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Explaining Exchange Rate Movements in New Member States of the European Union: Nominal and Real Convergence[*]

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Abstract

This paper uses the univariate and bivariate structural VAR variance framework to quantify real and nominal exchange rate volatility in the selective New Member States of the European Union, and identify factors responsible for movements of those rates. The scale and the nature of nominal and real exchange rate volatility are tightly linked to fulfilment of Maastricht criteria, real convergence, and the effectiveness of the nominal exchange rate in absorbing asymmetric real shocks. Given that there is no consensus on the appropriate definition of real convergence, **a**nd since the degree of real exchange rate volatility reflects the scale of idiosyncratic shocks, as well as overall flexibility of the economy to adjust to these shocks, this paper measures the degree of real convergence by the degree of real exchange rate variability. The results indicate that (i) real asymmetric shocks are not insignificant when compared with the poorer Old Member States of the European Union, (ii) the nominal exchange rates, in general, do play a stabilising role, 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, only Estonia and Slovenia are ready to give up monetary and exchange rate independence.

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

TABLE OF CONTENTS

Part I: Introduction

In recent years there has been tremendous growth in interest, and in the literature addressing, Central and Eastern Europe's economic development. Part of the reason for this, of course, is the fact that until the early 1990s very little empirical work had been undertaken on these countries. At the same time, however, economic researchers have been particularly interested in the economic transformation that has taken place in this region since the 1990s, its deepening economic integration, and perhaps most significant, its negotiation of entry, and finally its accession into the European Union. With the entry of the New Member States (NMSs)[3] into the EU in May 2004, attention is now shifting to the issue of their accession into the European Monetary Union (EMU) – in particular to their participation in the third stage of monetary union,[4] which is intended to complete the project of creating a single European market.

The main objective of this paper is to quantify real and nominal exchange rate volatility in the selective NMSs and identify factors responsible for movements of those rates. Nominal and real exchange rate volatility are each important for a number of reasons relating to the European Monetary Union accession:

First, real and nominal exchange rate movements are tightly linked to the issue of Maastricht criteria, without fulfilment of which candidate countries cannot join the eurozone. Second, given that movements of real exchange rates should reflect underlying economic conditions, analysis of these movements and their nature, should allow for the assessment of the scale of existing asymmetries between a particular NMS and the eurozone, and thus serves as a measure of real convergence. Third, by identifying factors that drive nominal and real exchange rate developments, one is able to shed some light on the effectiveness of the nominal exchange rate as a real 'shock-absorber' and thus its facilitating role in the catching-up process. Despite the fact that the NMSs cannot choose to stay outside the European currency zone forever, they can still decide on the *time* they want to participate in the Exchange Rate Mechanism II (ERMII) and on the European Monetary Union (EMU) entry. Thus, the evaluation of nominal and real exchange

[3] The NMSs include: the Czech Republic, Cyprus, Estonia, Hungary, Latvia, Lithuania, Malta, Poland, the Slovak Republic and Slovenia. For the purpose of this paper the NMSs include all the aforementioned counties excluding Cyprus and Malta. The economic structures of Cyprus and Malta have little in common with those of Central and Eastern European Countries (CEECs), and are not considered in the study. Therefore, throughout the paper the term NMSs refers to the Central and Eastern European New Members of the EU.

[4] Prior to the EU membership, in accordance with the Article 109 of the Treaty, all participating countries were obligated the full compliance with the set of requirements related to the second stage of the monetary integration as well as the independence of their national central banks.

rates behaviour constitutes an integral part of the ERMII/EMU monitoring and decision-making process. These issues are explored below.

Nominal exchange rate stability is an *explicit Maastricht criterion* and assumes no exchange rate realignments for at least two years prior the membership.[5] More specifically, preceding EMU entry, counties are obligated to join the ERMII, a framework designed to promote exchange rate stability among the EU Member States outside the euroarea. In the ERMII, currencies are allowed to float within a range of ±15 per cent with respect to an agreed central rate against the euro (bands narrower than ±15 per cent can be set upon the individual request of a country). No 'severe tension', including devaluation in particular, is allowed (i.e., Article 3 of the Protocol on the convergence criteria referred to in Article 121 (1) of the Treaty).[6] 'Severe tension' is generally measured by the degree of deviation of exchange rates from the agreed ERMII central rate against the euro.[7]

The importance of *real exchange rate variability* for the EMU decision making process becomes clear when understood as an indicator of underlying economic conditions. For instance, if it is assumed that prices and wages are rigid, than it can be said that a less volatile real exchange rate between investigated economies (currency zones) implies the following:

- the existence of flexible adjustment mechanisms other than the nominal exchange rate (for example, a high degree of factor mobility (labour and capital), stabilising role of fiscal policy, flexibility and shock-absorbing capacity of a financial sector);
- a low degree of idiosyncratic shocks (for instance, due to similar levels of prices and GDP per capita, synchronisation of business cycles, structural similarity (for instance, unemployment rates), convergence of the interest rate differential, a high degree of trade openness and trade diversification, stable terms of trade, financial market integration);
- more/larger symmetric effects of monetary policies in response to symmetric shocks across economies.

This is because, first, in the onset of an unexpected real shock (nominal shocks do not require exchange rate adjustments), flexibility of the given economy allows smooth tuning of

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[5] The other criteria include: a budget deficit of less than 3 per cent of GDP or which oscillates around 3 per cent for a longer period of time, a stock of public debt of less than 60 per cent of GDP, inflation rate no higher than 1.5 percentage points above the three best performing countries in the existing euro area, long-term interest rates no higher than 2 percentage points above the three best performing countries in the existing euro area.

[6] However, the some official statements (i.e., Pedro Solbes, 2003) suggest that exchange rates would be regarded as stable if they remained within a much narrower band (such as ±2.25 per cent).

[7] Clearly, there are two issues here. One relates to the appropriate choice of the central rate against the euro and therefore the task of estimating equilibrium exchange rates; the other a smooth entry and participation in the ERMII and EMU. This study concentrates on the latter with some insights into the former.

macroeconomic imbalances and thus limits the need for exchange rate adjustments; second, symmetric shocks do not require adjustments in the relative price and as such 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)).[8]

Against this argument it can be said that even when the flexible adjustment mechanisms are not in place, the flexibility of real exchange rates ensures responses to shocks with *asymmetric* effects on the domestic and foreign economies, even if in a currency union. Because the real exchange rate is defined as the relative price of domestic and foreign production, it can still influence the balance of supply and demand between domestic and foreign goods and services. Thus, changes to this price can play a stabilizing role in the economy. So why does one need to worry about large asymmetries between economies which wish to formulate a common currency region if they can always be tackled by changes in the real exchange rate? The answer to this question is not always straightforward (see footnote 8), but it relates to the fact that in the situation when neither a nominal exchange rate nor other adjustment mechanisms are in place the ways the real exchange rate can respond to macroeconomic imbalances are different. *In* the currency zone (such as EMU), in order to restore the equilibrium, real exchange rate changes can only happen via changes in a relative price; *outside* this zone however, the independent monetary and exchange rate policy allows tuning via changes in a flexible nominal exchange rate in addition to changes in relative prices.

