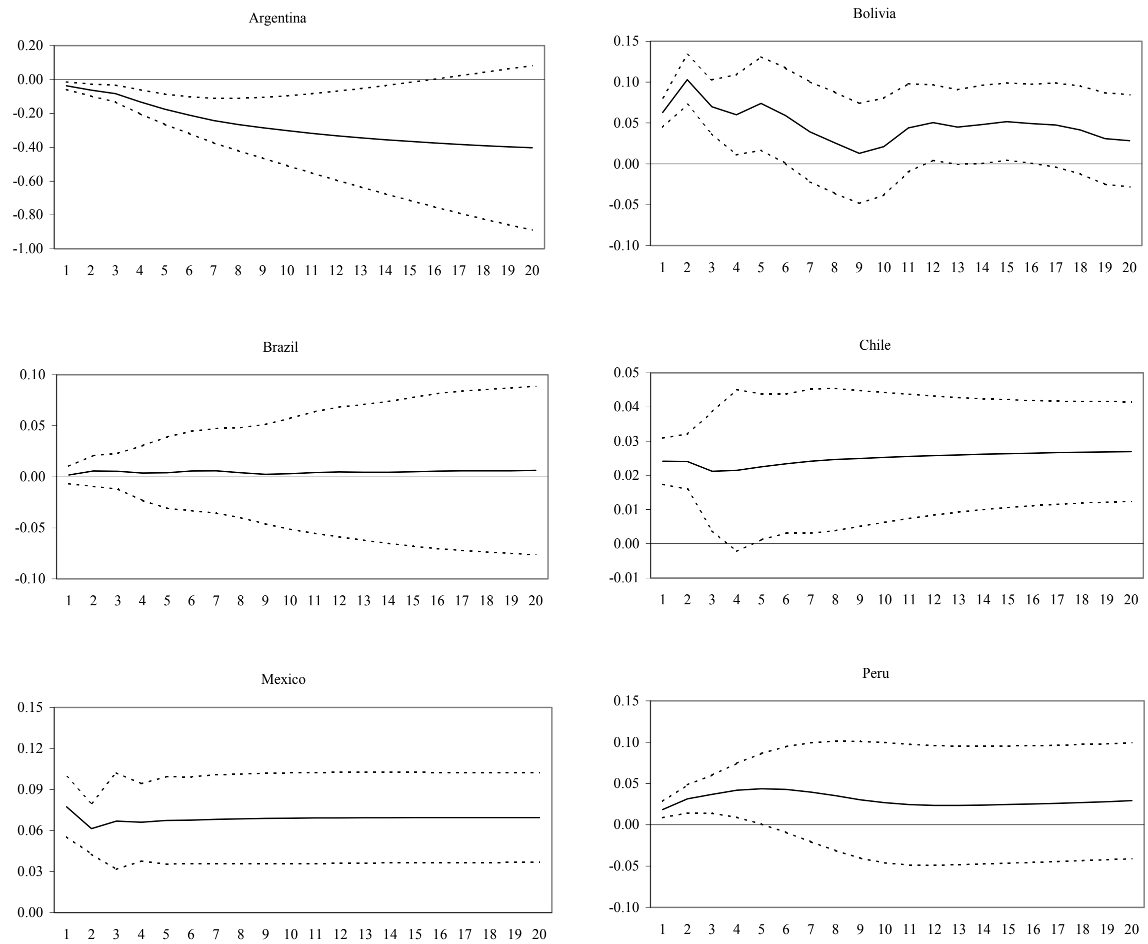
Note. Percentage contribution of each structural shock in explaining the forecast error variance decomposition of the real exchange rate. ϕ, υ, φ, η indicate relative productivity, relative labour, relative fiscal and relative preference shocks, respectively. The column "Supply" is the aggregate contribution of ϕ and υ disturbances. The column "Demand" is the aggregate contribution of φ and η disturbances.

