Real Exchange Rate Volatility: A Measure of Real Convergence in the Central and Eastern European Euro Area Accession Countries[1]

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Abstract

This paper sets out an analysis of the degree of real convergence between Central and Eastern European Countries and the selected euro area Members, in terms of real exchange rate volatility. It uses univariate variance analysis (GARCH) and a structural VAR methodology with Blanchard and Quah (1989) decomposition to quantify real exchange rate volatility in the selective New Member States of the European Union, and to distinguish between the real and nominal components of real exchange rate movements. The SVAR technique is also used to assess the role of the nominal exchange rate in these New Member States in accommodating real asymmetric shocks. The results indicate that: (i) real asymmetric shocks are significant when compared with those experienced by the poorer Old Member States of the European Union in their accession to the eurozone; (ii) nominal exchange rates, in general, do play a stabilising role in the New Member States; and that (iii) nominal shocks, on average, do not move real exchange rates. Therefore, based on the analysis conducted in this paper, it appears that among the New Member States, at present, only Estonia and Slovenia are ready to give up monetary and exchange rate independence.

Key Words: Real Exchange Rate Volatility, Convergence, European Monetary Integration, Structural Vector Autoregression, Heteroskedasticity, Structural Break, Small-Sample Confidence Intervals.

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PART 1: INTRODUCTION

In 2004 eight Central and Eastern European Countries acceded to the European Union and at the same time became active members of the third stage of the Economic and Monetary Union (EMU)[3]. By doing so, they committed to participate in the Exchange Rate Mechanism II (ERMII), and eventually adopt a common European currency, the euro. The basis on which these countries time their accession to the ERMII and adopt the euro is of considerable policy importance, and the focus of this paper.

Under the Maastricht Treaty, the binding criteria for these eight New Member States (NMSs) into the eurozone are exclusively described in nominal terms[4]. However, the fulfilment of Maastricht criteria by no means ensures that the NMSs will enjoy the net benefits of the monetary union. The extent to which the NMSs will benefit from giving up their monetary and exchange rate independence, in addition to the broader issue of the sound functioning of the enlarged eurozone, is generally discussed in terms of *real* factors, in particular, the degree of real convergence between the NMSs and the participating euro area countries.

There is no generally accepted indicator of real convergence. The European Commission itself, in various contexts, refers to such indicators as the balance of payment position, and to financial and product market integration (Convergence Report 2004**)**. Other research papers (Frankel (2004), Fidrmuc and Korhonen (2004), Kocenda et. al (2006)) focus on narrowing gaps of productivity or real income between respective countries and the euro area average, or concentrate on the correlation of business cycles.

This paper proposes a definition of real convergence which is based on real exchange rate volatility, i.e., it measures the degree of real convergence between a particular NMS and the selected group of eurozone countries by comparing the degree of real exchange rate volatility in this country and the average real exchange rate volatility estimated for the proposed eurozone members. While real exchange rate volatility analysis is not new, and in the context of optimal currency areas goes back to Vaubel (1976, 1978), to the best of the author's knowledge it has not yet been explicitly applied as a measure of real convergence.

What makes the scale of real exchange rate volatility a useful measure of the degree of real convergence? Real exchange rate volatility reflects underlying economic conditions in a number of ways. Under the assumptions of price and wage rigidity, the magnitude of real exchange rate volatility between a particular NMS and the eurozone captures:

- The extent to which flexible adjustment mechanisms exist in that NMS, other than the nominal exchange rate (i.e., the degree to which real exchange rates react to real asymmetric shocks). Such mechanisms might include, *inter alia*, factor mobility, fiscal policy, and the flexibility and shock-absorbing capacity of the financial sector.

- The degree to which the real exchange rate between the NMS and the eurozone is exposed to real asymmetric shocks. The degree of this symmetry would in turn depend on

[3] These were the Czech Republic, Estonia, Hungary, Latvia, Lithuania, Poland, the Slovak Republic and Slovenia. Although during the time of writing Slovenia joined the eurozone (in January 2007), it was left in the sample for comparative purposes.

[4] Maastricht criteria relate to the nominal exchange rate, the budget, public debt, inflation rate and long-term interest rates.

the similarities in price levels and GDP per capita, labour mobility, the synchronisation of business cycles, structural similarities, convergence of the interest rate differential between that NMS and the eurozone, the degree of trade openness and trade diversification of that NMS, the degree of stability in terms of trade, and financial market integration.

- The degree to which monetary policies in the NMS and the eurozone react symmetrically to symmetric shocks.

The lower the degree of real exchange rate volatility, the greater the extent of adjustment mechanisms other than the nominal exchange rate, and/or the lower the exposure of the NMS to asymmetric shocks, and/or the greater the degree of symmetry between the effects of monetary policy in response to symmetric shocks[5].

The existence of flexible adjustment mechanisms other than the nominal exchange rate, in the onset of unexpected real shocks, allows smooth tuning of macroeconomic imbalances, limiting the need for an exchange rate's adjustments. Given that symmetric shocks do not require adjustments in relative prices, they do not distort equilibrium. Consequently, *a less volatile real exchange rate indicates less scope for monetary and exchange rate independence*. These economic conditions, which guarantee a more stable real exchange rate, are also traditional arguments behind the successful creation of optimal currency areas (Mundell (1961), McKinnon (1963), Kenen (1969))[6].

An additional advantage of the proposed definition of real convergence is that the real exchange rate volatility criterion does not depend on the exchange rate in place, nor on the fact that a system actually chosen is optimal for the country. It only relies on the assumption that *national price stability is desirable*, and that therefore the flexibility of the nominal exchange rate may be justified to avoid changes in the real exchange rate that entail inflation or deflation above or below the eurozone average.

However, without empirically verifying the shock-absorbing role of a nominal exchange rate it is not possible to assess to what extent giving up monetary and exchange rate independence actually constitutes a cost of euro adoption, and to what extent nominal flexibility facilitates convergence. Having confirmed a *high* degree of real exchange rate volatility due to *real* shocks, the only inference one is able to convincingly draw - based on the proposed definition – is that it is not yet advisable to join ERMII or the eurozone (or both)[7]. Given the nature of the shock and the catching-up process in the NMSs, in many ways such a conclusion could be sufficient (i.e., it may

[5] Managing large real internal or external imbalances in countries with sizable asymmetric real shocks may prove to be difficult, especially during the ERMII period. Further elaboration is included in the longer version of this paper (Blaszkiewicz-Schwartzman (2006)).

[6] The traditional arguments however have not gone unchallenged. Based on Mundell (1973), it has been argued that if members of a currency zone are financially integrated, then a high degree of symmetry of the shocks among them, although desirable, is no longer a prerequisite. This is because, in a currency area, asymmetric shocks can be smoothed through risk sharing – i.e. through portfolio diversification and pooling of foreign exchange reserves (see Blaszkiewicz-Schwartzman and Wozniak (2005) for an overview of this literature). However, the risk-sharing argument does not change the fact that, in the presence of nominal rigidities, fewer asymmetric shocks call for a smaller need to adjust, and that giving up monetary and exchange rate independence represents a cost of monetary unification. This remains the basis of the approach set out in this paper.

[7] Eichengreen (1991) argues that real exchange rate variance analysis is not able to distinguish between the size of a shock and the ability to cope with it. Even if this were true, here it is argued that it does not matter if the volatility is high due to the degree of asymmetry or because the absorbing potential of other adjustment mechanisms is low. The outcome is the same: it is costly to join the common currency area.

be that the only appropriate way to move these countries to higher income levels is exactly via higher inflation, implying that these countries should not rush to give up their own currencies). Nevertheless, this conclusion could be further reinforced were one able to empirically confirm the nominal exchange rate's theoretical ability to induce rapid adjustments in the onset of idiosyncratic real shocks. Were this is the case, it would be possible to argue that the nominal exchange rate *is* an important channel for the real exchange rate changes, and thus plays a positive role during the convergence process (i.e., stabilises real shocks in the absence of other adjustment mechanisms and sluggish prices). Loosing this instrument represents the cost of the monetary integration and can have negative implications for countries' economic performance[8]. Of course, if this role is not confirmed, given high real asymmetries, recommendations on the timing of ERMII/ euro area accession would not change. Unless greater real convergence is achieved it may be too costly to share a common monetary policy. Still, it would be also obvious that since shocks cannot be addressed by monetary policy, the only way to achieve real convergence is via implementation of structural reforms. Moreover, if the nominal exchange rate did *not* play a shock-stabilising role, the scale of real shock asymmetry would indicate the degree of flexibility of other adjustment mechanisms (i.e., labour mobility or real wages)[9].

The approach developed in this paper builds on two strands of empirical literature with roots in the early theory of optimal currency areas (OCAs). The first of these focuses on the degree of real asymmetry between countries or regions wishing to constitute currency areas (Vaubel (1976, 1978), von Hagen and Neumann (1994), Gros and Hobza (2003), Blaszkiewicz-Schwartzman and Wozniak (2005)). The second strand of empirical literature attempts to test the main assumption of the OCA theory, and detect whether exchange rate flexibility is a significant stabilizer of real asymmetric shocks (Bayoumi and Eichengreen (1989), Clarida and Gali (1994), Canzoneri et al., (1996) in the context of developed countries, and Dibooglu and Kutan (2001) and Borghijs and Kuijs (2004) in the context of the NMSs). All these papers utilise standard assumptions of open macroeconomy sticky-price models in the spirit of Mundell-Fleming-Dornbusch, to classify shocks in different SVAR systems.

In common with this first strand of literature, this paper focuses on real exchange rate volatility. In common with the second strand of literature, this paper attempts to separate shocks governing movements of real and nominal exchange rate into their nominal and real components, and to detect the responses of nominal exchange rates to different types of disturbances.

The paper draws on a two-step univariate / bivariate methodology. In step one, the univariate approach measures the degree of unexpected real exchange rate variance (and thus the degree of real convergence) through a GARCH econometric methodology. In step two, the bivariate

[8] Whether or not exchange rates serve as effective shock stabilisers depends to a large extent on the price strategies governing firms' decisions (as stressed by New Open Economy Macro Models). For example, under conditions of local-currency-pricing, nominal currency changes would not change either real or nominal prices in the short run. However, in the context of the NMSs this is unlikely to be the case. Were it the case, observed real exchange rate volatility would have to be induced by market incompleteness and exporters' ability to discriminate against different markets, requiring that relative prices stay constant. Yet in the NMSs, inflation rates fell dramatically during the 1990s. (See also Engel (2002), who shows that if importer-distributors face pass-through to import-prices, then some flexibility may be still desirable. Similarly, Obstfeld (2002) brings empirically supporting evidence that there is still an important role for exchange rate flexibility to play in changing relative prices). Therefore, it is probably fair to assume that nominal exchange rates are not totally 'disconnected' from the real economy in the NMSs and - at least to a certain degree - are able to provide equilibrating real exchange rate adjustments.

[9] Buiter (2000) emphasize that the decision to join a monetary union, is a monetary issue. If prices of goods are flexible, relative-price behaviour is usually independent of the monetary regime. The choice of monetary regime only matters for short-run changes – the period during which nominal prices are adjusting. In this paper it is however argued that in the context of catching-up economies this decision does depend on the degree of real convergence, as the only way to reach higher income levels is via higher than the eurozone average growth rates, and thus inflation.

approach comprises a structural VAR analysis with a Blanchard and Quah decomposition (1989) (BQ-SVAR), and facilitates the identification of nominal and real factors driving real and nominal exchange rate movements and a potentially stabilising role of nominal exchange rates.

The two-step strategy is necessary for two reasons. First, it is essential for the accurate measurement of real convergence. This is because the univariate variance approach cannot convincingly distinguish between nominal and real shocks in real exchange rate movements, and therefore on its own is not well designed to accurately assess the degree real convergence. Therefore, the BQ-SVAR methodology is used to separate and measure the magnitudes of real and nominal components in real exchange rate movements. Second, the BQ-SVAR approach provides an indication of the shock-stabilising role that the nominal exchange rate plays in any NMS – i.e. the methodology establishes if the nominal exchange rate indeed responds to asymmetric real shocks, and moves together and in the same direction as the real exchange rate in order to ensure the necessary change in relative prices. Thus, the structural VAR analysis makes it possible to assess to what extent giving up monetary and exchange rate independence actually constitutes a cost of the euro adoption, and to what extent nominal flexibility facilitates convergence.

Although Dibooglu and Kutan also utilize a two-dimensional BQ-SVAR methodology (on a differenced real exchange rate and prices) in investigating the sources of real movements in Hungary and Poland, the aim of their paper differs from the one pursued here[10]. The authors examine the proposition that different fiscal and monetary policies in transition countries should lead to the predominance of real shocks in some countries, but nominal shocks in others and covers the period between 1990 and 1999. Their results suggest that during that time the Polish real exchange rate was mainly driven by nominal shocks (in the short-run) whereas, the Hungarian real exchange rate was driven by real shocks. However, the span of their sample includes a period of little nominal exchange rate flexibility and therefore cannot address the issues discussed in this paper. Also, they do not measure the size of real exchange rate volatility and therefore, based on their paper, very little insight about the process of real convergence can be gleaned.

The purpose of the Borghijs and Kuijs paper is to find out whether for five New Member States - the Czech Republic, Hungary, Poland, the Slovak Republic, and Slovenia - nominal exchange rate flexibility is a useful absorber of real shocks or an unhelpful propagator of monetary and financial shocks. The authors work within the three-equation model in the spirit of Clarida and Gali, and answer similar questions to Canzoneri et al., but their SVAR model includes a nominal exchange rate instead of prices since they argue that the loss of nominal exchange rate flexibility is the key cost of euro area participation. However, as pointed out above, the role of the nominal exchange rate as a shock absorber is only relevant if there are large real asymmetries between the economies wishing to form a common currency zone – the issue addressed in this paper. This is because even if a nominal exchange rate were not addressing macroeconomic imbalances and its movements were only a reflection of money and financial market shocks, one could not say that there is no cost from loosing monetary and exchange rate independence.

Application of the two-step methodology proposed here, as described in more detail in the body of this paper, suggests that real asymmetric shocks (i.e., the degree of real exchange rate volatility scaled down for the presence of nominal shocks) in the NMSs (with the exception of Slovenia and Estonia) outsize those experienced by Old Member States (OMSs) at the time of their euro

[10] The papers which identify sources of nominal and real exchange rates fluctuations within the bivariate BQ-SVAR models in developed countries, among others, include the works of Lastrapes (1992) and Enders and Lee (1997).

adoption process. This finding suggests that the NMSs are still converging in real terms on the basis of the proposed indicator. Additionally, it was found that in the NMSs, the nominal exchange rate *does* play a stabilising role (with the exception of Slovenia), and that nominal shocks do not, on average, move real exchange rates. Given that the benefits of monetary union are not immediately obvious at present, some caution should be exercised in timing the ERMII accession and euro adoption.

While the methodology used in this paper provides results useful to the policy questions raised by the prospect of the euro adoption, it is not without limitations, noted below. The broad decomposition of shocks into real and nominal components is both a strength and weakness. On one hand, the methodology does indicate whether or not nominal exchange rates move in the same direction as real exchange rates at the onset of a real shock, pointing to the stabilising role of the nominal exchange rate. On the other hand, it is not able to assess fully the destabilising role of the nominal exchange rate at the onset of the nominal shock. Even if the ex-post data revealed that variations in the nominal exchange rate were caused by a different type of shock to variations in the real rate, this could not be conclusively interpreted as an indication of the ineffectiveness of the nominal exchange rate to stabilise nominal shocks. An equivalently valid explanation could be that a nominal exchange rate fully cushioned the impact of a nominal shock on a relative price. This argument could be even stronger, given that nominal shocks represent a whole range of temporary shocks, such as supply, demand or monetary and financial shocks[11]. The only inference one would be able to make from such a result, would be that neither monetary policy nor fiscal policy can change competitiveness of a given country (and vice-versa, provided nominal shocks turned out to be important in real and nominal exchange rate movements). However, to the extent that the primary interest of this paper is to assess the importance of permanent movements in the real exchange rate, and the potential role of flexible regimes in stabilising permanent shocks (i.e., demand and supply shocks related to the convergence process) this decomposition is sufficient.