The difference between the channels of possible adjustments *in* and *outside* a common currency area possesses a number of challenges. For example, in the presence of excess demand for home production, to restore equilibrium, a real exchange rate appreciation is needed. Under both fixed and flexible exchange rate regimes the expenditure-switching mechanism eventually

[8] The newer findings of the theory (Mundell (1973)), however, point to the fact that if members of a currency zone are financially integrated, a high symmetry of shocks among them, although desirable, is no longer a prerequisite. This is because in a currency area asymmetric shocks can be handled by portfolio diversification and by pooling foreign exchange reserves (for details see Blaszkiewicz-Schwartzman and Wozniak (2006)). Notwithstanding, the risk-sharing argument does not change the fact that in the presence of nominal rigidities (or in the case when insurance markets are incomplete and an effective tax-transfers system does not exist), fewer asymmetric shocks call for a smaller need to adjust, and that giving up exchange rate flexibility represents the cost of monetary unification. Moreover, Ching and Devereux (2000) develop a model in which they show that the benefits associated with risk sharing are likely to be smaller than the costs related to the loss of a flexible exchange rate as an adjustment device. The recent empirical evidence from the current EMU members also indicates modest risk-sharing (De Grauwe and Mongelli, 2005). Similarly, Krugman (1993) contests Kenen (1969) and argues that greater openness of the economy leads to a greater degree of specialisation and thus a higher probability of asymmetric shocks. But this does not have to be true as long as structures of production are similar. Even though the issue remains yet to be resolved, De Grauwe and Mongelli (2005), based on to date developments in the euro area, conclude that it is more likely that the similarity of economic shocks probably increases and does not decrease with greater economic integration.

restores the balance by re-directing demand towards foreign production and supply towards domestic markets (as predicted by Mundell-Fleming-Dornbusch types of models). The adjustment mechanism however depends on the exchange rate regime in place. Under the flexible regime, the required appreciation may happen (fully or partially) via movements in the nominal exchange rate; in the situation when the nominal exchange rate cannot adjust, necessarily, all the pressure is put on the domestic price level which, in this case, must increase, causing domestic inflation. Likewise, if a shock to external demand were negative, in the absence of nominal exchange rate adjustments, without immediate improvements in productivity, real exchange rate depreciation would have to be accommodated by declining wages and prices. This would result in unemployment and political tensions. Additionally, given frictions in prices and wages and in the absence other flexible adjustment mechanisms, positive and negative external demand shocks could result in a prolonged period of real exchange rate misalignments and persistent regional disparities; depending on the shock this situation could have negative implications for countries' terms-of-trade (i.e., a drop in the price of commodities which is not accompanied by the immediate price change results in exporters' loses). Of course there are also other types of shocks, like internal demand shocks, which are probably better tackled with fiscal instruments. Nonetheless, fiscal policy for political reasons is not always easily/promptly available. Besides, the choice of the adjustment mechanism depends on the nature of the shock and internal and external conditions. For instance, for catching-up economies like the NMSs, the right response to macroeconomic imbalances steaming from increases in internal demand might be to accept higher inflation accompanied by current and trade account deficit in anticipation of future surpluses as well as increases in real income and not a fiscal contraction. Also, since the NMSs have a high growth potential, they may run into the problem of overheating after reaching the equilibrium level of economic activity. To slow down wage increases may be needed (see Blanchard (2001) on the Irish and Spanish case).[9]

Another issue is that existing volatility does not have to be an equilibrating reaction to an undesirable shock impinging over economies (which itself does not have to be destabilising either; nevertheless it does point to the existence of asymmetries and the adjustment needs). It can basically be an equilibrium phenomenon, a result of improvements in tradable-sector productivity (i.e., the Balassa-Samuelson-Harrod effect), improvements in equilibrium employment, capital

[9] One can argue (in line with McKinnon, 2000) that if macroeconomic shocks were themselves induced by poor policies, creating a currency area with another country or a group of countries would minimize those shocks (for example by imposing fiscal discipline). Nevertheless, the more similar the structures of economies wishing to formulate a currency union, the lower the likelihood that common shocks will have asymmetric rather than symmetric effects.

accumulation, liberalisation of prices, or increases in the initially undervalued exchange rates, to name a few. These trends have been already observed (see Blaszkiewicz, et. al., 2005) and are likely to continue for quite some time in the NMSs, since these economies are characterised by relatively low income levels. If indeed underlying changes in the domestic price level in these countries simply meant higher steady state inflation and gradually appreciating equilibrium real exchange rate, then the appropriate policy would be again to accept them since they are necessary for achieving higher standards of living.

Whatever the circumstances, managing large real external or internal imbalances in countries with sizable asymmetric real shocks may prove to be difficult, especially during the ERMII period. This is because the fulfilment of Maastricht criteria requires simultaneous stabilisation of all the conditions, including inflation, nominal exchange rates as well as balanced budgets. Meeting these conditions could be further complicated if coupled with large and volatile capital flows and credit booms. Under common monetary policy, on the other hand, unless these asymmetries are absorbed, major trend movements in domestic wages and price levels can be expected. Clearly, were this the case, there is a danger that monetary conditions set by the European Central Bank (ECB) will be too loose for the individual NMSs and thus not optimal. Moreover, given that the real convergence process can be expected to continue for quite some time - were the eurozone to have a fairly constant price level - large asymmetric changes in relative prices could risk price competitiveness in the absence of nominal exchange rate adjustments.[10] *Summa summarum*, if real asymmetries are not insignificant, participating in the ERMII and/or EMU can have unwarranted implications for growth and employment in the NMSs.

In the above examples, it was assumed that nominal exchange rate flexibility can facilitate real exchange rate changes needed to adjust to a particular shock, and that therefore it can be used to stabilise idiosyncratic real shocks. But whether or not exchange rates are good shock stabilisers is another issue and depends to a large extent on the price strategies governing firms' decisions. For example, if nominal prices are set in producers' currencies (i.e., producer-currency-pricing), a flexible exchange rate is more desirable (i.e., they are perfect substitutes for flexible prices). In the situation where nominal prices are set in advance in the currency of the consumers (i.e., local-currency-pricing), nominal currency changes, in the short-run, do not change either real or nominal prices. Here, fixing an exchange rate can be more attractive. However, as Engel (2002)

[10] It is often argued that fixed exchange rates are beneficiary for trade (see Rose 2000, Jeanne and Rose, 2002). But trading companies still remain subject to the real exchange rate changes. Despite the fact that in the currency union they have more time to adjust to price changes (i.e., domestic prices tend to move more slowly than nominal exchange rates), volatile real exchange rates do change price competitiveness for exporting and importing companies.

shows, 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. Also, in the context of the NMSs it is difficult to imagine that the local-currency-pricing prevents nominal exchange rates from acting as a shock stabiliser. Were it nevertheless the case, the observed real exchange rate volatility would have to be induced by market incompleteness and exporters ability to discriminate against different markets. But then relative prices would have to stay constant. Yet, in the NMSs, inflation rates fell dramatically during the 1990s. Therefore, it is probably not much of a shortcut 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.