Table 10 – Robustness Analysis(2) – FEVD using an alternative specification of private outuput

	Model specification					
	FD		*HP*		*LD*	
	Demand shocks	φ	*Demand shocks*	φ	*Demand shocks*	φ
Argentina	*82.01*	21.27	*80.14*	3.24	*54.36*	8.83
Bolivia	*89.96*	8.76	*80.61*	7.03	*72.28*	16.24
Brazil	*73.07*	2.29	*87.66*	13.01	*58.81*	22.72
Chile	*79.68*	2.54	*80.71*	4.96	*51.18*	3.66
Mexico	*83.61*	0.02	*56.58*	25.74	*9.62*	4.42
Peru	*87.16*	4.37	*45.07*	0.43	*31.81*	0.21

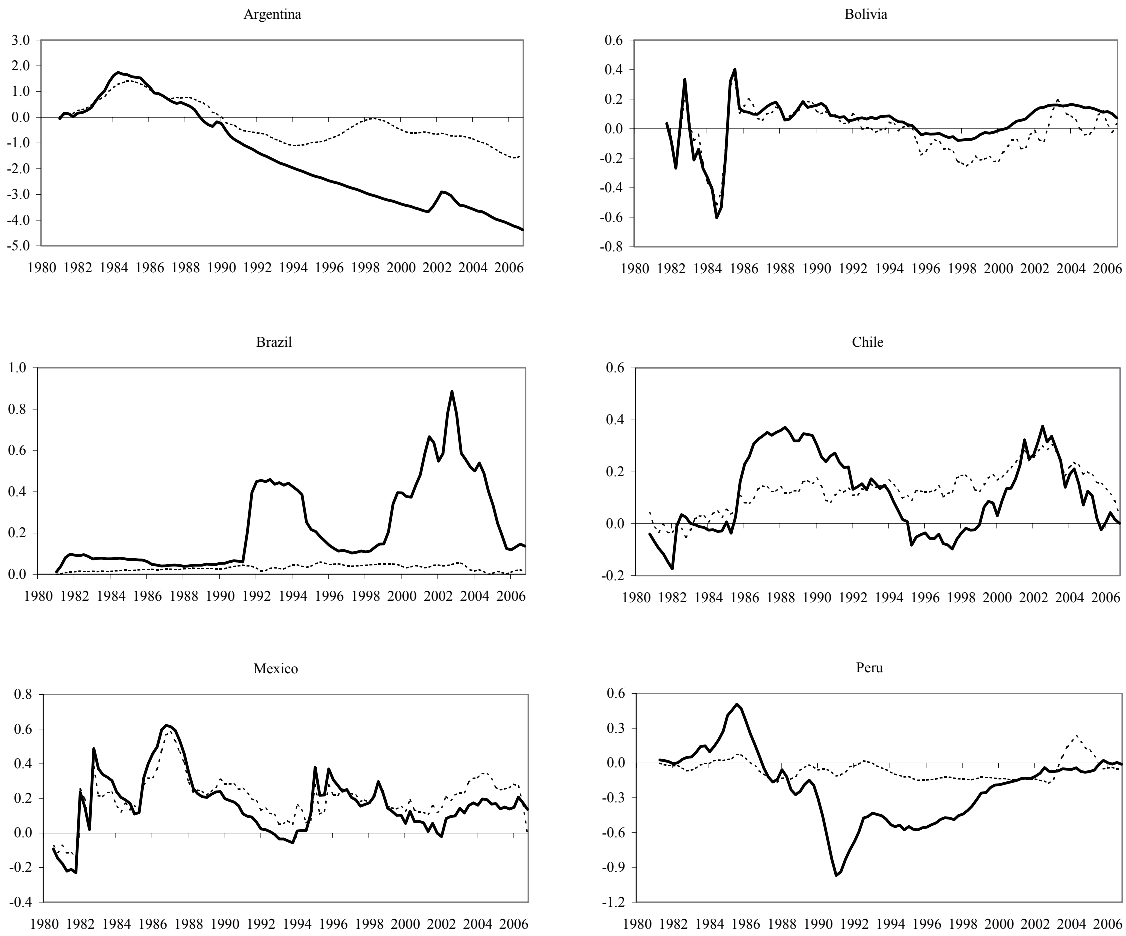
Note. Average percentage contribution of demand and relative fiscal shocks (φ) in explaining real exchange rate fluctuations at different cyclical frequencies over a simulation horizon of 20 quarters. FD, HP and LD indicates first differences, HP (Hodrick and Prescott, 1997) and linear detrending filters, respectively.

Figure 1 – Response of the real exchange rate to a fiscal shock



Note. Response of the real exchange rate to a relative fiscal shock. The vertical axis denotes changes from the pre-shock level (solid lines). The horizontal axis indicates quarters after the shocks. Confidence bounds (dashed lines) are generated by Monte Carlo with 5000 replications.

Figure 2 – Real exchange rate dynamics and the component driven by the fiscal shock



Note. In each graph, the solid line indicates the real exchange rate, while the dashed line plots its component driven by the fiscal shock.