The reminder of the paper is organised as follows: Part 2 analyses the choice of the sample and data properties, including recent developments of nominal and real exchange rates, as well as prices in the NMSs, the evolution of exchange rate regimes, and data integration properties. Part 3 sets out and explains methodologies utilised in the univariate estimation of nominal and real exchange rate variances, as well as in the BQ-SVAR model, which is used to identify two structural shocks (i.e., temporary and permanent). Part 4 presents the results. Part 5 concludes.

PART 2: DATA ANALYSIS

The rationale for the choice of countries used in the sample, as well as the sample time span for both the univariate and bivariate analysis, are set out below. Since the data employed in the study should be stationary, the recent evolution is discussed of the real and nominal exchange rate as well as price movements in the selected countries, as a pre-step towards detecting integration properties of the data. Finally, the formal unit root tests are conducted.

[11] The same arguments apply to the 3-equation VAR system, and therefore the estimation of such a system would also fail to fully resolve the question of whether flexible exchange rates are destabilizing or not. It is true that a three-variable SVAR model distinguishes between demand and supply shocks (which the bivariate system cannot), but again its identification specification is not able to unambiguously separate between impact of the temporary supply and monetary/financial shocks on the short-run behaviour of the nominal exchange rate.

2.1 Sample Choice and Size

The NMSs analysed in this study include the Czech Republic, Estonia, Hungary, Latvia, Lithuania, Poland, the Slovak Republic and Slovenia.

The sample for inclusion in the two step univariate / bivariate methodology used in this paper gives rise to the following issues:

- First, while all NMSs can be included in the univariate sample (i.e., because of a real exchange rate flexibility), bivariate SVAR analysis (with real and nominal exchange rates) can only be applied to countries with relatively flexible nominal exchange rates. As a result, not all the countries included in the univariate analysis were included in the bivariate analysis.

- Second, countries included in the SVAR analysis must have *de facto* variable exchange rates. In some cases, *de facto* exchange rates differ from officially announced exchange rate regimes. In order to distinguish between different exchange rate regimes, officially announced exchange rate arrangements (as published by the IMF) were cross-checked with the classification developed by Reinhart and Rogoff (2002)[12]. According to both classification schemes, the exchange rates of Czech Republic, Hungary, Poland, the Slovak Republic, and Slovenia may be regarded as relatively flexible exchange rates. The exchange rates of Estonia, Latvia, and Lithuania are regarded as fixed, and could not be included in the bivariate SVAR analysis[13]. Box 1 reviews the evolution of nominal exchange rate regimes in the five NMSs which may be regarded as having flexible exchange rates.

- Third, meaningful structural analysis requires sufficiently long data span. Unfortunately, for countries under consideration, reliable data only exists from the beginning of the 1990s. As a result, for the univariate analysis the estimation period spans 1993M1 to 2007M11. Prior to this period, the data is contaminated by structural changes related to the transition process. The data span used for the bivariate analysis is based on the *de facto* flexible exchange rate regime in place, as described in Box 1.

The sample of current euro area countries chosen for comparison with the NMSs includes Italy, Greece, Portugal and Spain (the so-called Club Med countries) as well as France and Germany. The Club Med countries are regarded as belonging to the periphery of the EU, while France and Germany are chosen to represent the core of the EU. The span of data for chosen eurozone countries runs from January 1993 to December 1998. This choice reflects the following factors: First, 1993 marks the end of the European Monetary System, which allowed nominal exchange rates to fluctuate within a band of +/-15 percent. This ensures minimum policy coordination between countries and is important for comparative purposes. Data after December 1998 is not considered, as for the purposes of this study, the performance of countries after their entry into the Eurozone in January 1999 is not of interest.

Unfortunately, according to the Reinhart and Rogoff's classification, the only country within this selected group with a *de facto* floating exchange rate regime was Germany. Nevertheless, because

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[12] Because Reinhart and Rogoff's study goes back only to December 2001, exchange rate regimes between 2001M12 and 2007M11 were classified in accordance with the IMF code.

[13] From now on, whenever the reference is made to the *de facto* exchange rate regime, it refers to Reinhart and Rogoff's classification.

between 1993 and 1998 the Club Med countries as well as France adopted some kind of peg or crawling band regime against the DM, at the same time fluctuating freely around the ECU, they were all included in the SVAR modelling[14].

2.2 Data Source and Transformation

For all NMSs, monthly data on period average nominal exchange rates, against the euro, up to November 2007 were sourced from Eurostat. Eurostat also provided data for the euro area consumer price index (HICP). Consumer price indices (CPIs) for the New and Old Member States, as well as former eurozone national currencies vs. euro/ECU considered in the sample were taken from the IMF IFS[15]. All series were transformed into logarithms, and scaled with the base period set to 100 in 2005 for the NMSs, and to 1995 for the OMSs. The individual real exchange rate indices were calculated as nominal NMS/euro rates, deflated by the relevant consumer price indices (i.e. CPI for NMSs and HICP for the eurozone) [16].

2.3 Graphical Presentation

Figure 2, Annex 1, presents the developments of real and nominal exchange rates as well as price ratios (defined as P^{EMU}/P) between 1993M1 and 2007M11, for the NMSs included in the univariate and/or bivariate variance analysis. It shows that in countries with relatively flexible nominal exchange rates, real and nominal exchange rates tend to move together, as indicated by the coinciding turning points[17]. This outcome is confirmed by the simple correlation between real and nominal exchange rates in these countries. This coefficient is approximately 0.9 for the Czech Republic, Hungary, Latvia, Poland and Slovakia and equal to approximately 0.8 for Lithuania, and Slovenia. The lowest comovements between real and nominal exchange rates are observed in Estonia, where the correlation coefficient is 0.3 (see Table A.1.1 in Annex 1 for details). Despite a high correlation, over time, nominal and real exchange rates diverge or move in different directions in the case of Hungary and Slovenia[18]. The differences in the short and long-time dynamics of real and nominal exchange rate point to the presence of two different types of shocks affecting these countries: one temporary and one permanent in nature. This is consistent with the predictions of the broad class of structural open macro models (i.e., Dornbusch's 'overshooting' model or Stockman's 'equilibrium' model). Given that the divergence of the rates occurs quickly, there exists a strong pre-assumption that permanent shocks dominate real exchange rate movement.

[14] The results of analysis conducted on the OMSs treated in this paper are not discussed in detail, as they only serve a point of comparison. Further details may be obtained in the longer version of this paper.
[15] HICP indices for NMSs are not available over the time period estimated in this study.
[16] An increase in the index indicates currency depreciation relatively to the euro.
[17] These observations are not unique. Enders and Lee (1997), for instance, have noted similar trends in Canada and Japan.
[18] Given the objective of monetary authorities to keep the real exchange rate constant in Slovenia, and to limit initial flexibility in Hungary until 2001, this is not surprising.

Box 1: Evolution of Exchange Rate Regimes in the Czech Republic, Hungary, Poland, the Slovak Republic, and Slovenia (Choice of Sample Size)

In the *Czech Republic,* exchange rate flexibility was limited before 1996. Initially the official exchange rate was tied to a currency basket and then to the ECU. *De facto*, however, the country had a crawling band system around the DM (with a band width of ±2%). More flexibility was introduced in May 1997. The Czech koruna was officially classified as a pre-announced crawling band around the DM with a band width of ± 7.5% (*de facto* the band width was ±5%). Because, between 1993M1 and 1996M3, the official regime was less flexible than indicated by the *de facto* regime, the final sample for the Czech Republic spans from 1996M3 to 2007M11.

In *Hungary,* until December 1998, the exchange rate regime was a *de facto* crawling band around the DM, with a band width of ±5%, until May 1994, and ±2% between May 1994 and January 1999. From January 1999 to December 2001, the exchange rate was *de facto* classified as a pre-announced crawling band around the euro. Officially, more flexibility was introduced in May 2001. The crawling band was widened from ±2.25% to ±15%. While more official flexibility was announced in 2001, it is not possible to conduct analysis on so few data points. Given this, and the fact that there was already some flexibility before 2001, the estimation period used for Hungary covers the years 1993M1-2007M11.

The sample size for *Poland* starts in June 1995 since before that a *de facto* exchange rate regime was either classified as freely falling (i.e., period of hyperinflation) or dual market. From mid-May 1995 up to February 1998, the *de facto* regime was classified as a crawling band around the euro (ECU) with a band width of ±5%; there was a pre-announced crawling band around the DM and the US dollar of ±7%. Between February 1998 and April 2000, the band width was systematically widened (up to ±15%). In April 2000, a float was introduced (i.e., a *de facto* managed float). The regime has not changed since then. The final sample size spans 1995M6 to 2007M11.

In *Slovakia* exchange rate flexibility was introduced gradually. Between 1993M1 and 1996M7, the currency was *de facto* governed by a crawling band regime around the DM with a band width of ±2%. The band width did not change up to September 1997, but between August 1996 and September 1997 the pre-announced crawling band was progressively widened to ±7%. As of September 1997, *de facto* the band was widened to ±5% and a pre-announced crawling band of ±7% was maintained. Even though the managed float system was introduced in October 1998, according to Reinhart and Rogoff, between October 1998 and December 2001, all the observations remained within a ±5% band of DM/euro. Taking into account policy changes in the exchange rate regime, the estimation period starts in 1996M8 and ends in 2007M11.

Between the years 1993 and 2004, the nominal exchange rate in *Slovenia* was governed by a *de facto* crawling band around either the DM, or euro with a band width of ±2% (euro/ECU replaced DM in October 1996). From June 2004 Slovenia has been participating in the ERMII system in which the exchange rate is allowed to fluctuate by ±15%. Unfortunately, the period of greater *de jure* flexibility is not long enough to perform the estimation. Therefore, estimation based on data spanning 1993M4 to 2006M12 has been used (before April 1993 a *de facto* regime was classified as freely falling; Slovenia adopted the euro on January 1 2007). The final results, however, are presented for the period 1996M1 to 2006M12 (this relates to issues of heteroskedasticity, and will be discussed in further depth below).

Source: Compiled by the author based on Reinhart and Rogoff (2002) and the IMF classification.

2.4 Integration Properties

This Section formally tests the unit root hypothesis for the data series used in this study. In the case of the univariate analysis, nonstationarity of data in levels would imply that real exchange rate movements cannot be characterised by their average values. In such cases it would be inappropriate to use standard measures of volatility, such as variance/ standard deviation of the series. Stationary data is also required for the Blanchard and Quah decomposition of the SVAR model. Moreover, the variables in a VAR should not be cointegrated if the data in levels is non-stationary. To test cointegration between the pairs of exchange rates entering the VAR, it is enough to check the integration properties of price ratios in levels (i.e., the price ratio between the eurozone and country of interest inflation) (see Enders and Lee (1997)). Only when all the variables are I(1) and no cointegrating relationship exists, is it appropriate to test the VAR in first differences. The results of the formal unit root analysis discussed below should however be treated with great caution as the time span on which the tests are conducted is very short.

Following Maddala's and Kim's (2002) argument that the Dickey-Fuller, augmented DF, and Philips-Perron unit root tests do not have enough power to meaningfully reject the null hypothesis, these tests are not used. Instead, in this study the DF-GLS test of Elliot-Rothenberg and Stock (1996) as well as the class of MZt and MZa tests of Ng and Perron (2000) are applied. As suggested by Ng-Perron (based on Monte Carlo simulations), in order to maximize the power, all tests are based on GLS detrending; likewise, in order to minimise the size distortion under the null (and not over-parameterise under the alternative), the choice of the lag length is selected on the basis of the Modified Akaike Information Criteria (MAIC). The maximum number of lags is set in accordance with the rule suggested by Schwert (1989).

Given that DF-GLS and MZ-GLS tests may not be appropriate for variables with an apparent structural break (see Perron (1989), Christiano (1992), Zivot and Andrews (1992)), the unit root test of Perron (1997) which allows a parsimonious single structural break is also carried out. The structural break date is treated as unknown and chosen so as to minimize the t-statistic on the α coefficient (i.e., model 3 in Perron (1997, p. 358)). The number of lags is determined by the 'general-to-specific' procedure with the maximum number of lags specified as in the previous tests.

The performed DF-GLS and MZ tests indicate non-stationarity of the investigated data (i.e., the data is I(1) - see Annex 1 Table A.1.2) for all the series to be used in the univariate and bivariate modelling, but for the real exchange rate for Slovenia. The result for Slovenia, however, was not confirmed by the unit root test with a break (see Annex 1 Table A.1.3). In the case of Hungary and France, there is some evidence of the stationarity of the nominal exchange rate and price ratio based on the unit root test with a break. Therefore, as a check, the Kwiatkowski, Phillips, Schmidt and Shin's (KPPS) test was done. This test sets a stationarity hypothesis as the null and was suggested by Kim and Maddala (2002) as confirmatory analysis. In this case, a stationarity hypothesis for the Slovenian real and Hungarian nominal exchange rates as well as the price ratio series for France could not be accepted at the 1%, 5% and 10% level.

Based on the results from the unit root test, the series of real and nominal exchange rate used in this study enter univariate and bivariate estimations in first differences. Additionally, all the countries selected in Section 2.1 above, are included in the SVAR modelling, since: (i) all the exchange rate series of countries previously proposed for the structural VAR analysis are non-stationary in levels (with some uncertainty as far as the real and nominal exchange rate in Slovenia and Hungary are concerned); and (ii) the integration properties of the ratio of prices in levels do not suggest cointegration between respective pairs of nominal and real exchange rates in those countries (with some uncertainty in the French case). Once again, given the short span on the data, the performed tests are rather indicative than conclusive[19].

PART 3: ECONOMETRIC METHODOLOGY

The univariate and bivariate components of the econometric methodology employed in the study are explained in further detail below.

3.1 Univariate Variance Analysis: Technical Aspects

Empirical analysis of real exchange rate movements involves estimating the *unexpected* (i.e., conditional) real exchange rate variances between the respective NMSs and the selected euro area members treated as a group. As set out in Part I, this approach draws on Vaubel and is similar to that of von Hagen and Neumann, Blaszkiewicz and Wozniak and Gros and Hobza[20]. Again, given the unit root processes in real exchange rate series, the *unexpected* component of real exchange changes (i.e., fluctuations which cannot be explained by past RER movements) for each country of interest is obtained using AR(p)-GARCH(p,q) econometric methodology, by regressing real exchange rate changes on their own lags, as follows:

$$\Delta rer_{i,t} = b_0 + b_1 \Delta rer_{i,t-1} + b_2 \Delta rer_{i,t-2} + ... + b_{12} \Delta rer_{i,t-12} + u_{i,t} \tag{5}$$

where $\Delta rer_{i,t}$ is the change in the real exchange rate[21]. Residuals obtained from these regressions represent conditional real exchange rate shocks. Next, the standard deviations of these shocks (i.e., the measure of real convergence) are measured.