The fact that real exchange rate volatility is a good indicator of underlying economic conditions makes it a good starting point in analysing a potential trade-off between low and stable inflation and a stable nominal exchange rate in the ERMII. It also sheds the light on the scale of the problems related to the management of real external and internal imbalances in the eurozone, and thus is an obvious candidate for a measure of real convergence. According to the European Commission (EC), the degree of real convergence in a country must be that, without independent monetary and exchange rate policy, asymmetric developments arising from shocks are limited.[11]

Since there is no consensus on the appropriate definition of real convergence[12], **a**nd since the degree of real exchange rate volatility reflects the scale of idiosyncratic shocks, as well as overall flexibility of the economy to adjust to these shocks, this is the definition proposed in this paper. Unlike the indicators analysed in the 2004 EC Report, the real exchange rate volatility criterion does not depend on the exchange rate in place, or 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. It ultimately reflects adjustments to shocks.[13] This argument was already pointed out by Vaubel (1976-1978) in the context of optimal currency areas.[14]

[11] Despite the fact that only the nominal criteria described in the Treaty of Maastricht are binding, real convergence is analysed by the European Commission and the European Central Bank in their Convergence Reports (CR, 2004). Because the non-euro area Member States can only become an integral part of the eurozone upon the positive assessment of the degree of convergence based on those Reports, this issue goes beyond counties own interest.

[12] When discussing real convergence the EC looks at such indicators as the balance of payment position, financial and product market integration. Other research papers, however, look for narrowing gaps of productivity or real income between respective countries and the EMU average, or concentrate on the correlation of unanticipated shocks.

[13] Also, the choice of indicators analyzed by the EC Convergence Report is rather ad hock, and therefore cannot be conclusive. For example, Dabrowski (2005) notices that the Report does not take into account labour market

Without a doubt, measuring the degree of nominal and real exchange rates' volatility is an important part of the analysis of nominal and real convergence. The measure of nominal exchange rate variability (or 'severe tension') adopted in this paper, even if relatively straightforward, necessarily has to be different from what is measured for the purposes of evaluating the Maastricht nominal exchange rate stability criterion. This is because, on one hand, not all of the NMSs participate in the ERMII, and therefore they do not have an agreed central parity from which their nominal exchange rates could deviate; on the other hand, the countries which do participate in the system have not done it for the required time period of two years.[15] Consequently, the deviations of those rates from their 3-year mean are evaluated as a measure of nominal exchange rate volatility vis-à-vis the euro in the NMSs. As far as the analysis of a real exchange rate is concerned, this paper analyses the *unexpected* (i.e., conditional) real exchange rate variances. The latter is done in order to capture the magnitudes of the real exchange rate's variation which cannot be explained by its past movements.[16]

For policy purposes and given the links between the exchange rate volatility, and the conditions upon which the NMSs are going to be admitted to the EMU (i.e., Maastricht), the scale of nominal and real exchange rate variances is necessarily measured on a relative basis. The magnitude of the variance estimated for the NMSs is compared with that estimated for the selected previous EMU candidates (i.e., Germany, France, Italy, Greece, Portugal, and Spain). If conditions in the previous accession counties at the onset of their EMU membership were the same as they are in the new EMU candidates, it would make sense to assume that, at least from the perspective of the scale exchange rate adjustments, it is not less desirable for the NMSs to join the eurozone.

Although relative estimates of nominal and real exchange rates variances in the NMSs are useful in analysing whether ERMII participation will be problematic, and whether joining the

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integration. It is not claimed that the definition adopted in this paper is optimal. But it definitely gives a broader picture on the issue. This is because the behaviour of a real exchange rate should reflect not only the scale of asymmetric real shocks, but also the adjustment ability of factors – other than nominal rates - recognized as important for the creation of monetary unions (for example, factor mobility, the degree of market diversification, fiscal integration, degree of openness, etc).

[14] The detailed analysis of optimal currency literature and its empirical application to the NMSs can be found in Blaszkiewicz-Schwartzman and Wozniak (2005).

[15] Estonia, Lithuania and Slovenia joined ERMII in June 2004; Latvia in May 2005; and the Slovak Republic in November 2005. The reminder of the NMSs do not participate in the ERMII yet.

[16] The univariate analysis on real exchange rate movements employed in this dissertation draws on Vaubel (1976) and is similar to that of von Hagen and Neumann (1994). The results of the univariate nominal and real exchange rates' analysis for the NMSs were published in Blaszkiewicz-Schwartzman and Wozniak's paper on the topic of the fulfilment of optimal currency area criteria by the NMSs (with data up to November 2004). This should not be surprising given the close relationship between broadly understood real convergence and the 'cost' part of the OCA analysis. In this paper the univariate approach is supplemented by the SVAR analysis which further explores and strengthens Blaszkiewicz-Schwartzman and Wozniak's findings on the issue of real exchange rate volatility. The papers should be thus seen as complimentary.

EMU could be more or less attractive than retaining the status quo, they are only partially meaningful. This is because the simple univariate variance approach proposed above ignores the fact that exchange rate fluctuations are caused by a variety of factors, making it difficult to isolate effects of a particular event. It does not allow for separating out changes in exchange rates caused by nominal factors, such as monetary policy or financial market movements from changes caused by real factors, such as improvements in productivity or increased degree of openness. This concern should not be ignored, as the nature of shocks could have different implications for policy recommendations. For example, it is frequently argued that exchange rate flexibility often destabilises rather than stabilises the economy, particularly when volatile asset market flows dominate exchange rate movements. Where this is the case, it could be argued that the NMSs should join the EMU as soon as possible to avoid negative outcomes of such movements on the convergence process (i.e., if, under the current exchange rate regime, shocks to the real exchange rate were mostly nominal, it would make sense to give up monetary and exchange rate independence, as this autonomy is more likely to slow down the convergence process).[17] Of course, the final decision of an individual country to join the EMU is subject to the fulfilment of Maastricht stability criteria, and therefore the nominal exchange rate fluctuations must past the 'severe tension' test irrespective of types of shocks causing them. Nevertheless, given the potential shock–absorbing role of flexible nominal exchange rates, detecting factors behind nominal exchange rate fluctuations still matters in the decision on the *timing* of ERMII and EMU participation. Also, the recent statements from the high rank EU officials suggest that the assessment of the exchange rate stability will take into account factors which could have led to nominal exchange rate appreciation.