In order to see whether the degree of volatility has been changing over time, this exercise is done for 3 sub-samples (1993-1995, 1996-1998, 1999-2006/7). The sub-samples are chosen to roughly represent the periods of nominal exchange rate regimes' movements toward greater flexibility in the countries in question (see Egert and Kierzkowski (2003)). To check whether these volatility changes are statistically significant over time (i.e., test for variance equality between sub-samples), equation 5 is estimated on two sub-samples (1993-1998 and 1996-2006/7) by OLS[22]. Various statistical tests are then performed. Von Hagen and Neumann (1994) propose White's tests for heteroskedasticity. Additionally, an ARCH test is carried out, as financial market data often follow an ARCH/GARCH process.

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[19] Kutan and Dibooglu (2001) argue that assuming non-stationary real exchange rates in transition economies is reasonable as purchasing power parity, implying stationary real exchange rates, holds under very restrictive conditions, which are extremely unlikely to be met in the case of the transition economies. Moreover, equilibrium real exchange rates in these countries should exhibit an upward trend over time due to the catching up process and as productivity and real wages increase over time. Because such shocks are generally stochastic in nature, there is a strong presumption that real exchange rates should have a permanent component during the time-span covered by their study. The same arguments should hold for the NMSs in the period of 1993-2007.

[20] However, the estimation methods in these studies differ from the method utilized in this paper. Moreover, Gros and Hobza look at observed rather than unexpected exchange rate variability.

[21] The final number of lags in individual AR(p)-GARCH(p,q) equations was determined by the 'general-to-specific' approach.

[22] Notice, that GARCH models for the three sub-samples were estimated regardless of the ARCH test's results performed on equivalent models estimated by OLS. This is because ARCH test is an asymptotic test and the estimated sample size is very small. Moreover, the OLS estimates, which are available from the author upon request, in most cases, brought similar results.

Finally, for each country, AR(p)-GARCH(p,q) models for a change in the real exchange rate for the whole sample period are estimated (i.e. 1993-2006/7). This is done in order to obtain the plots of the estimated time varying conditional variances for real exchange rates. Similar to conditional real exchange rate shocks measured for different sub-samples, these plots should help assess whether the NMSs are indeed converging over time in real terms. This indicator also avoids the choice of a somewhat arbitrary split into sub-samples.

To assess the magnitude of these real exchange rate variances (i.e., to decide when the variance should be considered large, and when small), averages of estimates of the observed real exchange rate volatility of the selected current euro area members are also provided. These were obtained by applying the same methodology to the two sub-samples (i.e., 1993-1995 and 1996-1998) as well as the whole sample (i.e. 1993-1998).

The univariate variance approach is only well designed to accurately assess the degree real convergence if – among other things - it can precisely measure the degree of real shock asymmetry, and thus the degree of real convergence. To this end, nominal shocks should be eliminated from real exchange rate movements. This is not an easy task and constitutes one of the main drawbacks of unvariate variance analysis. An attempt however is made to tackle this problem. Von Hagen and Neumann (1994) make a strong assumption that high-frequency (monthly) real exchange rate changes mostly reflect nominal shocks, and low-frequency (quarterly) real exchange rate changes are principally due to real shocks. However, because the analysis carried out on quarterly data necessarily means using fewer data points, estimations for the respective sub-samples are not possible. Therefore, a quarterly data analysis is only carried out on the whole sample. The AR(p)-GARCH(p,q) models are estimated, and the plots of the time varying conditional variances are analysed. This, together with the BQ-SVAR analysis described below, is the basis for evaluating the differences between the scale of asymmetric real and nominal shocks in real exchange rates.

3.2 Bivariate Variance Analysis: Technical Aspects

Given that the variables of interest, real and nominal exchange rates (rer_t and ner_t, respectively), have a single unit root (and are not cointegrated), the VAR model considered in the study can be written as follows:

$$B\Delta y_t = \Gamma_0 + \Gamma(L)\Delta y_{t-1} + \varepsilon_t \tag{6}$$

where $\Delta y_t = (\Delta rer, \Delta ner)'$, B is a 2×2 invertible matrix, Γ_0 is a 2×1 matrix of constants, $\Gamma(L)$ is a 2×2 polynomial in the lag operator, and $\varepsilon_t = (\varepsilon_{1t}, \varepsilon_{2t})'$ is a vector of white-noise structural disturbances, i.e., $\varepsilon_t \sim iid(0, D)$ with D being a variance-covariance matrix of structural disturbances. ε_{1t} is interpreted as a real shock with possible permanent effects on nominal and real exchange rates.

ε_{2t} stands for a nominal shock with only short-run effects on a real exchange rate. This broad classification of shocks is consistent with the Dornbusch (1976) ‘overshooting’ model of a small open economy in which nominal shocks can have permanent effects on the nominal exchange rate, but only temporary effects on the real exchange rate (Lastrapes (1992), Enders and Lee (1997)).

Given that there are more parameters than equations to be estimated, the inference starts from estimating the flowing reduced form VAR model by OLS[23]:

$$\Delta y_t = C_0 + C(L)\Delta y_{t-1} + e_t \tag{7}$$

where $C_0 = B^{-1}\Gamma_0$, $C(L) = B^{-1}\Gamma(L)$ and $e_t = B^{-1}\varepsilon_t$

It is further assumed that $e_t \sim iid(0,\Omega)$ where Ω is a variance-covariance matrix of the reduced form error term. This matrix can be expressed as:

$$\Omega = B^{-1}D(B^{-1})' \tag{8}$$

Now, in order to recover to structural disturbances, ε_t, from the reduced form VAR, B^{-1} must be identified. As Blanchard and Quah (1989) show, this can be done by imposing long-run (infinite-horizon) restrictions on the matrix of structural dynamics multipliers $\Theta(1)$ which can be obtained by estimating the moving average representation of Δy_t and then by re-writing it in terms of <u>structural</u> shock:

$$\Delta y_t = \mu + \Theta(L)\varepsilon_t \tag{9}$$

Because it was assumed that ε_{2t} has no long-run effect on a real exchange rate, $\Theta(1)$ can be obtained as a lower triangular. Since $\Theta(1)$ equals $(I - C(1))^{-1}B^{-1}$:

$$B^{-1} = [(I - C(1))\,\Theta(1)] \tag{10}$$

Using this expression, the reduced form long-run variance-covariance matrix can be expressed as:

$$[(I - C(1))^{-1}]\,\Omega\,[(I - C(1))']^{-1} = \Theta(1)D\,\Theta(1)' \tag{11}$$

The left hand side of this expression can by fully obtained by estimating the reduced form VAR by OLS. Normalising D to the identity matrix and given the imposed long-run restriction on $\Theta(1)$ enables $\Theta(1)$ to be fully identified through the system of equations specified in eq. (16); when $\Theta(1)$ is identified, B^{-1} is also identified, and so are the *structural disturbances,* $\varepsilon_t = (\varepsilon_{1t,}\varepsilon_{2t})'$:

$$\lim_{s\to\infty}\begin{bmatrix} rer_{t+s} \\ ner_{t+s} \end{bmatrix} = \begin{bmatrix} \theta_{11}(1) & 0 \\ \theta_{21}(1) & \theta_{22}(1) \end{bmatrix}\begin{bmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \end{bmatrix} \tag{12}$$

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[23] As discussed in Hamilton (1994) separate VAR equation can be estimated by OLS without loosing efficiency since, with the normality assumption, OLS estimators are almost identical with the maximum likelihood (ML) estimators.

Given that $\Theta(1)$ is now fully identified, it is possible to test the **additional** identifying restriction $\theta_{22}(1) = 1$ (i.e., so far, in order to identify the system, it was assumed, in accordance with the broad class of open economy macro models, that the long-run effect of the nominal shocks on the real exchange rate is zero, i.e., $\theta_{12}(1) = 0$), which says that a nominal shock has a proportional effect on a nominal exchange rate (see Enders and Lee). Since a positive nominal shock should cause a currency to depreciate, even if the long-run impact of the shock is not proportional, the expected sign on the estimated coefficient $\theta_{22}(1)$ is positive.

PART 4: ESTIMATION RESULTS

4.1 Univariate Variance Analysis.

This Section presents the results of the univariate variance analysis for the New Member States. Magnitudes of real exchange rate fluctuations (on a monthly and quarterly basis) are estimated and contrasted with magnitudes of such fluctuations in the selected OMSs.

Table 4.1.1 summarises estimates of conditional standard deviations (CSDs) of monthly and quarterly real exchange rate shocks for the NMSs as well as selected members of the eurozone obtained from GARCH models. As postulated, the quarterly estimates attempt to eliminate nominal variability in real exchange rate movements.

Based on the estimates for the whole sub-sample, on a monthly basis, among the NMSs, Poland displays the highest real exchange rate volatility, Slovenia the lowest. The average real exchange instability in the group of the NMSs, compared with the Club Med countries, is higher by approximately 1.2 times (by 3.1 times when compared with the average for France and Germany and by 1.5 times when compared with the average for the group of Club Med countries plus France and Germany (OMS)).

In terms of the results for the particular sub-samples, there are 3 countries for which monthly standard deviations of real exchange rate shocks exhibit a consistent and decreasing trend. These are Estonia, Lithuania and Slovenia. Hungary, Latvia, and Poland managed to decrease the variance of real exchange rate shocks between the II and I sub-sample. In the III sub-sample, real exchange rates again became again more volatile. In the Czech Republic there is clear evidence of stabilizing policies between 1999 and 2006/07. As for the Slovak Republic, there has been a decrease in real exchange rate volatility between the II and III sub-sample. Based on the two sub-sample (1993-1998 and 1996-2007) estimates of equation 5 by OLS, statistically significant changes occurred in the Czech Republic, Latvia, Poland and Slovenia in the first sub-sample and in all considered NMSs but Poland and Slovakia in the second sub-sample.

When the average magnitudes of the Club Med real exchange rate shocks in the early 1990s, as well as in years preceding the creation of the eurozone are compared with the NMSs average for 1999-2006/07, the results show that, on average, the NMSs real exchange rate volatility is 1.5 times higher than the real exchange rate volatility of Club Med countries in years preceding eurozone membership (i.e., 1996 to 1998), and almost equal to the variance of Club Med countries in the early 1990s. It should be stressed however that for countries like Estonia and Slovenia, real exchange rate volatility in years 1999-2006/07 is smaller in both sub-samples not only when compared with the Club Med countries, but also when compared France and Germany are added to the group.

Table 4.1.1 Real exchange rate volatility

VOLATILITY CHANGES		CSD I	CSD II	CSD III	White Heteroskedasticity		ARCH		GARCH M	GARCH Q
		93-95	96-98	99-07	93-98	96-07	93-98	96-07	Full Sample	
Czech Rep.	No*/Yes*	0.67	1.83	1.27	**0.00**	0.13	**0.01**	**0.03**	1.35	1.31
Estonia	Yes/Yes*	1.72	0.64	0.34	0.00	**0.00**	0.00	**0.00**	0.78	1.02
Hungary	Yes/No*	1.76	1.14	1.54	0.75	0.98	0.77	**0.00**	1.48	1.35
Latvia	Yes*/No*	3.75	1.18	1.28	0.27	**0.08**	**0.06**	**0.01**	1.81	1.60
Lithuania	Yes/Yes*	3.92	1.93	1.47	0.95	**0.00**	0.13	**0.00**	1.66	1.99
Poland	Yes*/No	2.16	1.96	2.15	0.88	0.23	**0.03**	0.18	**2.09**	**2.08**
Slovak Rep.	No/Yes	1.25	1.75	1.54	0.75	0.16	0.79	0.80	1.38	1.44
Slovenia	Yes*/Yes*	1.11	0.68	0.38	**0.08**	**0.00**	0.36	**0.01**	0.64	0.91
Average		***2.04***	***1.39***	***1.25***					***1.40***	***1.46***
Average (FRA and DEU)		***0.63***	***0.36***						***0.45***	***0.40***
Average (OMS)		***1.03***	***0.67***						***0.95***	***0.82***
Average (ClubMed)		***1.23***	***0.83***						***1.19***	***1.02***

Note: Columns labelled 'CSD' report conditional standard deviations (CSDs) from GARCH models for selected sub-samples (on a monthly basis). Columns labelled 'GARCH M' and 'GARCH Q' report CSDs from models estimated for the whole sample (on a monthly and quarterly basis, respectively). Quarterly CSDs were normalized to monthly CSDs. "Yes" indicates convergence, "No" indicates divergence, i.e., we observe a decrease/increase in the CSD of the real exchange rate between the two tested sub-samples (93-95 and 99-07); '*' marks statistically significant changes in CSDs of the real exchange rate between the two sub-samples estimated by OLS (based on p-values of conducted White Heteroskedasticity and/or ARCH tests and 10% significance levels).

Source: Author estimates based on IMF IFS and EUROSTAT data.

Turning into results obtained from the whole-sample quarterly estimates, in all NMSs but the Czech Republic, Hungary and Poland, the magnitude of individual quarterly real exchange rate variance is higher than that of monthly changes. Given that the assumption was made that unexpected quarterly real exchange rate volatility reflects real shocks which are free of short-run disturbances, and the fact that - with the exception of Slovenia and Estonia – relative quarterly real exchange volatility for the NMSs is higher than it was for the Club Med countries (on average 1.4 times), it is clear that asymmetric real shocks are still an important source of real exchange rate volatility in these countries.

Unfortunately, because of the short data span, the only way to discern quarterly changes in variances over time is to examine the plots of the estimated time varying conditional variances for real exchange rates. They are presented in Annex 2, Figure 4, and mostly confirm the monthly basis patterns (compare with Annex 2, Figure 3). Slovenia not only has limited volatility, but the magnitudes of this volatility is low and decreasing. Estonia shows similar, but weaker characteristics with recently increased volatility. In Latvia, after a drop in magnitudes between 1993 and 1996, there are no changes in time varying conditional variance. In Lithuania, there is a considerable drop in the amplitude of volatility between 1993 and 2007 indicating convergence. In the Czech Republic and Slovakia, although the quarterly volatility has been decreasing, the magnitudes are still above those observed in the OMSs. It is hard to discern significant convergence in Poland (despite the fact that in recent years the shocks seem to be lower, their scale is still significantly above the one observed for the OMSs). In Hungary, a divergence is observed. Interestingly, on a monthly basis, the plots for Estonia, Latvia and Lithuania, despite the obvious decreases in volatility (and its scale), pick up the current downside risks present in these economies.

In summary, based on univariate analysis and the proposed definition of real convergence, it appears that only Slovenia Estonia achieved a level of real convergence which is comparable with that of the selected OMSs in the onset of the euro adoption.

4.2 Bivariate Variance Analysis.

Since Section 2.4 brought clear evidence of a unit root in the data, the variables of interest, nominal and real exchange rates, enter VARs in first differences. Next, the strategy described in Section 3.2 is implemented. The VAR lag order for each country is chosen based on the Akaike Information Criterion (AIC). The estimated lag length $\hat{p}$ is chosen for the value of p that minimises AIC(p), with the maximum number of lags of p_{max}=8 (and 6 for the OMSs). Although the AIC criterion tends to overestimate the number of selected lags, as shown by Kilian (2001), impulse response estimates tend to be highly sensitive to the underestimation of a lag order.

Model Specification and Checking.

Before one can move to the structural VAR analysis of shocks (i.e. to evaluating the contribution of nominal and real shocks to exchange rate movements), it is necessary to make sure that the errors from the estimated reduced form VAR models are normal and i.i.d. For the models to be correctly specified, the estimated residuals should be normally distributed, serially uncorrelated and homoskedastic. Moreover, since the structural VAR analysis needs to be conducted in constant economic structures, and since there have been frequent changes in the monetary and exchange rate regimes in the NMSs during the period under consideration (which could have disrupted a stable relationship between the variables), tests for structural changes are also performed.