The second reason why the proposed univariate exchange rates' approach is not fully meaningful, unless different types of shocks are separated, is that in this analysis the real exchange rate movements are defined so as to measure real convergence. Therefore, an accurate

[17] This having been said, investors are not completely irrational, and they are more likely to invest in an economy characterised by strong fundamentals, and to pull out when fundamentals are week. If this were the case, nominal exchange rate movements would stabilise, rather than destabilise the economy. Moreover, whether exchange rate flexibility plays a significant destabilising role is yet to be empirically resolved. To date, the results are mixed (see Mussa (1986), Baxter and Stockman (1989), Husain, et al., (2004)), but usually conclude that volatile nominal exchange rates do not typically affect other macroeconomic variables other then real exchange rates (i.e., the choice of the regime is neutral). Husain, et al., does, however, find that for developing countries with little exposure to international capital markets, pegs are notable for their durability and relatively low inflation; for developed countries floating systems are preferable; and that in the case of emerging economies the neutrality findings of Baxter and Stockman are nevertheless confirmed. This difference in results shows that the issue of exchange rate regimes is yet to be resolved. The results are nevertheless not in odds with the argument that the choice of the exchange rate regime is not random and depends on the degree of asymmetric real shocks. That is to say that when other adjustment mechanisms are not flexible enough to adjust to real shocks, it becomes preferable to float. Perhaps, given the differences in exchange rate regimes in the NMSs, separation of the shocks within real and nominal exchange rate movements will also shed the light on this issue.

assessment of the progression of this process could be distorted (i.e., overestimated) if the asymmetries originated in financial market sentiment, and temporary policy changes were not singled out from real exchange rate movements. Only when this is done would a high degree of real exchange rate volatility point to a low degree of real convergence, and thus in turn both to the existence of a sizable amount of asymmetric real shocks between a given NMSs and the eurozone, and to a lack of other adjustment mechanisms in addressing them. Because smooth participation in the ERMII and EMU depends on a country's economic situation and its policies, and given the differences in the adjustment channels *in* and *outside* the eurozone, it could be further concluded that because changes to the real exchange rates are asymmetric and permanent (i.e., real), until they become more symmetric, or more flexibility is developed in the labour market or real wages, an early abandoning of monetary and exchange rate instruments may only be possible at substantial macroeconomic cost. More precisely, it could be argued that the existing volatility is not due to volatility in macroeconomic policies or financial markets, but rather due to volatility in asymmetric real forces in the economy.[18]

In the context of the NMSs, the additional argument against premature membership could arise because of the catching-up process. As argued, the progression in the convergence process requires having a higher inflation rate as a right policy response. Therefore, unless there is some flexibility in the interpretation of the Maastricht price stability criterion, this process could be slowed down in the ERMII/ EMU.[19]

So far this discussion has concentrated on why, without the separation of shocks, the univariate variance approach is not well designed to accurately assess the degree of nominal exchange rate volatility and real convergence. It was shown that this is because it is not able to a) recognise factors driving nominal exchange rates, which can be important in the assessment of its stability; and b) to accurately measure the degree of real shock asymmetry and thus real convergence. But another related issue with which the univariate variance approach cannot convincingly deal, unless both the *nature* of shocks and their *dynamics* are recognised, is whether or not the loss of monetary and exchange rate independence (because of its facilitating role in the process of convergence) is a cost of the EMU integration.

Throughout the theoretical analysis presented so far it was assumed that a nominal exchange rate can react rapidly to changes in economic conditions. Therefore, given frictions in prices, it

[18] It is important to recognise that short-run changes to monetary and fiscal policy as well as temporary supply-side shocks (since they cannot affect a real exchange rate permanently) are categorised as nominal. Therefore, if short-run real exchange rate volatility were dominated by asymmetric real shocks, it could only be due to factors like changes in productivity, TOT or preferences.

[19] Also, during the ERMII period, the NMSs are likely to experience intensified capital inflows related to so-called 'convergence-play', making the achievement of Maastricht criteria even more difficult if real asymmetries were large.

represents an important channel of adjustments and thus facilitates convergence. It was further assumed that had a nominal exchange rate not responded at the onset of a real asymmetric shock, this would have to result in a relative price change, which may have an unwanted impact on growth, employment and competitiveness.

But without empirically verifying the shock-absorbing role of a nominal exchange rate (i.e., establishing if it indeed responds to real asymmetric shocks, and moves together and in the same direction as a real exchange rate whenever the necessary changes in relative prices are required) it is not possible to assess to what extent giving up monetary and exchange rate independence actually constitutes a **c**ost of EMU accession, and to what extent nominal flexibility facilitates convergence. The only inference one is able to convincingly draw having confirmed a *high* degree of real exchange rate volatility due to *real* shocks (i.e., after nominal movements were removed) is that it not advisable to join ERMII or EMU (or both) yet.[20] Given the nature of the shock and the catching-up process in the NMSs, in many ways, such a conclusion would be sufficient (i.e., as argued, it may be that the only appropriate way to move these countries to higher income levels is exactly via higher inflation and so these should not rush to give up their own currencies). Nevertheless, this 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 it 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. Of course, if this role were not confirmed, given high real asymmetries, the recommendations about the timing of the ERMII/ EMU accession would not change. Unless greater real convergence is achieved it may be too costly to share a common monetary policy. However, 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).[21]

[20] 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.

[21] 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

Although crude, one way of separating real asymmetric shocks from nominal asymmetric shocks within *real* exchange rate variability would by working with different frequency data (i.e., von Hagen and Neumann, Blaszkiewicz-Schwartzman and Wozniak, Gros and Hobza). This is because it could be argued that that high-frequency real exchange rate changes mostly reflect nominal shocks (i.e., monetary and fiscal policy changes, financial market movements) and low-frequency real exchange rates changes are principally due to real shocks (i.e., productivity, terms-of-trade, tastes). However, since nominal shocks can induce permanent changes to a nominal exchange rate, this methodology cannot be used to differentiate between the split of factors which drive these rates.

Another issue is that using different frequency data in order to separate shocks is potentially compromised. This is because an *a priori* assumption is made about the duration of the impact of nominal shock on the real exchange rate. But nominal shocks can be fairly persistent, effecting real exchange rates for quite some time. Therefore, the technique which seems to be better matched for this purpose is the econometric technique of structural vector autoregression (SVAR) with long-run identifying restrictions due to Blanchard and Quah (1989). This method, because of its multi-equation setup, makes it possible to decompose the time paths of the variation of not only the real but also the nominal exchange rate, into the proportion due to nominal and real disturbances. In addition it allows for testing dynamics between these variables and shocks to them, and thus is suitable for dealing with the issue of whether or not nominal exchange rates play a shock-absorbing role. Also, the Blanchard and Quah (BQ) identification method has the advantage that it does not impose contemporaneous boundaries, but allows the data to determine short-run dynamics motivated by a particular long-run model.

Various research papers including Lastrapes (1992), Clarida and Gali (1994), and Enders and Lee (1997) used this econometric procedure to identify sources of nominal and real exchange rates fluctuations in developed countries. Canzoneri et al., (1996) applies various SVAR models with the BQ decomposition (BQ-SVAR) to estimate potential costs of monetary integration for the selected Old Member States of the EU. All these papers utilise standard assumptions of open macroeconomy sticky-price models in the spirit of a Mundell-Fleming-Dornbusch to classify the shocks in the different VAR systems. For example, Lastrapes as well as Enders and Lee estimate the model with two variables: nominal and real exchange rates which can be affected by two types of disturbances: temporary and permanent. This broad classification of shocks is consistent with the Dornbusch (1976) 'overshooting' model of a small open economy in which nominal

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 EMU average growth rates, and thus inflation.

shocks can have permanent effects on the nominal exchange rate, but only temporary effects on the real exchange rate.[22] Alternatively, Clarida and Gali develop a three-equation open macro model and estimate a structural VAR on the real exchange rate, prices and output to identify three categories of shocks: demand and supply (i.e., IS and AD) and money/financial (i.e., LM).

From the perspective of this study, the results obtained by Canzoneri et al., are particularly interesting. By estimating various SVAR systems on the sample of seven EU Member Sates between 1970 and 1985, the authors come to the conclusion that unless short-run monetary and financial shocks are interpreted as short-run demand shocks, the exchange rate movements in examined countries acted more like an asset price than a 'shock-absorber', and thus larger unions which not only include Germany, Austria and Netherlands, but also France, Spain and the UK are possible (Italy was classified as a border case). Having said that, since the real exchange rate movements in these countries turned to be exogenous to the goods market, the large stabilisation gains from forming the EMU were not confirmed by the data. Therefore, the authors point out that the overall desirability of the EMU – due to its complexity – requires further analysis. As already argued in this paper, and as explored in more detail below, the 'shock-absorbing' role of the flexible exchange rate is only interesting in the context of its facilitating role in the process of real convergence, which can be approximated by the degree of real exchange rate volatility.

There are relatively few studies of this kind in the context of the NMSs. Dibooglu and Kutan (2001) and Borghijs and Kuijs (2004) are familiar exceptions. The aim of the Dibooglu and Kutan study is to 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. Their study concentrates only on two countries, Hungary and Poland, and covers the period between 1990 and 1999. The results suggest that during that time the Polish real exchange rate was mainly driven by nominal shocks whereas, the Hungarian real exchange rate was driven by real shocks. However, the span of their sample includes a period of little exchange rate flexibility and therefore cannot address the issues discussed in this paper. Also, they do not measure the size of nominal and 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 the for the 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 a spirit of

[22] A detailed illustration of the effects of different types of shocks on real and nominal exchange rate can be found in the Enders and Lee's paper.

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 EMU participation. Additionally, the data span is different for different countries (in order to cover flexible exchange rate regimes), but on average covers the years 1995-2003. Since their results reveal that the nominal exchange rate responds little to the shocks that affect output, and that LM shocks have contributed significantly to nominal exchange rate variability, with their contribution being bigger in the smaller and more open countries, they conclude that exchange rate flexibility is not a useful adjustment instrument. Moreover, on average, it appears to be a propagator of LM shocks rather than an absorber of IS shocks.

Given the usefulness of the BQ-SVAR technique in analysing the relationship between a set of variables and the shocks affecting their movements, this econometric method is also utilised in this paper.[23] Although Borghijs and Kuijs' paper, at first glance, seems to deal with similar issues, there are important differences between this study and their work. These differences involve not only a different set of questions, but also differences in analytical tools. In addition to the BQ-SVAR estimation technique - which still differs in the specification of equations entering the model (i.e., this analysis estimates the bivariate BQ-SVAR model similar to Lastrapes, Lee and Enders, and Dibooglu and Kutan as opposed to the trivariate Clarida and Gali type of a system estimated by Borghijs and Kuijs) and robustness checks - in order to measure the relative magnitudes of nominal and real exchange rate variances, the univariate variance approach is also utilised. The differences between the Borghijs and Kuijs analysis and this paper are contrasted in more detail below. As discussed, the primarily focus of this study is to discuss the issues related to the fulfilment of Maastricht nominal exchange rate stability criteria and real convergence. Thus the following questions asked are:

- How volatile are nominal and real exchange rates in the selected NMSs, and how this volatility compares to what was observed in the relatively poor Old Member States before they joined the EMU, as well as in France and Germany? (This is tacked with the univariate variance analysis of nominal and real exchange rates.)
- Has the existing volatility of real exchange rates been changing over time, indicating progress in the process of real convergence? (Again, the univariate variance approach is employed.)

[23] Unfortunately, given the fact that Estonia, Latvia, and Lithuania have adopted fixed exchange rate regimes, they had to be necessarily excluded from the SVAR analysis.

- Given volatility, what is the proportion of temporary (i.e., nominal) as opposed to permanent (i.e., real) shocks in nominal and real exchange rate movements? (The univariate low frequency-data variance analysis on the real exchange rate is performed and compared with the results obtained from the bivariate BQ-SVAR model.)

- Do nominal exchange rates move in the same direction as real rates due to asymmetric real shocks? (Variance decomposition and impulse responses of the estimated BQ-SVAR model are analysed.)

- Given different exchange rate regimes in different countries, is there evidence that less flexible systems have less volatile real exchange rates? (The univariate variance analysis of the real exchange rate brings an answer to this question).

Based on answers to these questions conclusions can be drawn about the actual level of real convergence in the NMSs, as well as the role of the real component in the nominal and real exchange rate movements in these countries. Were both rates moving due to the real shock, it could be reasoned that flexible regimes do absorb macroeconomic imbalances, and thus that these factors should be discounted in the evaluation of the Maastricht nominal exchange rate stability criterion. Likewise, based on the empirical results, it can be evaluated whether the loss of nominal exchange rate flexibility represents a cost of EMU enlargement, or is rather neutral to it. A matter not deeply elaborated is the 'shock-propagating' role of the nominal exchange rate.

Additionally, investigating the dynamics behind the factors which lead to the variance of nominal and real exchange rates allows for a discussion on the following issues:

- The ability of the authorities to affect competitiveness - i.e., if the nominal component in the real exchange rate fluctuations are large and lasting, it could be a reflection of a high degree of nominal price inertia (i.e., 'overshooting'), suggesting that nominal exchange rate changes can influence movements in the real exchange rate, and hence countries' terms of trade (Dornbusch (1976), Kutan and Dibooglu (2001)). Similarly, if real exchange rate fluctuations are dominated by real shocks, competitiveness could only be enhanced by improvements in productivity, permanent changes in a fiscal stance or unemployment.

- The 'exchange rates disconnected' theory: if nominal exchange rate movements are fully passed-through to the general consumer price index (i.e., they are fully offset by the relative price changes keeping real exchange rate constant), the nominal exchange rate flexibility would not be able to bring about the expenditure-switching mechanism.