Normality, Autocorrelation and Heteroskedasticity.

In order to test whether the estimated VAR residuals exhibit any remaining autocorrelation, the Portmanteau and LM autocorrelation tests are executed. To test the normality assumption, the multivariate test of Doornik and Hansen (1994) is employed. Homoskedasticity is checked by performing general White's tests (joint test and tests for individual components with (i.e., test for heteroskedasticity and specification bias) and without cross-products (i.e., test for pure heteroskedasticity)). White's heteroskedasticity tests are primarily chosen because they neither require explicit formulation of the form of heteroscedasticity, nor do they require normality under the null hypothesis (i.e., no heteroskedasticity). The results of those tests are presented in Annex 3, Table A.3.1) [24].

Despite the fact that there seem to be no autocorrelation left in the residuals of the estimated VAR models, the results should be treated with caution because the misspecification autocorrelation tests are derived under the assumption of normally distributed errors, which is clearly violated in all cases but Poland (for the selected OMSs, three countries – Germany, Greece and Italy – do not pass normality tests). As indicated by White's tests, in the cases of Czech Republic, Slovakia, Slovenia, Spain, Italy and Portugal, the lack of normality could be due to heteroskedastic errors. In other countries, it could be because the distribution is skewed or leptokurtic, or simply because of the small sample size, which could be too small to confirm asymptotical normality.

Structural Changes.

In order to tests for structural breaks, the techniques developed by Bai, et. al., (1998) and Hansen (2000) are employed. Both treat the break date as unknown. In light of various econometric

[24] Since the results of Portmanteau tests did not differ from the results of LM tests, only the former are presented in Table A.3.1. The results of the LM tests can be obtained however from the author upon request.

studies, which document that testing for structural breaks with an *a priori* determined break date can be misleading, the choice of the methods seem adequate (Banerjee, Lumsdaine, and Stock (1992), Christiano (1992), Zivot and Andrews (1992)).

The advantage of using the Bai, et. al. method is that it tests for common breaks in multivariate time series (more precisely, it looks for the simultaneous break date in mean growth rates, treating autoregressive parameters as nuisance parameters). In doing so the procedure implements the "supremum" test of Andrews (1993) (i.e., Sup-Wald) and the related "average" and "exponential" tests of Andrews and Ploberger (1994) (i.e., SupF, ExpF, AveF tests)[25]. As shown by Bai, et. al., testing for simultaneous structural breaks in the VAR system improves estimation precision. Moreover, the authors construct confidence intervals for the break date that increase the estimation accuracy. The difference between the Bai, et. al., procedure and the procedure implemented in this study, is in the lag-selection method. This is necessary in order to obtain models consistent with those employed in the SVAR analysis[26].

The disadvantage of the Bai, et. al., method is that it is based on the asymptotic distribution theory. Although, the asymptotic distribution is relatively easy to tabulate, it may be unreliable in finite samples. Additionally, because the Bai, et. al. test uses asymptotic critical values, calculated under the null of i.i.d. errors, it can be inadequate in persistent or/and heteroskedastic series. Given the results of the normality and heteroskedasticity tests performed on the estimated VARs residuals, the likelihood of obtaining misleading results may not be insignificant. For example, Hansen (2000) finds that asymptotic distributions of Andrews' test statistics depend on the presence of a unit root and/or structural change in the regressors (i.e., they are not robust to structural change in the marginal distribution of the regressors) and thus the stationarity assumption underlying those tests may result in inadequate inference. Also, Diebold and Chenn (1996) provide evidence of size distortions (i.e., tendency to over-reject) of supremum tests for a structural change in dynamic models. This poses a problem in testing conditional relationships, since these tests cannot differentiate between structural change in conditional and marginal distributions. As such, they are not of much use to policymakers. For example, the marginal model can be thought of as an instrument that can be moved in order to achieve some goal (i.e., expressed by the conditional model). For policy purposes, of interest is the question whether the conditional model has invariant parameters, despite changes in the marginal model. In the context of this study, the question is whether the parameters of the estimated VAR equations are stable, despite changes in the exchange rate regimes, changes to the rate of growth of money, etc.

To this end Hansen proposes the 'fixed regressor bootstrap' which allows for arbitrary structural change in the regressors, including the lagged dependent variable and heteroskedastic error process. He further shows that this bootstrap technique produces the correct asymptotic distribution and also leads to reasonable size properties in finite samples. Therefore, in this study, the results of structural break tests obtained by implementing the Bai et.al. procedure are contrasted with those obtained by implementing Hansen's bootstrap technique[27].

[25] Andrews (1993) and Andrews and Ploberger (1994) provide critical values for SupF, ExpF, and AveF tests. Hansen (1997) calculates p-vales for those tests. His Gauss program is available at http//www.ssc.wisc.edu/~hansen/progs/progs.htm

[26] Bai et. al.'s Gauss program was modified to select the break date on the basis of the minimum AIC value as opposed to BIC value preferred by the authors.

[27] Given that results of the unit root tests may not robust, and the potential lack of normality of the data, as well as the presence of heteroskedasticity in some cases, performing structural change tests due to Hansen is also helpful in detecting whether the data used to estimate the conditional models are stationary (i.e., the evidence of structural break based on Andrews and Andrews-Ploberger's p-values which cannot be confirmed by the bootstrap method can be an additional evidence against stationarity of the data).

Test Results.

Based on the Bai, et. al. test, a break date in the mean at a common break date of real and nominal exchange rates is statistically significant for Slovakia (in 2000M11) and Spain (in 1994M11). For the rest of the countries, there is no evidence of shifts in the mean growth rates (Annex 3, Table A.3.2). The results obtained from the test due to Hansen are different (Annex 3, Table A.3.3). According to the p-values, there is evidence of coefficient instability in the Czech Republic, Hungary, and Slovenia in the NMSs group (in Poland only one tests appears to provide evidence for a possible break in the nominal exchange rate); and Portugal in the OMS group. However, in Hungary and Slovenia, once the tests are robust to the presence of heteroskedasticity, a potential structural break is confirmed only by the AveF test. The fact that estimated break dates cannot be confirmed by both tests is somewhat disturbing. Therefore, the CUSUM structural stability tests were also performed. A plot of cumulative sum of residuals did confirm parameters instability in the case of Czech Republic at the 5% level for both VAR equations[28]. In Slovenia, a plot of cumulative sum of squared residuals indicated parameter or variance instability in the nominal exchange rate equation. In the case of Hungary and Poland, neither CUSUM not CUSUM squared tests revealed the presence of a structural break. In Portugal, both tests strongly confirmed the presence of the structural break.

As documented by Diebold and Chenn (1996), the Bai at. al. test suffers from over-rejection, and since the CUSUM tests mostly confirmed the results obtained from Hansen's tests, the later two are treated as superior. To this end, the final VAR models include a shift dummy variable in the case of Czech Republic and Portugal (i.e., the dummy variable, respectively, equals one from 1999M3 and 1994M1 onwards). Since a heteroskedasticity corrected bootstrap did not confirm a structural break in the Slovenian exchange rate, it was concluded that the significance of the CUSUM squared test was due to the variance and not parameter instability. Moreover, the included dummy variables in the VAR equations were insignificant. No break was assumed for Slovenia. Similarly, given the outcome of the CUSUM tests for Hungary, and insignificance of the tested dummy variables in VAR equations, no break was assumed for Hungary.

Once estimated with structural breaks taken into account, the results of misspecification tests did not change much (Annex 3, Table A.3.1). Despite various attempts, normal and homoskedastic errors could not be obtained for countries in which these problems were initially detected[29]. Since the structural VAR form is derived from the reduced VAR representation (as a one-to-one transformation), the reliability of results from the structural analysis may be dubious. In order to mitigate normality issues, it is important to put some confidence on impulse responses and variance decompositions obtained from the SVAR models. Since, as a consequence of heteroskedastic errors, the structural shocks are not 'purely' exogenous and may depend on the values of variables in the system (i.e., the conditional variance of the nominal or real exchange rate may change with the past values taken by those rates), the White robust variance estimate for the errors is needed[30].

Small-Sample Bootstrap Confidence Intervals.

Kilian (1998a, 1998b) shows that if the innovations in a VAR system are not normally distributed, standard methods of generating confidence intervals for impulse responses - such as those proposed by Lütkepohl (1990) or Sims and Zha's (1995) - bring unsatisfactory results.

[28] The results can be obtained from the author upon request.
[29] In the cases of the Czech Republic, Slovakia, and Slovenia, dummy variables were also tested for periods of Asian, Czech, and Russian financial/banking crises, as well as for regime changes identified by Reinhardt and Rogoff.
[30] The Gauss programming language was used to obtain results presented in this and the next sections.

Following this approach, the bootstrap-after-bootstrap method is implemented. In Kilian's bootstrapping technique, the non-normality of VAR innovations is accounted for through adjustments for the bias in the OLS coefficient estimates of the VAR system. The bias term in the original OLS estimator is approximated by the following procedure:

1) standard nonparametric bootstrap methods are applied to draw 1000 realisations of $\hat{C}^{(i)}$ from the estimated VAR (p) models (i.e., equation (8));
2) then, the bias term $bias = E[\hat{C} - C]$ is approximated by $\hat{bias} = 1/1000 \sum \hat{C}^{(i)} - \hat{C}$;
3) next, stationarity correction is applied if the bias-corrected estimates imply that the VAR becomes non-stationary;

Once the stability conditions are satisfied, the biased corrected coefficients are used to generate 2000 new bootstrap replications of $\hat{C}^{(i)}$. These bias-corrected estimates are next used to compute the empirical distribution of impulse responses. Confidence intervals on impulse responses are constructed using modified percentile method of Davidson and McKinnon (1993). The same, biased corrected coefficients are used to calculate confidence intervals for variance decompositions[31].

The nonparametric standard bootstrap method proposed in Step 1 draws on Runkle (1987) - i.e., it generates bootstrap innovations e_t^* by resampling with replacement from the empirical residuals e_t. Pseudo-data Δy_t^* is constructed with the use of VAR (p) coefficients and is conditional on the vector of initial observations $\Delta y_0^* = \{\Delta y_1^*, ..., \Delta y_{25}^*\}$, which are selected randomly with replacement from the original VAR residuals (see Berkowitz and Kilian (1997)). These initial observations are then discarded so that the final pseudo-sample equals to Δy_t^*.

Additionally, each bootstrap loop takes into account the lag order uncertainty resampling by choosing the number of lags in each draw by minimising the AIC criterion. As showed by Kilian (1998b, p. 545) failing to do so leads to misleading inference - i.e. ignoring the lag order uncertainty may seriously undermine the coverage accuracy of bootstrap confidence intervals for impulse responses[32]. His results further suggest that in small and moderate samples the coverage accuracy of bootstrap confidence intervals for VAR impulse response estimates is much closer to the nominal coverage for the AIC criterion than it is for more parsimonious criteria.

White Robust Variance Estimate for the Errors.

As discussed, not taking into account heteroskedasticity in the structural variance decomposition results in bias in the relative importance of random innovations in the forecast error (i.e., it is influenced by past singular events). Since the purpose of the variance decomposition is to identify the importance of shocks which hit the economy regularly and within constant economic structures, it is important to correct for the presence of heteroskedasticity before such structural inference can be conducted. Therefore, in countries where better specification of VAR models could not be achieved, in order to obtain the White robust variance estimate for the errors, e_t,

[31] See Kilian (1998a), p.220 for details.
[32] Because the lag order uncertainty is taken into account, the short-cut proposed by Kilian in step 2a (1998a, p.220) could not be taken. Additional 2000 loops had to be estimated.

each of the equations in a reduced form VAR system is estimated by WLS instead of OLS (in order to obtain accurate confidence intervals, WLS VAR estimation also replaces OLS in the above described bootstrapping).

In the WLS estimation conducted for the Czech Republic, Italy, Portugal, Slovakia, and Spain the previously heteroskedastic residuals from the estimated reduced form VAR models turned out to be i.i.d.. Unfortunately, in the case of Slovenia the appropriate weights could not be found. Given that single digit inflation in Slovenia was reached only in 1996, the sample was set to start in January 1996. Since this time the Whites' tests were not found to be significant, the final structural analysis for Slovenia was performed on the sample spanning from 1996M1 to 2006M12.

Variance Decomposition.

Once the reduced form VAR models were correctly specified, the structural Blanchard and Quah decomposition was executed. Table 4.2.1 shows the contributions of temporary (i.e. nominal) and permanent (i.e. real) shocks to explaining the forecast error variance of nominal and real exchange rates in the Czech Republic, Hungary, Poland, the Slovak Republic and Slovenia. Column I and III in Table 4.2.1 reflect contributions of the real and nominal shocks, respectively, to the forecast variance error of the real exchange rate. Columns V and VII contain the contributions of the same shocks to movements of the nominal exchange rate. Finally, the numbers in columns II, IV, VI and VIII represent the bootstrapped confidence intervals calculated for a particular percentage of variance decomposition.

The results are striking. In the case of Hungary, the Slovak Republic, and Slovenia over 90% of shocks to the real exchange rate are real in nature. In the Czech Republic real shocks explain as much as 88% of the forecast variance error of the real exchange rate. Poland is somewhat different, with a nominal shock playing a substantial role in the variation of its real exchange rate (in the first month, the extent is 40%). Nevertheless, after a year, the significance of the nominal shock drops significantly, with real shocks explaining approximately 80% of the forecast variance error of a Polish real exchange rate.

The dominance of real shocks in the real exchange rate fluctuations confirms the finding of the unit root tests which suggest that real rates are not stationary.

Variance decomposition of nominal exchange rates is more heterogeneous. Nominal shocks overwhelmingly dominate the variation of the Slovenian tolar (over 80% of movements are due to this type of shocks), and are an important part of the volatility of the Polish zloty (around 50% irrespective of the forecast horizon), the Czech koruna (on average 30% within the first three months), the Hungarian forint (which despite a very minimal initial impact, after a year increases to around 20% and higher) and to some extent the Slovakian koruna (which remains around 15% after a year).