- The possibility that real and nominal exchange rates movements in the NMSs represent equilibrium adjustments. If the responses of real and nominal exchange rates due to real shocks are of the same magnitude, it would imply that models which explicitly model the equilibrium exchange rate as a function of real economic fundamentals, and in so doing allow for a time-varying equilibrium path of the real exchange rate (i.e., the Stockman equilibrium model (1987), BEER and PEER approaches proposed by Clark and MacDonald (1999)), are more adequate than models which rely on a strict PPP assumption (i.e., the Harrod-Balassa-Samuelson productivity differential model, or disequilibrium sticky price models in the spirit of Dornbusch (1976)).

In contrast, the Borghijs and Kuijs' paper concentrates on the stabilising role of the nominal exchange rate. The authors reason that because the real exchange rate is always able to adjust, even in a currency union, the primary interest of the research which looks at costs of currency unification should be on the shock-absorbing role of the nominal exchange rate. However if it is true that a loss of nominal exchange rate flexibility represents the cost of monetary unification, as already emphasised in this paper, then 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. Therefore, Borghijs and Kuijs' argument is only partially useful for future policy directions in the NMSs. Even if a nominal exchange rate were not addressing macroeconomic imbalances, and its movements were a reflection of money and financial market shocks, one could not say that there is no cost from loosing monetary and exchange rate independence. Moreover their conclusion that the costs of losing exchange rate flexibility are limited, if positive, cannot be really drawn from their results. This is because nominal exchange rates in these five NMSs do move in response to the IS and AS shocks. True, the responses to the AS shocks are moderate, but so are the responses of output to the LM shocks indicating no significant stabilisation gain from EMU membership either (the responses of nominal rates to the IS shocks are on average larger than output responses to these shocks, suggesting a stabilising role of the nominal exchange rate). Also, the finding that LM shocks do move exchange rates, but do not have a significant impact on the real economy *is* equally consistent with it being a shock absorber. Without analysing what would have happened had the exchange rate not moved, it is not possible to definitely say that nominal exchange rate movements are destabilising. The only inference one is able to make based on the Borghijs and Kuijs results it that because output is mostly affected by asymmetric real forces, unless there are other adjustment mechanisms at work, these shocks cannot be mitigated - either *in* or *outside* the eurozone - other than by changes in a relative price. With the magnitude of shocks being unknown, it is however hard to say whether there is a need for greater symmetry. Since Borghijs and Kuijs do not present misspecification tests of their reduced form VAR errors, nor perform robustness checks of estimates obtained through infinite-horizon identifying restrictions of their SVARs, it is difficult to assess the quality of their results. Given the small span of the data in their VAR models, it is surprising that Borghijs and Kuijs present impulse response functions only in terms of standard deviations of shocks and do not provide confidence intervals either for impulse responses or for the variance decomposition. As well documented by the literature (see for example Runkle (1987) in Hamilton (1994) or Kilian (1998a)), normal distributions are not good approximations for impulse responses in small samples. In fact, in this case their distributions are biased and skewed.

In the estimations of the next sections, however, the bootstrap-after-bootstrap algorithm is adopted following Kilian (1998a), which first corrects for the bias in OLS estimates, and than applies bootstrap methods to compute confidence intervals (i.e., a nonparametric residual-based resampling method is applied with the endogenous selection of the VAR lag order).[24] Likewise, since the specification tests performed in this study confirmed the existence of heteroskedasticity, which in some cases could not be eliminated, some of the reduced-form VAR models were estimated using the Weighted Last Squares (WLS) instead of Ordinary Last Squares (OLS), in which case the heteroskedasticity-corrected bootstrap methods were used. This significantly changed obtained results. Finally, Faust and Leeper's (1997) critique of the fact that identification of structural shocks in a VAR by infinite-horizon restrictions, may not be reliable in finite samples is tested.

The broad decomposition of shocks into real and nominal is both the strength and weakness of this paper (since explicit sources of real and nominal shocks can only be identified by adding more variables to the model and imposing additional restrictions). On one hand, it does indicate whether or not nominal exchange rate moved in the same direction as real, in the onset of the real shock, pointing to its stabilising role. On the other hand, it is not able to fully assess its destabilising role in the onset of the nominal shock. This is because, given the methodology, the answer to this question cannot not be conclusive. Even if the ex-post data revealed that variations in a nominal exchange rate were caused by a different type of shock to variations in a real rate, this could not be conclusively interpreted as nominal exchange rate ineffectiveness 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

[24] Kilian (1998a) shows - through extensive Monte Carlo evidence - that confidence intervals for impulse responses generated by this bootstrap-after-bootstrap procedure are much more accurate than standard asymptotic confidence intervals. This bootstrap algorithm seems also to perform well in small and large samples and stationary and non-stationary models (Kilian (1998b)).

stronger, given that nominal shocks represent a whole range of temporary shocks, such as supply, demand or monetary and financial shocks. The only inference one is able to make from such a result is 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). 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., shocks related to the convergence process) this decomposition is sufficient.[25]

The other potential disadvantage of the bivariate SVAR model estimated in this paper is that it is not capable of distinguishing between the types of real and nominal shocks. But, in general SVAR models are not well design to do so. Also, even if real shocks, which dominate real exchange rate volatility, are demand-side shocks, unless they become more symmetric, they may complicate the fulfilment of Maastricht criteria.

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

Parts 2 presents 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. 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.

[25] Since the same arguments apply to the 3-equation VAR system (i.e., the system estimated by Borghijs and Kuijs), the estimation of such a system cannot 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, 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 Choice of the Sample and Its Size

This study focuses on the New Member States of the European Union excluding Cyprus and Malta. The countries considered in the study are: the Czech Republic, Estonia, Hungary, Latvia, Lithuania, Poland, the Slovak Republic and Slovenia. The final choice of the sample and its size poses a number of challenges, related to the following factors:

First, in order that results be comparable as well as compatible with the theoretical model of interest and its SVAR implementation, the countries included in the sample should have relatively flexible exchange rate arrangements, i.e., be free from systematic policy interventions. Unfortunately, the New Member States of the EU adopted a broad range of exchange rate regimes, ranging from the most fixed (i.e., currency boards) to floating. This fact necessarily leads to a discrepancy between countries incorporated in the univariate and bivariate variance analysis since the former does not depend on the exchange rate regime in place. Including countries with fixed exchange rates in the proposed SVAR model would result in misspecification. Moreover, the importance of a nominal shock in real exchange rate movements would be overestimated, as nominal shocks would be transmitted into real rate movements. As a result, for some countries, the separation of shocks into real and nominal components is obtained exclusively from univariate analysis carried out for different frequency data. In these cases, for reasons discussed in Part I, results should be treated with caution as they may be biased. Also, in cases when only the univariate variance analysis is possible, the absorbing role of the nominal exchange rate becomes harder to analyse.

Second, since the choice of countries for the SVAR models must be based on exercised exchange rate regimes, the well-known problem arises that *de facto* exchange rate regimes substantially differ from officially announced regimes. For example, the results of Reinhart and Rogoff's (2002) study show that when the official exchange rate regime is a form of peg, almost half the time the true underlying monetary regime is fundamentally different (the same is true for officially classified floats).

Third, any meaningful structural analysis should involve sufficiently long data span. Unfortunately, for countries under consideration, due to substantial structural changes related to the transformation process, reliable data only exists at the beginning of the 1990s.

Because constrains of the univariate variance analysis relate mostly to the span of the data (i.e., exclusion of the early stages of the transition period), all eight NMSs could be included in the sample.[26] The estimation period spans in 1993M1 to 2005M8. For the SVAR analysis, in order to distinguish between different exchange rate regimes the officially announced exchange rate arrangements (as published by the IMF) were cross-checked with the classification developed by Reinhart and Rogoff.[27] According to both classification schemes countries of interest which exhibit relatively flexible exchange rate regimes are: the Czech Republic, Hungary, Poland, the Slovak Republic, and Slovenia. This excludes Estonia, Latvia, and Lithuania.[28] **Box 1** reviews the evolution of nominal exchange rate regimes in these five NMSs and establishes the final data span for the SVAR modelling.

For comparative analysis, the following previous EMU accession countries were chosen: Italy, Greece, Portugal and Spain (the so-called Club Med countries) as well as France and Germany. This selection was based on the fact that the Club Med countries belong to the periphery of the EU (i.e., they are relatively poor Members of the EU), whereas France and Germany belong to the core. The span of data for chosen countries runs from January 1993 to December 1998; and is justified by two factors. First, 1993 marks the end of the European Monetary System which allowed nominal exchange rates to fluctuate within a band of +/-15 percent (since August 1993). This ensures minimum policy coordination between countries and is important for comparative purposes. Secondly, the fact that the study looks at past data exchange rate variations and compares those with the exchange rate volatility of the New Member States allows the question of endogenouity to be addressed. For example, the study examines the time period where conditions for the Club Med countries as well as France and Germany were not influenced by structural changes induced by the creation of the monetary union itself. Unfortunately, according to the Reinhart and Rogoff's classification, the only country within this group with a *de facto* floating exchange rate regime was Germany. Nevertheless, because between 1993 and 1998 the Club Med countries as well as France adopted some kind of peg or

[26] One could argue that exchange rate regime also matters for univariate variance analysis since, for example, currency boards with an anchor different than the euro naturally exhibit higher euro exchange rate fluctuations. Given that some NMSs have adopted currency boards anchored to the USD, SDR or DM, this would be an important drawback of this approach. However, all NMSs except for Latvia either already have euro-based currency boards or euro-dominated reference baskets (see for example Egert and Kierzenkowski, 2003). Moreover, in order to be admitted to the EMU, Latvia, as all EMU candidate countries, must limit fluctuations of its national currency against the euro. Hence the interest lies in fluctuations of nominal exchange rate against the euro and not the USD or any other currency (see Blaszkiewicz-Schwartzman and Wozniak).

[27] Because Reinhart and Rogoff's study goes back only to December 2001, exchange rate regimes between 2001M12 and 2005M8 were classified in accordance with the IMF code.

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

crawling band regime against the DM at the same time fluctuating freely around the ECU, they were all included in the SVAR modelling.[29]

[29] As primarily interest of this study lies in the NMSs, the Club Med countries as well as France and Germany would not always be discussed to the extend that it serves comparative analysis or is necessary for models' specification.

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

In the *Czech Republic* before 1996 exchange rate flexibility was limited. Initially the official exchange rate was tied to a currency basket and then to the ECU. *De facto*, however, it was a crawling band system around the DM (with the band width of ±2 per cent). 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 per cent (*de facto* the band width was ±5 per cent). Because, between 1993M1 and 1996M3, the official regime was less flexible than indicated by the *de facto* regime, two different samples were investigated: one spanning from 1993M1 to 2005M8 and another from 1996M3 to 2005M8.

In *Hungary*, until December 1998, the exchange rate regime was a *de facto* crawling band around the DM, with the band width of ±5 per cent until May 1994, and ±2 per cent between May 1994 and January 1999. From January 1999 to December 2001, it 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 per cent to ±15 per cent. Since more official flexibility was only announced in 2001, analysis conducted on so few data points would be questionable. Therefore, and given some flexibility within the *de facto* crawling band regime, the estimation period covers the years 1993M1-2005M8.

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 euro (ECU) with a band width of ±5 per cent; there was a pre-announced crawling band around the DM and the US dollar of ±7 per cent. Between February 1998 and April 2000, the band width was systematically widened (up to ±15 per cent). 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 from 1995M6 to 2005M8.

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 per cent. 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 up to ±7 per cent. As of September 1997, *de facto* the band was widened to ±5 per cent and a pre-announced crawling band of ±7 per cent 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 per cent band of DM/euro. Taking into account policy changes in the exchange rate regime, the estimation period starts in 1996M8 and ends in 2005M8.

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 per cent (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 within ±15 per cent. Unfortunately, the period of greater *de jure* flexibility is not long enough to perform the estimation. Therefore, the estimation based on the data spanning from 1993M4 to 2005M8 was tried (before April 1993 a *de facto* regime was classified as freely falling). The final results, however, are presented for the period 1996M9 to 2005M8.[30]

Source: Compiled by the author based on Reinhart and Rogoff study and the IMF classification.

[30] This is relates to the heteroskedasticity problem and will be discussed in depth in the next section.

2.2 Data Source and Transformation

Monthly data on period average nominal exchange rates against the euro up to August 2005 - for all NMSs – were soured from Eurostat. Eurostat also provided data for the euro area consumer price index (HICP). Former eurozone national currencies vs. euro/ECU, as well as consumer price indexes for Old Member States considered in the sample were taken from the IMF IFS. All series were transformed into logarithms. Price indexes were seasonally adjusted (with the use of the software Demetra), and then scaled with the base period equal to one in 1995. The individual real exchange rate indexes were calculated as nominal NMSs/euro rates deflated by the relevant consumer price indexes (CPI and HICP).[31].