Table 4.2.1 Variance Decomposition (NMSs)

Variable	RER				NER			
Variance Decomposition	RER shock		NER shock		RER shock		NER shock	
CZECH REP.	**I**	**II**	**III**	**IV**	**V**	**VI**	**VII**	**VIII**
1-month	87.8	60.2-100	12.2	0.0-39.8	62.8	23.1-100	37.2	0.0-76.9
3-month	91.9	72.0-100	8.1	0.0-28.0	70.1	35.3-99.7	29.9	0.3-64.7
12-month	95.0	82.1-100	5.0	0.0-17.9	76.0	47.5-99.6	24.0	0.4-52.5
24-month	97.5	90.3-100	2.5	0.0-9.7	80.9	58.7-97.6	19.1	2.4-41.3
60-month	99.5	98.0-100	0.5	0.0-2.0	86.0	71.0-99.0	14.0	1.0-29.0
HUNGARY								
1-month	99.5	58.6-100	0.5	0.0-41.4	96.9	38.0-100	3.1	0.0-62.0
3-month	99.8	57.6-100	0.2	0.0-42.4	93.6	33.0-100	6.4	0.0-67.0
12-month	99.5	69.1-100	0.5	0.0-30.9	83.0	23.7-99.5	17.0	0.5-76.3
24-month	99.7	80.2-100	0.3	0.0-19.8	72.4	18.2-98.9	27.6	1.1-81.8
60-month	99.8	91.5-100	0.2	0.0-8.5	60.8	7.8-94.4	39.2	5.6-92.2
POLAND								
1-month	60.6	26.1-100	39.4	0.0-73.9	47.1	15.0-100	52.9	0.0-85.0
3-month	69.1	36.1-100	30.9	0.0-63.9	53.1	20.1-100	46.9	0.0-79.9
12-month	82.9	64.8-100	17.1	0.0-35.2	49.9	24.6-98.2	50.1	1.8-75.4
24-month	90.4	81.2-100	9.6	0.0-18.8	45.7	23.6-94.2	54.3	5.8-76.4
60-month	96.3	93.5-100	3.7	0.0-6.5	40.4	25.0-95.7	59.6	4.3-75.0
SLOVAK REP.								
1-month	100	86.2-100	0.0	0.0-13.8	74.1	41.0-100	25.9	0.0-59.0
3-month	99.6	88.5-100	0.4	0.0-11.5	78.4	48.0-100	21.6	0.0-52.0
12-month	99.9	96.5-100	0.1	0.0-3.6	84.4	63.1-97.0	15.6	3.0-36.9
24-month	100	98.2-100	0.0	0.0-1.8	85.2	66.5- 97.3	14.8	2.7- 33.5
60-month	100	99.3-100	0.0	0.0-0.7	85.7	67.3 97.2	14.3	2.8- 32.7
SLOVENIA								
1-month	95.4	38.9-100	4.6	0.0-61.1	14.9	0.0-81.6	85.1	100-18.4
3-month	89.9	39.3-100	10.1	0.0-60.7	22.2	0.0-87.0	77.8	100-13.0
12-month	92.9	56.7-99.9	7.1	0.1-43.3	8.1	3.3-65.2	91.9	96.7-34.8
24-month	95.5	74.6-99.9	4.5	0.1-25.4	3.8	0.9-61.9	96.2	99.1-38.1
60-month	98.0	90.3-100	2.0	0.0-9.7	4.3	0.3-66.9	95.7	99.7-33.1

Source: Author's estimation based on IMF IFS and EUROSTAT data.

Table 4.2.2: Test of Long-Run Over-identifying Restrictions

	L_BAND	D(2,2)	U_BAND
Czech Rep.	0.18	**0.40**	0.50
Hungary	0.74	**3.84**	4.24
Poland	0.61	**2.02**	2.43
Slovakia	0.13	**0.67**	0.87
Slovenia	0.16	**0.72**	0.46

Note: Columns L_Band and U_Band stand for the lower and upper band, respectively, of 90% bootstrapped confidence intervals.

Source: Author's estimation based on IMF IFS and EUROSTAT data.

With regard to the OMSs (Annex 4, Table A.4.1), there is no doubt that real shocks are responsible for real exchange rate movements in all countries but Spain, where at the 3-month forecast horizon, a nominal shock explains 10 percent of real exchange rate volatility. Real shocks also move nominal exchange rates in Germany and France and to the lesser extend in Italy. For Greece, the nominal shock is persistent and amounts to around 30 percent of the nominal exchange rate volatility for all forecast horizons. For Portugal it remains at the 20 percent level. For Spain it drops to around 15 percent already after a year. Interestingly, despite the fact that all countries included in the OMS group except for Germany adopted some form of *de facto* pegged exchange rate regime to the DM, relatively little distortion caused by that fact seems to arise. The temporary component in the real exchange rate forecast error variance in all the countries is virtually nonexistent.

Impulse Responses.

Overall the shocks seem to be well identified. As required by the identification assumption, in all cases, the impact of the nominal shock on the real exchange rate is temporary (see Figure 1). Testing the hypothesis that a positive nominal shock has a proportional, long-run, effect on a nominal exchange rate (i.e., imposing the restriction D(2,2)=1) brought mixed results, but overall showed that a positive nominal shock leads to currency depreciation (i.e., D(2,2) is always positive, see Table 4.2.2).

This effect is less than proportional in the case of the Czech Republic, Slovakia and Slovenia. Despite the fact that the overshooting effect is present in Poland and Hungary, the bootstrapped 90% confidence bands are rather wide. Finally, a positive real shock, in all countries, causes long-run nominal **a**nd real exchange rate depreciation (perhaps with the exception of Slovenia, where the real exchange rate appreciates over time). This brings some evidence against the ‘exchange rate disconnect’ theory. The remainder of this section deals with individual country cases[33]:

Czech Republic

In the Czech Republic, following a real shock, the real exchange rate jumps in the same direction but more than the nominal rate. After the initial jump, both rates return to their long-run values within eleven and nine months, respectively. Given that the response of the real rate is larger than the response of the nominal rate indicates that permanent changes to the real exchange rate occurs through the nominal rate, but also through the relative price. The fact that the nominal rate depreciates less in response to a real shock than the real rate does, indicates that - in response to positive real shocks - domestic prices go down. One possible explanation of this outcome is that real shocks are supply-side shocks.

Along with the imposed identification restriction (D(1,2)=0), a nominal shock has no long-run effect on a real exchange rate. In the short-run, the real exchange rate increases, but this effect is not large and approaches zero in less than a year. In response to the nominal shock, the nominal exchange rate jumps away from zero, but the jump is of a very small and short-lasting magnitude, casting doubts about its significance. Finally, because nominal shocks have no long-run effects on real rates, and since nominal shocks do affect prices to some degree, nominal shocks have to affect prices by an equal but opposite amount.

[33] Impulse responses for the OMSs, as they are of little relevance for the study, are available upon request.

Hungary

In Hungary, real shocks cause a long-term depreciation of both real and nominal exchange rates, indicating that the nominal exchange rate does absorb shocks which cause real exchange rate movements. The fact that the nominal rate depreciates more in response to a real shock than the real rate does, indicates that - in response to positive real shocks - domestic prices go up. One possible explanation of this outcome is that real shocks are demand-side shocks. In response to a nominal shock, after the initial appreciation, the real exchange rate comes back to the zero line – as predicted by the identification restriction (the effect does not last more than a year). On the other hand, following the nominal shock, the nominal exchange rate depreciates substantially and this depreciation is <u>permanent</u>.

Poland

In Poland, as in previous cases, a real shock causes real and nominal exchange rates to depreciate. The adjustment path for the real rate nevertheless differs from that of the nominal rate. Following the real shock, the real rate goes through a brief period of revaluation relative to the initial depreciation. After a year a steady depreciation is observed. In the case of the nominal rate, the final depreciation is lower than the initial response. The fact that the real exchange rate depreciates more than the nominal rate in response to the real shock, suggests that a positive real shock causes domestic prices to decline and thus improves the country's price competitiveness (as in the case of Czech Republic).

In response to a nominal shock a real exchange rate overshoots its long-run value. The impact is long lasting. Were nominal shocks monetary in nature, this would indicate that monetary policy can influence both real and nominal exchange rates. Nevertheless, in the short-run, the response of the real exchange rate to a real shock is greater.

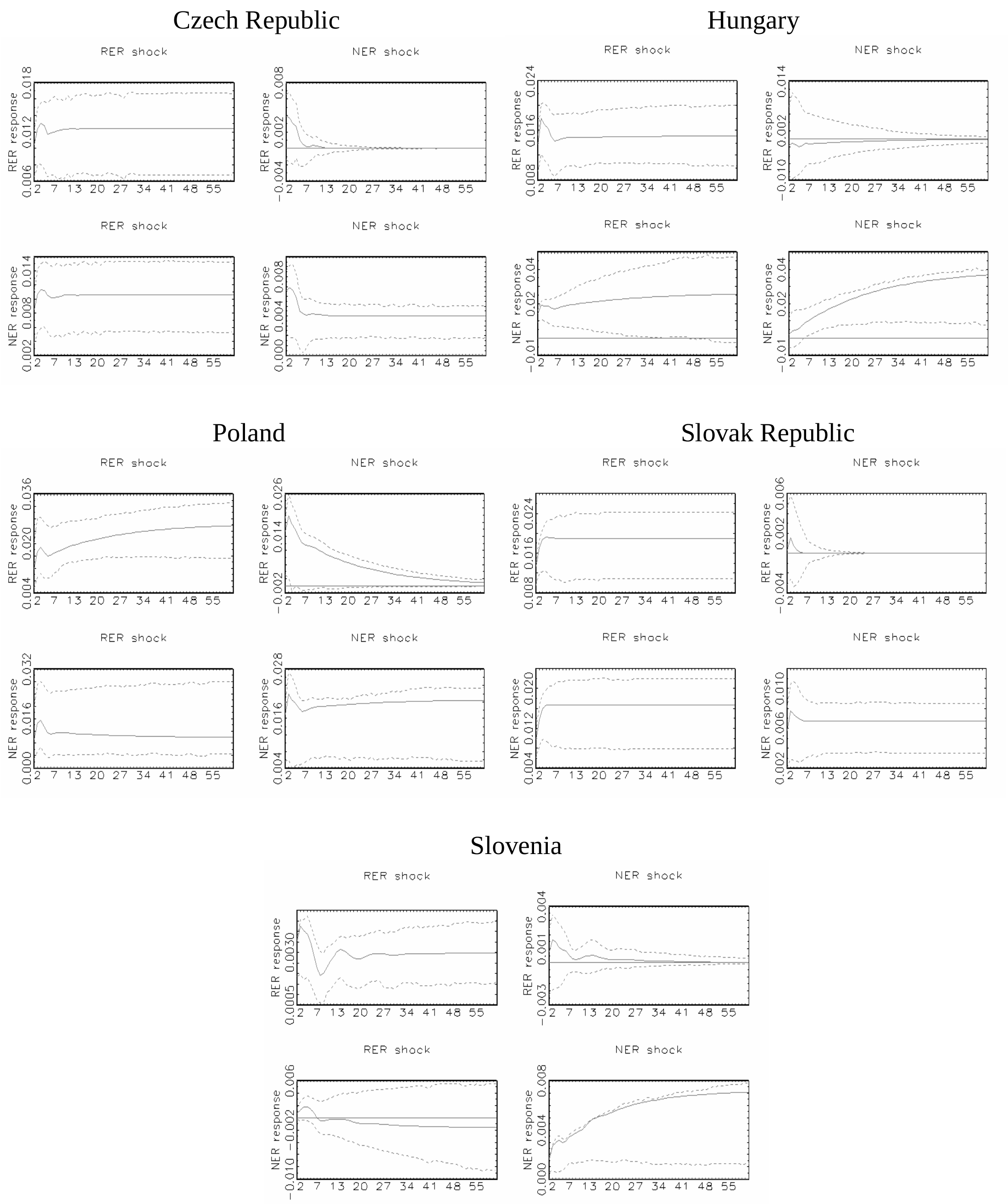
The same is not true for the nominal exchange rate. The response of the Polish zloty to a nominal shock is greater than it is to a real shock; there is also evidence of moderate 'overshooting' (although less so than in Hungary). However, the nominal exchange rate reaction function indicates that the nominal exchange rate achieves its new long-run level in less than a year.

The Slovak Republic

An interesting feature of the impulse response analysis conducted for the Slovak Republic is that, although the magnitudes are somewhat different, the shapes of reaction functions are the same for the real and nominal exchange rates in response to a real and nominal shock, respectively. In the short-run, the real exchange rate depreciates more in response to the real shock than the nominal exchange rate does. Again, as in cases of Czech Republic and Poland, this indicates an improvement in the country's price competitiveness.

Looking at the effects of nominal shocks on real and nominal exchange rates, it is hard to see any significant evidence of overshooting. Even if both rates do move in response to nominal shocks affecting them, the jump is less than proportional (this is confirmed by the 90% confidence bands) and relatively small.

Figure 1: Impulse Responses (NMSs)



Note: RER – real exchange rate, NER - nominal exchange rate. The top two, out of four, panels presented for each country represent impulse responses of the RER to a unit of real and nominal shocks, respectively; the bottom two panels are impulse responses of the NER to a unit of real and nominal shocks, respectively.

Source: Author's estimations based on IMF IFS and EUROSTAT data.

Slovenia

In Slovenia, the impulse response functions of real and nominal exchange rates are different depending on the shock hitting the economy. In response to a real shock, initial fluctuations of the real exchange rate are observed. The long-run value is only modestly higher when compared with the 'starting' level. As the real shock causes a long-run moderate appreciation of the nominal exchange rate, and the real exchange rate is almost constant, this suggests that the nominal rate has been moving in the opposite direction to the relative price. In response to a nominal shock, the real exchange rate jumps above its long run value, but the jump does not seem to be significant. The nominal exchange rate does not jump, but rather depreciates steadily. The fact that the nominal shock dominates nominal exchange rate movement is not surprising given the discretionary exchange rate policy in Slovenia. Overall, the scale of impulse responses in Slovenia, with the exception of the response of the nominal rate to the nominal shock, is minimal.

Robustness Checks.

Since the Faust and Leeper (1997) critique relating to the potential invalidity of imposing long-run restrictions to the finite sample is particularly important for VAR models with a large lag order, it is important to check the robustness of the performed structural VAR analysis. As showed by Lastrapes (1998) the robustness checks can be performed by re-estimating the bivariate SVAR model with the identifying restrictions imposed at different finite horizons. Figure 5, Annex 4, sets out the results of this analysis, and shows that impulse responses change very little in terms of original dynamics. Therefore, the structural VAR analysis can be said to be robust to the Faust and Leeper's critique, and that the horizon of 60 months can sufficiently approximate the long-run[34].

PART 5: CONCLUSIONS

Based on the results obtained from the univarite variance analysis as well as the structural VAR, the following comments and conclusions can be drawn.

As the estimates show, based on the proposed definition of real convergence, to lessen the costs of the eurozone membership, the levels of volatility in the NMSs will have to be reduced (with the exception of Slovenia[35] and Estonia). This is because, average volatility for the NMSs is currently 1.9 times higher than the average volatility of the selected OMSs analysed in this paper in the mid-1990s (1.5 higher than it was for the average of Club Med countries and 3.5 times higher than the average estimated for France and Germany). The plots of the estimated quarterly time varying conditional variances for real exchange rates seem to support these findings.

The fact that real shocks dominate real exchange rate movements implies that the univariate analysis undertaken in this paper provides an accurate measure of real exchange rate variance (at least for countries for which VAR analysis was possible). Given the moderate impact of nominal shocks to real exchange rate variability at the 3-month forecast horizon, it is not implausible to assume that monthly and quarterly volatility changes represent the 'true' magnitude of real convergence in these countries. However, in the case of Poland, monthly and quarterly real

[34] The results for the OMSs are available upon request.

[35] As Slovenia is already in the eurozone, it can be said that – based on the proposed indicator – it was ready to give up its monetary and exchange rate policy and should benefit from the euro adoption.

exchange rate volatility should be scaled down by approximately 40 and 31 percent, respectively; in the case of the Czech Republic by 12 and 8 percent. Likewise, because the variance decomposition for the Club Med countries as well as France and Germany shows that real shocks dominated real exchange rates in these countries, the benchmark magnitudes for real convergence can be treated with confidence.

Although at the time of writing Estonia does not fulfil the nominal Maastricht inflation criterion, it seems that – based on the proposed definition – it could benefit from full integration into the EMU. Given that Estonia has managed to cope without monetary independence for over ten years, and given its comparatively flexible labour and product markets, Estonia could benefit from the euro adoption[36]. As for the rest of countries, more effort is needed in reducing idiosyncratic real shocks. The recent recommendation of the European Commission that Slovakia be allowed to adopt the euro as of January 2009 is not supported by the proposed indicator for real convergence. Giving up monetary independence may be premature because existing asymmetries are not insignificant, and in fact are higher than those observed in the Czech Republic, Latvia and Lithuania (in the third sub-sample). Finally, analysis in the case of Poland, once real exchange rate volatility is scaled down for the presence of nominal shocks, imply that the country has achieved a level of real convergence with the eurozone comparable to other countries in the NMSs' group[37].

The structural VAR analysis showed that in all countries except for Slovenia, for which it was possible to estimate VAR models, the nominal exchange rate <u>does move</u> in the same direction as the real exchange rate at the onset of a real shock, and thus <u>does play</u> a shock-stabilising role. Therefore, loss of the nominal exchange rate as an adjustment instrument would represent a cost of the euro area membership.

In the case of Poland (and to a lesser degree in the Czech Republic), the importance of nominal shocks in the real exchange rate forecast error variance does not indicate that there is no cost in loosing the nominal exchange rate as an adjustment instrument. This is because nominal shocks seem not to be destabilizing (similarly to other investigated countries), indicating that there is still some room for short-run effectiveness of monetary and exchange rate policy in changing the real exchange rate in this country. However, it is also true that more stable monetary policy could result in a greater real exchange rate stability[38].

The results of the B&Q variance decomposition show that the nominal component of nominal exchange rate movements in the five countries included in the SVAR modelling is not insignificant (perhaps to a lesser degree in the case of Slovakia)[39]. Therefore, elimination of these movements could be a positive benefit of speedy accession to the euro area. However, given the need to fulfil Maastricht criteria first, and the relatively significant degree of real asymmetry, there may be risks to premature ERMII participation (i.e., due to increased capital flows or consumption booms). Additionally, significant nominal exchange rate responses to real shocks in

[36] Moreover, the policies needed to target current imbalances, in the currency union, would not be different to those required under a currency board. At the same time Estonia could enjoy the benefits of the common currency.

[37] In the case of Poland, without the SVAR analysis, the univariate approach would produce distorted results.

[38] The possibility of effectiveness of monetary and exchange rate policies in changing Polish competitiveness was also concluded by Diboolu and Kutan (2001). However, in their study, a nominal shock dominates a real exchange rate forecast error variance to a greater degree.

[39] The nominal exchange rate volatility in the OMSs was mainly driven by real factors (to the lesser degree in Spain and Greece).

all five NMSs except for Slovenia need to be analysed and taken into account in assessing nominal exchange rate stability at the end of the ERMII period[40].

Finally, given (i) that real asymmetric shocks are not insignificant when compared with the Club Med countries, (ii) the stabilising role of nominal exchange rates (with the exception of Slovenia), and (iii) the fact that nominal shocks, on average, do not move real exchange rates, the NMSs may be well advised to take their time in joining the eurozone (except for Estonia and Slovenia (ex post)). In the interim they may do well to concentrate on enhancing structural reforms, until they are ready to give up monetary and exchange rate independence.

[40] Given the broad classification of shocks, at this stage, without further identification of shocks, it is impossible to investigate the split between supply and demand shocks.

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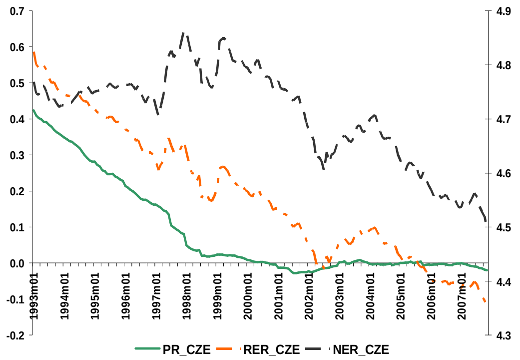
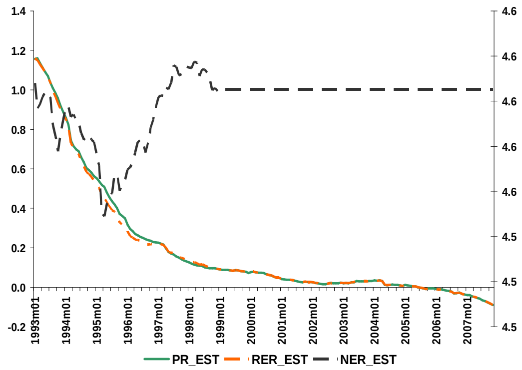
Annex 1: Graphical Presentation and Integration Properties

Figure 2: Exchange Rates and Price Ratio Developments in the NMSs

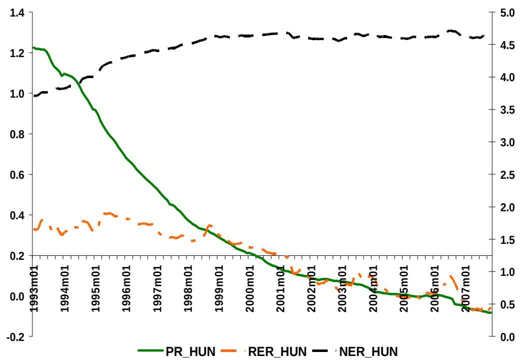
Czech Republic



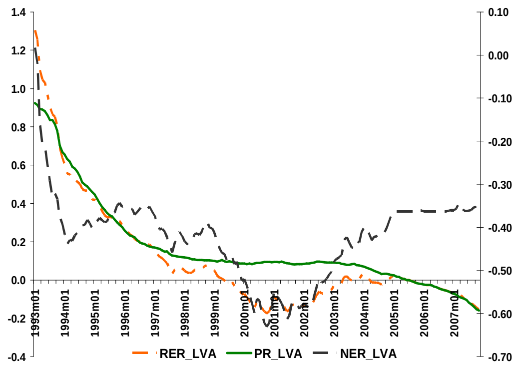
Estonia



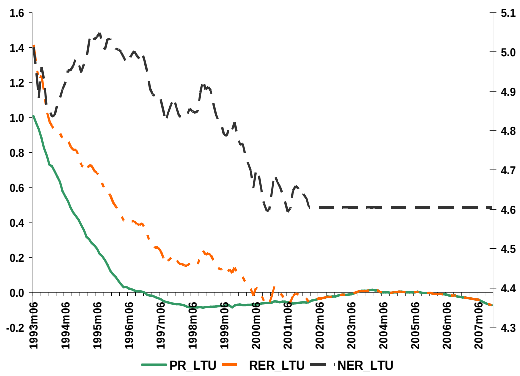
Hungary



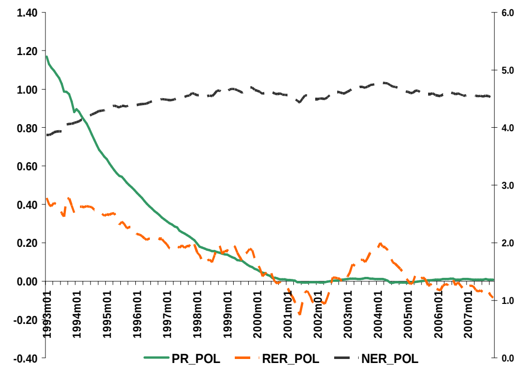
Latvia



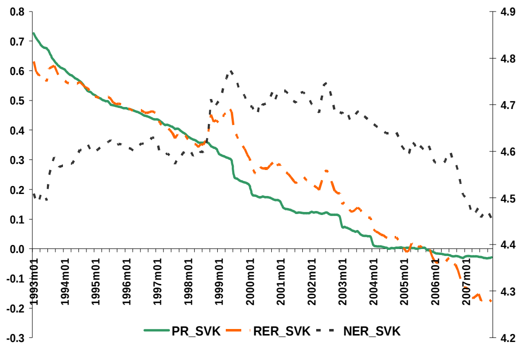
Lithuania



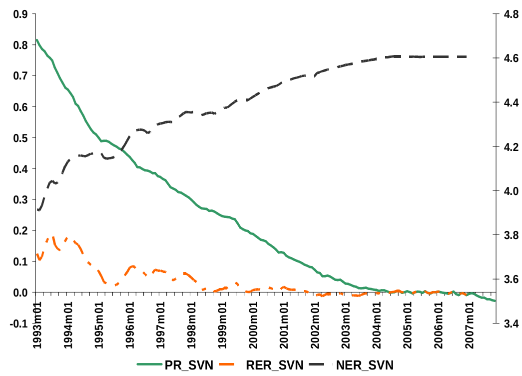
Poland



Slovak Republic



Slovenia



Source: Author's calculation based on data from the Eurostat and IMF IFS. The magnitudes of relative prices and real exchange rates' indices are on the LHS axis; the magnitudes of nominal exchange rates' indices on the RHS axis. Indices are defined with the base of 2005=100. Decrease in the exchange rate index indicates appreciation.

Table A.1.1 Properties of Real Exchange Rates in a Data (NMSs)

FULL SAMPLE	Czech Rep.	Estonia	Hungary	Latvia	Lithuania	Poland	Slovak Rep.	Slovenia
Persistance								
Nominal Exchnage Rate	0.19	0.27	0.36	0.47	0.18	0.33	0.27	0.70
t-stat	2.58	3.83	5.11	7.11	2.42	4.66	3.73	12.86
Real Exchange Rate	0.14	0.69	0.24	0.53	0.37	0.27	0.22	0.53
t-stat	1.94	12.68	3.30	8.43	5.59	3.71	3.02	8.28
Cross-Correlations								
Real and Nominal Exchnage Rates	0.94	0.31	0.92	0.91	0.83	0.91	0.89	0.77

Source: Author's calculation based on IMF IFS and EUROSTAT data.

Table A.1.2 Unit Root Tests

	DF-GLS			MZt			MZa			KPSS		
	NER	RER	PR	NER	RER	PR	NER	RER	PR	NER	RER	PR
CZE	-0.9	-1.6	-1.1	-0.9	-1.7	-1.4	-2.4	-5.8	-4.8	…	…	…
EST	…	-2.1	…	…	-1.0	…	…	-2.4	…	…	…	…
HUN	-0.9	-1.8	-0.9	-1.1	-1.8	-1.5	-3.2	-6.5	-5.5	0.4	…	…
LTU	…	-0.3	…	…	-0.1	…	…	-0.1	…	…	…	…
LVA	…	-0.4	…	…	-0.4	…	…	-0.5	…	…	…	…
POL	-0.2	-1.7	-1.7	-0.2	-1.7	-0.8	-0.3	-5.8	-2.0	…	…	…
SVK	-0.5	-2.6	0.1	-0.5	-2.5	0.1	-1.3	-13.8	0.3	…	…	…
SVN	-0.2	**-3.0**	-0.5	-0.3	**-3.0**	-0.7	-0.5	**-18.1**	-1.8	…	0.3	…
FRA	-2.0	-2.2	-1.1	-1.9	-2.1	-0.9	-7.0	-8.7	-1.9	…	…	0.2
DEU	-1.5	-1.6	-1.6	-1.5	-1.6	-1.5	-4.6	-5.0	-4.7	…	…	…
ESP	-1.2	-1.4	-1.3	-1.0	-1.1	-0.6	-2.3	-2.7	-1.2	…	…	…
GRC	-1.7	-2.0	-0.4	-1.7	-2.1	-0.5	-6.3	-9.1	-1.1	…	…	…
ITA	-1.5	-1.5	-0.8	-1.5	-1.5	-0.8	-5.0	-4.9	-1.8	…	…	…
PRT	-1.1	-1.2	-1.5	-0.8	-0.9	-1.5	-1.5	-2.1	-4.3	…	…	…

Note: NER - nominal exchange rate; RER – real exchange rate; PR – price ratio. Bolden magnitudes indicate unexpected results at the 5 per cent significance level; i.e., stationarity of exchange rate or/and price ratio series in levels; kmax=int(12*(T/100)^(0.25)).

Source: Author's calculation based on IMF IFS and EUROSTAT data.

Table A.1.3 Unit Root Test with a Break

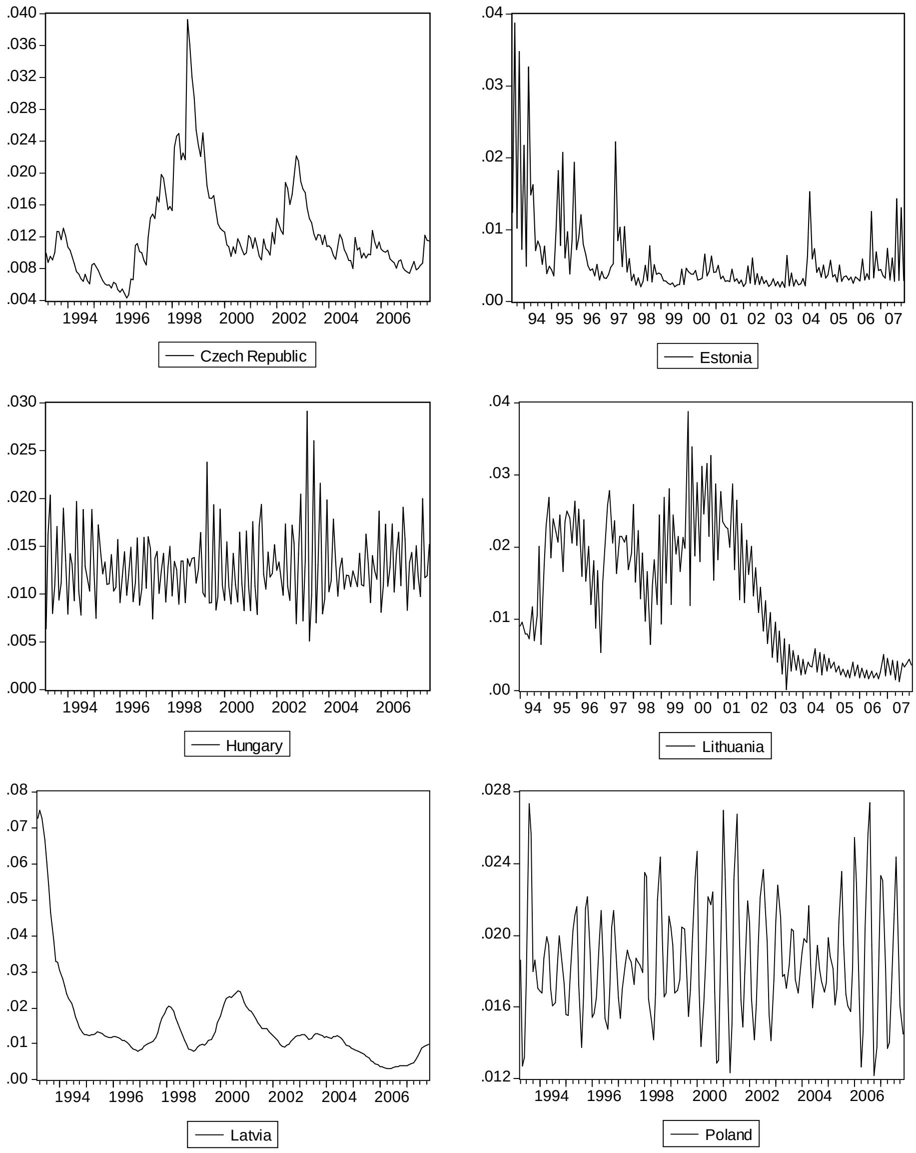
Perron (1997)	NER								NER					
	_CZE	_EST	_HUN	_LTU	_LVA	_POL	_SVK	_SVN	_FRA	_DEU	_ESP	_GRC	_ITA	_PRT
Break point	77.0	n/a	64.0	n/a	n/a	54.0	109.0	110.0	65.0	65.0	14.0	24.0	29.0	14.0
Dummy_coeff.	0.0	n/a	0.0	n/a	n/a	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0
T-stat	-22.2	n/a	-53.2	n/a	n/a	-23.7	-27.9	-17.9	0.5	0.6	-12.5	-9.3	-10.6	-10.3
Trend_coeff.	0.0	n/a	0.0	n/a	n/a	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0
T-stat	9.4	n/a	72.1	n/a	n/a	29.6	23.1	66.0	-0.8	3.1	15.0	15.7	11.0	11.8
Fixed lag	10.0	n/a	11.0	n/a	n/a	11.0	1.0	7.0	0.0	7.0	7.0	5.0	8.0	9.0
y(a-1)	-0.1	n/a	-0.2	n/a	n/a	-0.1	-0.1	-0.1	-0.1	-0.1	-0.4	-0.3	-0.2	-0.2
ADF	-3.0	n/a	**-4.4**	n/a	n/a	-3.4	-3.7	-5.1	-2.2	-1.7	-3.7	-3.4	-2.7	-2.9
	RER								**RER**					
Break point	161.0	161.0	33.0	157.0	161.0	161.0	113.0	151.0	65.0	65.0	14.0	14.0	28.0	14.0
Dummy_coeff.	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0
T-stat	2.6	3.6	-11.2	4.6	3.1	1.9	-13.0	4.4	0.1	0.4	-9.3	0.9	-9.3	-6.5
Trend_coeff.	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0
T-stat	-56.9	-18.6	4.9	-20.5	-15.0	-18.7	-46.2	-18.7	6.3	6.0	10.0	-2.0	7.5	6.1
Fixed lag	10.0	13.0	10.0	12.0	8.0	10.0	11.0	7.0	0.0	7.0	1.0	5.0	7.0	7.0
y(a-1)	-0.1	0.0	-0.1	0.0	0.0	0.0	-0.2	-0.1	-0.1	-0.1	-0.2	-0.1	-0.2	-0.2
ADF	-2.4	-2.2	-3.3	-1.7	-2.1	-2.5	-4.0	-3.1	-2.4	-1.7	-3.1	-2.0	-2.2	-3.1
	PR								**PR**					
Break point	161.0	n/a	161.0	n/a	n/a	161.0	18.0	151.0	24.0	44.0	65.0	86.0	14.0	65.0
Dummy_coeff.	0.0	n/a	0.0	n/a	n/a	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0
T-stat	5.8	n/a	6.0	n/a	n/a	6.1	7.0	9.2	-15.4	-7.7	4.0	6.6	2.0	-1.1
Trend_coeff.	0.0	n/a	0.0	n/a	n/a	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0
T-stat	-23.0	n/a	-35.6	n/a	n/a	-23.1	-11.2	-52.2	29.0	17.4	-31.1	-39.7	-6.2	-27.6
Fixed lag	10.0	n/a	3.0	n/a	n/a	13.0	12.0	13.0	0.0	9.0	9.0	4.0	9.0	1.0
y(a-1)	0.0	n/a	0.0	n/a	n/a	0.0	0.0	0.0	-0.5	-0.3	-0.1	0.0	-0.1	-0.1
ADF	-2.1	n/a	-2.5	n/a	n/a	-2.6	-1.1	-2.6	**-4.9**	-4.2	-2.0	-2.7	-2.6	-2.6

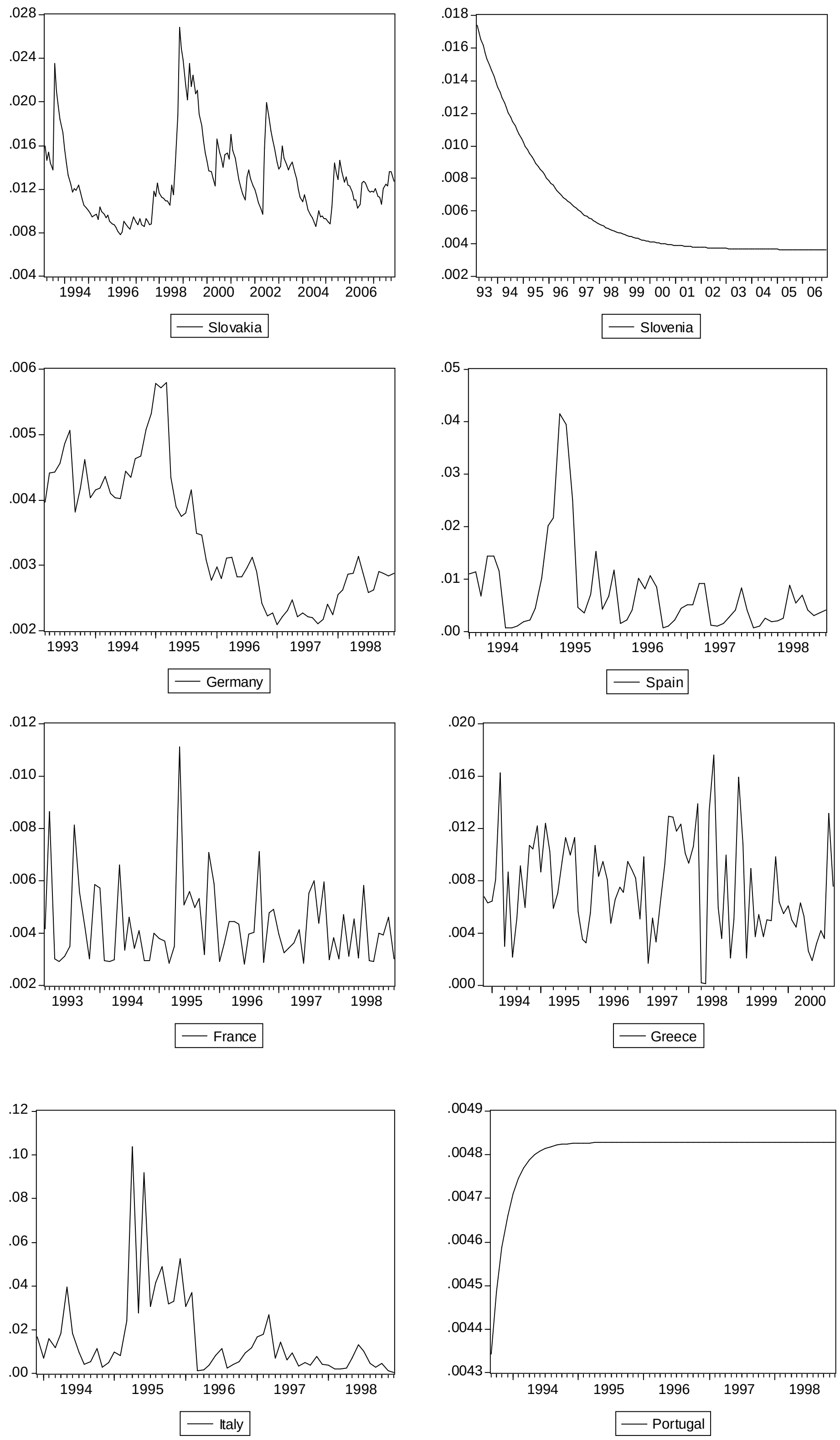
Note: NER - nominal exchange rate; RER – real exchange rate; PR – price ratio. kmax=int(12*(T/100)^(0.25)).

Source: Author's calculation based on IMF IFS and EUROSTAT data.

Annex 2: Time Varying Conditional Variances (Real Exchange Rates)

Figure 3: Time Varying Conditional Variances (NMSs, OMSs, Monthly)

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.004
1994 1996 1998 2000 2002 2004 2006
Slovakia
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93 94 95 96 97 98 99 00 01 02 03 04 05 06
Slovenia
.006
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1993 1994 1995 1996 1997 1998
Germany
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.03
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1994 1995 1996 1997 1998
Spain
.012
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1993 1994 1995 1996 1997 1998
France
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1994 1995 1996 1997 1998 1999 2000
Greece
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1994 1995 1996 1997 1998
Italy
.0049
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.0045
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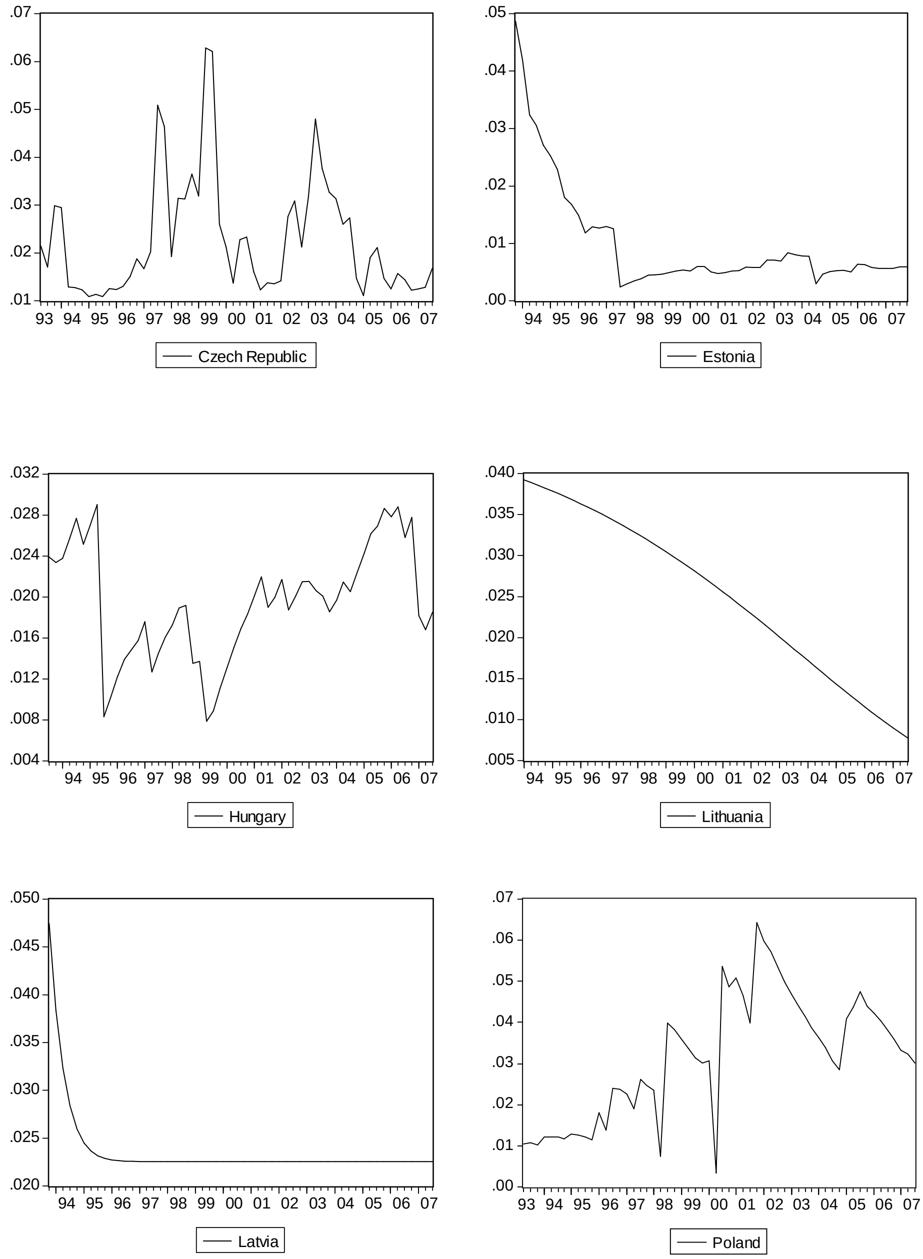
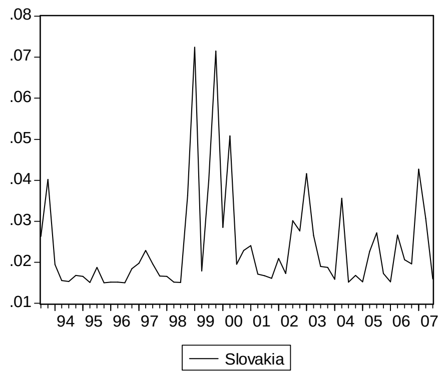
1994 1995 1996 1997 1998
Portugal

Figure 4: Time Varying Conditional Variances (NMSs and OMSs, Quarterly)

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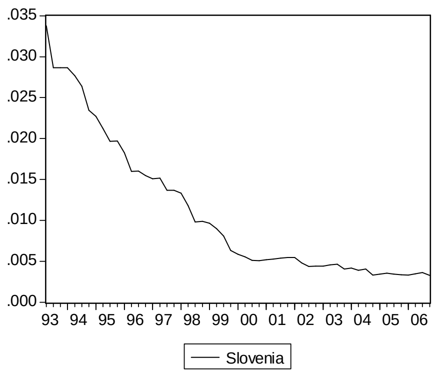
Slovakia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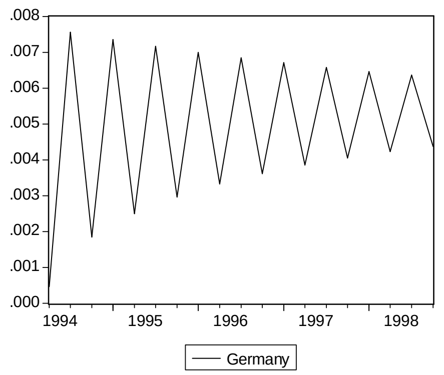
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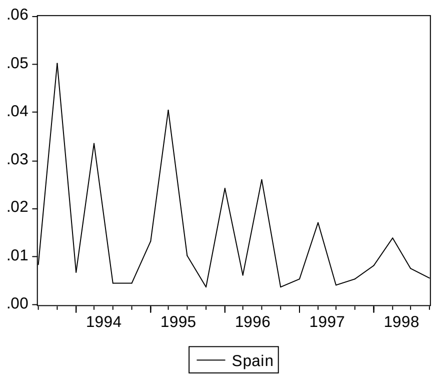
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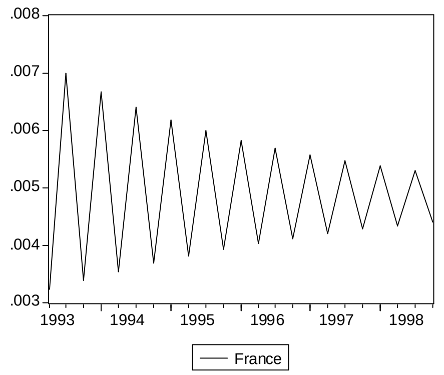
Spain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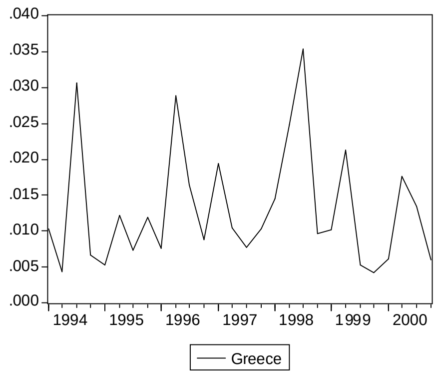
1993 1994 1995 1996 1997 1998
France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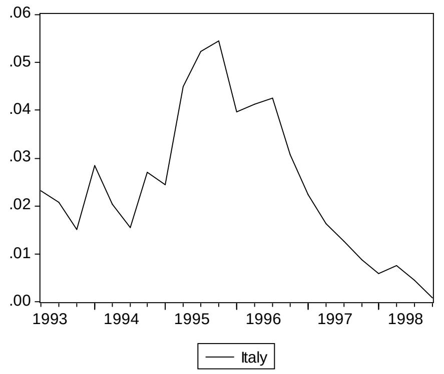
1994 1995 1996 1997 1998 1999 2000
Greece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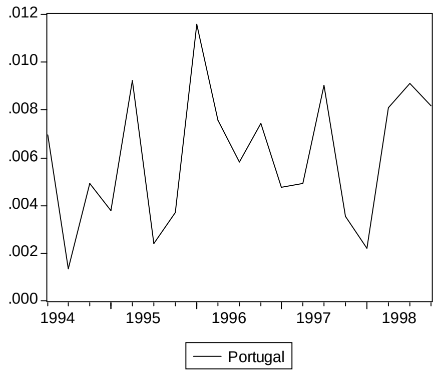
1993 1994 1995 1996 1997 1998
Italy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1994 1995 1996 1997 1998
Portugal

Annex 3: Model Specification and Checks

Table A.3.1 Misspecification Tests

Country	Lags	Normality	Autocorrelation		White Hetero	
		Doornik-Hansen	Portmanteau	p-val	no cross terms	cross terms
CZE	6	**0.0000**	0.22		**0.00**	**0.00**
CZE_DUM	6	**0.0001**	0.07		**0.00**	**0.01**
HUN	6	**0.0000**	0.19		0.59	0.20
POL	3	0.1139	0.30		0.14	0.08
SVK	1	**0.0000**	0.10		**0.00**	**0.00**
SVK_93	7	**0.0121**	0.09		**0.00**	**0.00**
SVN_96	6	**0.0023**	0.39		**0.05**	**0.01**
DEU	1	**0.0000**	0.21		0.81	0.79
ESP	1	0.1965	0.61		**0.01**	**0.02**
FRA	1	0.5493	0.26		0.83	0.90
GRC	3	**0.0000**	0.22		0.62	0.58
ITA	3	**0.0000**	0.13		**0.00**	**0.00**
PRT	1	0.0970	0.33		**0.01**	**0.02**
PRT_DUM	1	0.6609	0.19		**0.00**	**0.02**

Note: The lines CZE_DUM and PRT_DUM present results from estimating models specified with the dummy variable detected by structural break tests; lines SVN_93 and SVN_96 present results from estimating the models with the data spanning form 1993M4 to 2006M12 and from 1996M1 to 2006M12, respectively. The column marked 'Lags' includes number of lags chooses by the AIC criterion for the final estimation of VAR models. Columns 'Normality', Autocorrelation' and 'White Hetero' present p-values attached to estimated tests; figures in bold indicate that the null hypothesis cannot be rejected at the 5 per cent significance level.

Source: Author's estimates based on IMF IFS and EUROSTAT data

Table A.3.2 Bai, et. al, Structural Break Test

Country	Sample	Lags	Sup-W-15%	Exp-W-15%	Est Break	90% Conf. Int.
CZE	1996:10-2007:11	6	0.58	0.80	_1999:3	(1998:12, 1999:6)
HUN	1993:8-2007:11	6	1.00	1.00	_1997:5	(1996:9,1998:1)
POL	1996:3-2007:11	3	0.15	0.19	_2000:3	(2000:1,2000:5)
SVK	1997:3-2007:11	1	**0.02**	**0.01**	_2000:11	(1999:8, 2002:2)
SVN	1996:8-2006:12	6	0.48	0.57	_2002:3	(2001:11, 2002:7)
FRA	1993:3-1998:12	1	0.45	0.41	_1995:3	(1993:8,1996:10)
DEU	1993:3-1998:12	1	0.25	0.35	_1994:11	(1993:10,1995:12)
ESP	1993:3-1998:12	1	**0.01**	**0.01**	_1994:11	(1994:6,1995:4)
GRC	1993:5-2000:12	3	0.19	0.10	_1996:2	(1995:8,1996:8)
ITA	1993:5-1998:12	3	0.61	0.74	_1996:8	(1996:3,1997:1)
PRT	1993:3-1998:12	1	0.87	0.84	_1994:2	(X,1995:8)

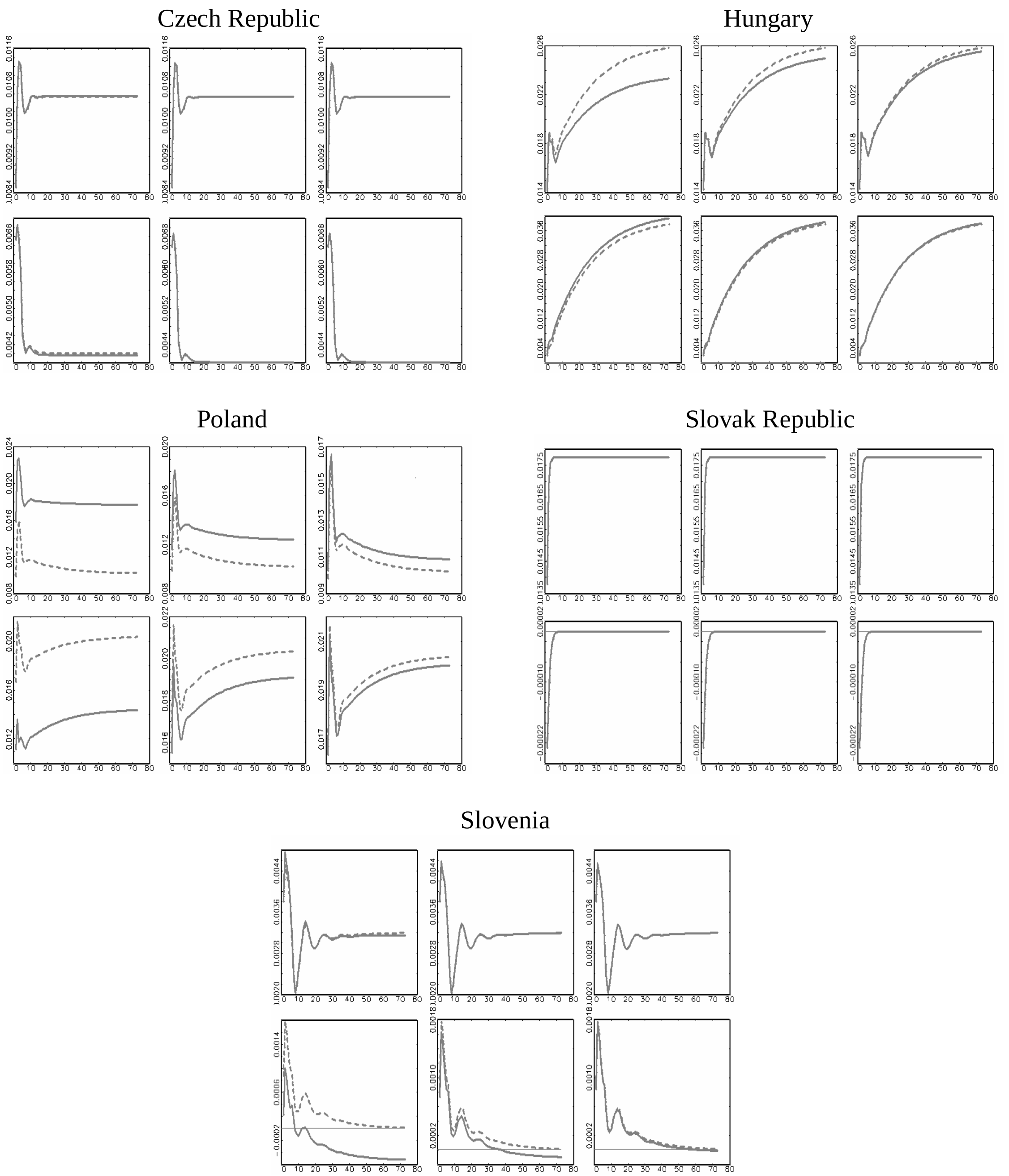
Source: Author's estimates. The highlighted p-values indicate the significance of the structural break at the 5 per cent level.

Table A.3.3 Hansen Structural Break Test

TEST / EXCHANGE RATE	Lags	Breakpoint	Andrews			Bootstarp			Hetero-Corrected		
			SupF	ExpF	AveF	SupF	ExpF	AveF	SupF	ExpF	AveF
NER_CZE	6	29	**0.00**	**0.00**	0.26	**0.00**	**0.01**	0.29	**0.03**	**0.03**	0.22
RER_CZE	6	29	**0.00**	**0.00**	0.29	**0.00**	**0.00**	0.32	**0.04**	**0.04**	0.31
NER_HUN	6	151	**0.00**	**0.00**	**0.00**	**0.01**	**0.01**	**0.00**	0.25	0.23	**0.00**
RER_HUN	6	113	**0.00**	**0.00**	**0.00**	**0.00**	**0.01**	**0.01**	0.12	0.10	**0.00**
NER_POL	3	101	0.13	0.10	0.07	0.16	0.14	0.06	0.22	0.17	**0.03**
RER_POL	3	101	0.20	0.17	0.12	0.25	0.22	0.11	0.28	0.23	0.06
NER_SVK	1	25	0.07	0.17	0.19	0.06	0.18	0.18	0.26	0.38	0.28
RER_SVK	1	25	0.47	0.85	0.94	0.48	0.89	0.94	0.55	0.86	0.92
NER_SVN	6	22	**0.00**	**0.00**	**0.02**	**0.00**	**0.01**	**0.02**	0.26	0.27	0.10
RER_SVN	6	66	**0.01**	**0.01**	**0.01**	**0.02**	**0.03**	**0.01**	0.10	0.11	**0.01**
NER_FRA	1	18	0.72	0.56	0.61	0.69	0.60	0.66	0.76	0.70	0.72
RER_FRA	1	17	0.91	0.72	0.75	0.87	0.75	0.77	0.85	0.75	0.75
NER_DEU	1	25	0.23	0.50	0.74	0.19	0.50	0.75	0.14	0.41	0.65
RER_DEU	1	25	0.13	0.31	0.63	0.13	0.34	0.67	0.10	0.30	0.56
NER_ESP	1	10	0.24	0.22	0.43	0.23	0.26	0.46	0.44	0.45	0.47
RER_ESP	1	10	0.29	0.31	0.57	0.24	0.36	0.61	0.47	0.51	0.58
NER_GRC	3	60	0.96	0.98	0.98	0.93	0.97	0.96	0.64	0.66	0.59
RER_GRC	3	15	1.00	1.00	1.00	1.00	1.00	1.00	0.89	0.73	0.61
NER_ITA	3	14	0.08	0.07	0.61	0.13	0.14	0.68	0.15	0.14	0.23
RER_ITA	3	14	0.09	0.06	0.60	0.15	0.14	0.69	0.12	0.10	0.24
NER_PRT	1	8	**0.00**	**0.00**	0.16	**0.00**	**0.00**	0.19	**0.03**	**0.04**	0.22
RER_PRT	1	8	**0.00**	**0.00**	0.31	**0.00**	**0.01**	0.31	**0.03**	**0.03**	0.35

Source: Author’s estimates. The highlighted p-values indicate the significance of the structural break at the 5 per cent level.

Figure 5: Robustness Checks (NMSs)



Note: For each country, the first three panels in a first row present **nominal** exchange rates responses **to real shocks** with restrictions imposed on 1-year, 3-year and 5-year horizon, respectively. The second row is presented accordingly, but indicates **nominal** rates' responses **to nominal shocks**. Dashed lines are the lines with the identifying restrictions imposed at infinite horizons; solid lines, in respective panels, represent restrictions imposed on 1-year, 3-year and 5-year horizons.

Source: Author's estimations based on IMF IFS and EUROSTAT data.

Annex 4: Variance Decomposition and Impulse Responses (OMSs)

Table A.4.1 Variance Decomposition (OMSs)

Variable	RER				NER			
Variance Decomposition	RER shock		NER shock		RER shock		NER shock	
GERMANY								
1-month	99.8	94.1-100	0.2	0.0-5.9	96.6	86.0-100.0	3.4	0.0-14.0
3-month	100	98.0-100	0.0	0.0-2.0	97.5	92.6- 99.9	2.5	0.1-7.4
12-month	100	99.5-100	0.0	0.0-0.5	97.8	94.6-99.7	2.2	0.3-5.4
24-month	100	99.8-100	0.0	0.0-0.2	97.9	95.0- 99.7	2.1	0.3-5.0
60-month	100	99.9-100	0.0	0.0-0.1	97.9	95.1- 99.7	2.1	0.3-4.9
SPAIN								
1-month	87.2	14.0-100	12.8	0.0-86.0	67.4	0.0-94.6	32.6	5.3-100
3-month	89.9	21.8-100	10.1	0.0-78.2	73.6	9.1-99.6	26.4	0.4-90.9
12-month	97.2	74.0-100	2.8	0.0-26.0	84.8	34.2-99.1	15.2	0.9-65.8
24-month	98.7	85.5-100	1.3	0.0-14.5	87.9	44.5-99.9	12.1	0.1-55.5
60-month	99.5	89.5-100	0.5	0.0-10.5	89.8	45.3-100.0	10.2	0.0-54.7
FRANCE								
1-month	99.9	96.1-100	0.1	0.0-3.9	93.2	83.2-100	6.8	0.0-16.8
3-month	100	97.8-100	0.0	0.0-2.2	93.9	87.6-98.3	6.1	1.7-12.4
12-month	100	99.4-100	0.0	0.0-0.6	94.1	89.1-97.6	5.9	2.4-10.9
24-month	100	99.7-100	0.0	0.0-0.3	94.1	89.4-97.7	5.9	2.3-10.6
60-month	100	99.9-100	0.0	0.0-0.1	94.1	89.7-98.1	5.9	1.9-10.3
GREECE								
1-month	96.2	72.6-100	3.8	0.0-27.4	76.7	31.2-100	23.3	0.0-68.8
3-month	96.0	76.3-100	4.0	0.0-23.7	74.5	30.9-99.9	25.5	0.1-69.1
12-month	98.6	92.3-100	1.4	0.0-7.7	70.2	29.9-97.1	29.8	2.9-70.1
24-month	99.3	96.1-100	0.7	0.0-3.9	68.9	28.4-96.7	31.1	3.3-71.6
60-month	99.7	98.4-100	0.3	0.0-1.6	68.0	27.9-96.8	32.0	3.2-72.1
ITALY								
1-month	98.2	57.8-100	1.8	0.0-42.2	90.9	36.9-100	9.1	0.0-63.1
3-month	99.5	71.9-100	0.5	0.0-28.1	96.5	61.3-100	3.5	0.0-38.7
12-month	99.9	89.4-100	0.1	0.0-10.6	98.0	82.1-99.8	2.0	0.2-17.9
24-month	99.9	93.5-100	0.1	0.0-6.5	98.2	86.0-99.7	1.8	0.3-14.0
60-month	100	96.9-100	0.0	0.0-3.1	98.4	87.8-99.9	1.6	0.1-12.2
PORTUGAL								
1-month	99.3	76.0-100	0.7	0.0-24.0	74.9	32.2-100	25.1	0.0-67.8
3-month	99.8	83.9-100	0.2	0.0-16.1	71.4	32.3-99.1	28.6	0.9-67.7
12-month	99.9	95.2-100	0.1	0.0-4.8	69.1	30.7-95.0	30.9	5.0-69.3
24-month	100	97.5-100	0.0	0.0-2.5	68.7	27.1-94.4	31.3	5.6-72.9
60-month	100	99.0-100	0.0	0.0-1.0	68.4	26.6-95.9	31.6	4.1-73.5

Note: Column I and III reflect contributions of the real and nominal shocks, respectively, to the forecast variance error of the real exchange rate; columns V and VII contain the contributions of the same shocks to movements of the nominal exchange rate; the numbers in columns II, IV, VI and VIII represent the bootstrapped confidence intervals calculated for a particular percentage of variance decomposition.

Source: Author's estimations based on IMF IFS and EUROSTAT data.