2.3 Graphical Presentation

Figure 2 presents developments of real and nominal exchange rates as well as price ratios between 1993M1 and 2005M8, for the NMSs included in the univariate and/or bivariate variance analysis. There are a few general points to make. First, the observed dynamics of nominal exchange rates confirm the choice of the sample for the SVAR analysis; i.e., nominal exchange rates in Estonia, Latvia and Lithuania (from 2001) are more or less constant. Second, the graphical presentation of the data reveals clear evidence of the unit root in the series (i.e., almost all series exhibit strong trend movements). Possible exceptions are nominal exchange rates in Estonia and Latvia, and potentially a real exchange rate in Slovenia. Of course, given currency boards in Estonia, and Latvia, and the objective of Slovenian authorities to keep the real exchange rate constant, this picture is not surprising. In terms of the direction of real and nominal exchange rate movements, in all cases, real exchange rates have been appreciating during the time under consideration; nominal exchange rates either appreciated or depreciated; the price ratio series exhibit a downward trend, pointing to deflation processes taking place in all NMSs. The third point to make is that real and nominal exchange rates of all countries with relatively flexible nominal exchange rates tend to move together, as indicated by the coinciding turning points. Nonetheless, over time, they diverge (or move in different directions as in the case of Hungary).[32] This observation was also made by Enders and Lee in terms of developed countries (Canada, Germany and Japan). 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 them: one

[31] An increase in the index indicates currency depreciation relatively to the euro.

[32] Even if it is less clear for Slovenia and Hungary (again, given the objective of monetary authorities to keep the real exchange rate constant in Slovenia and limited initial flexibility in Hungary until 2001, this is not surprising), there is still some evidence of the short-run co-movements.

temporary and one permanent in nature. This is consistent with predictions of the broad class of structural open macro models (i.e., Dornbusch 'overshooting' model or Stockman '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.

2.4 Integration Properties

This Section formally tests the unit root hypothesis in series in question. In the context of the univariate variance approach adopted in this study formal testing is necessary because the lack of stationarity of data in levels would imply that nominal and real exchange rate movements cannot be characterised by their average values, it would be inappropriate to use standard measures of volatility such as variance/ standard deviation of the series. Were it the case, the appropriate way to proceed would be to estimate nominal and real exchange rate movements using data in first differences.

Also, the Blanchard and Quah decomposition of the SVAR model requires series to be in a stationary form. 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 ration between the eurozone and country of interest inflation). Only when all the variables are I(1) and no cointegrating relationship exists, it is appropriate to test the VAR in first differences. Therefore, below, the results of various unit root tests for countries selected in Section 2.1 are discussed.

DF-GLS and MZ Tests

As Maddala and Kim (2002) stress, because Dickey-Fuller, augmented DF and Philips-Perron unit root tests lack power against the meaningful alternative, they should not be used anymore. Therefore, 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).[33] These tests are then applied to nominal

[33] Schwert formula chooses the maximum lag between 12 and 13. However, when the maximum lag is determined somewhat arbitrary, i.e., it is set to 8 or 6, DF-GLS and MZ-GLS tests, which previously indicated the existence of unit root for first differences of nominal and real exchange rates in Hungary, Poland and Slovenia, now indicate stationarity

and real exchange rate series in levels with a null hypothesis of unit root. Because all three tests indicate non-stationarity of the data (i.e., they are I(1), see Annex 1, **Table A.1.1** and **Table A.1.2**), the tests on the series in first differences are conducted.[34]

Unexpectedly, again, the unit some hypothesis cannot be rejected for Estonian, Hungarian, Latvian, Lithuanian, Polish and Slovenian exchange rate. In the case of Estonia, Latvia, Lithuania, Poland and Slovenia differenced nominal and real exchange rates are non-stationary based on all the tests run; in Hungary the results are similar except for the results from the DF-GLS test for the real exchange rate, which – as expected - turned out to be stationary. Similar results are obtained for the Club Med group as well as France and Germany. All the series are non-stationary in levels, but stationarity *could not* be confirmed for the data in first differences (except for Portugal, for which the results are mixed depending on the test applied, and Spain, for which all tests indicate a presence of unit root).

As far as results for price ratio series are concerned (see Annex 1,

Table A.1.3), except for Hungary and Slovenia, for the MZ-GLS tests, all data in levels is non-stationary confirming the lack of cointegration between bilateral pairs of nominal and real exchange rates.[35] Given the puzzling outcome of MZ-GLS tests (see the dynamics of the data in Figure 2), the Modified Schwarz (MSIC) criterion for the choice of lags in the unit root equations was tried and unit root tests were re-run. Here the null hypothesis of a unit root could not be rejected.

Unit Root Tests with a Break

Given that DF-GLS and MZ-GLS tests may not be appropriate for the variables with an apparent structural break (see Perron (1989), Christiano (1992), Zivot and Andrews (1992)), and given the number of unexpected results, the unit root tests which allow a parsimonious single structural break are also performed. To test the data in levels, the test proposed by Perron (1997) is used; to test data in first differences, the Perron-Vogelsang's (1992) test that uses the broken-trend stationary model for measures of purchasing power parity is applied. In both tests the structural break date is treated as unknown and chosen so as to minimize the t-statistic on the α

of those rates. The only exception is the integration property of the nominal exchange rate in Hungary, which remains I(1) irrespective of the maximum number of lags included in MZ-GLS tests (the smaller number of the maximum lag did however change the outcome for the DF-GLS test).

[34] All the unit root tests but MZ-GLS, which is available from the Ng-Perron webside, were written by the author in the Gauss programming language and are available upon request.

[35] Since the purpose of testing integration properties of price ratio series is to check possible cointegration of the nominal and real exchange rates entering the SVAR models, the unit root tests were only performed for the data in levels and excluding Estonia, Latvia and Lithuania (i.e., countries not included in the sample chosen for SVAR modeling).

coefficient; the number of lags is determined by the '*general to specific*' procedure with the maximum number of lags specified as in the previous tests. Perron's strategy is broken into two steps. First, the series are "detrended" by running the following regression by the OLS:

$$y_t = \mu + \beta t + 1(t \succ T_b)(t - T_b) + \tilde{y}_t \qquad (1)$$

Second, the residuals $\tilde{y}$ are tested for the presence of a unit root using the augmented Dickey-Fuller test:

$$\tilde{y}_t = \alpha \tilde{y}_{t-1} + \sum_{i=1}^{k} c_i \Delta \tilde{y}_{t-1} + e_t \qquad (2)$$

The estimation of the Perron-Vogelsang's model entails testing (again by OLS) an equation of the form:

$$y_t = \mu + \delta DU_t + \theta D(TB)_t + \alpha y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-1} + e_t \qquad (3)$$

In both cases, the null hypothesis of a unit root is tested (i.e., it is tested whether or not α is equal to one).[36] The detailed results of those tests are gathered in Annex 1, Table A.1.4-5)

It is clear that for exchange rate series in levels, the previously obtained results are confirmed. Nevertheless, the break coefficient is not always statistically significant, ruling out this type of specification. For series in first differences, stationarity could not be confirmed for Hungary (nominal exchange rate) and France (real exchange rate), with Spain being a border case. Again, the break coefficient is not statistically significant in Latvia, Poland, France, Germany and Greece (for both rates tested).[37] Interestingly, including a smaller number of maximum lags in the case of Hungary and Latvia confirms the existence of break and rules out the unit root hypothesis in these countries.[38]

In terms of price ratio series in levels, the strong presence of a break was identified in Hungary and Slovenia (see Annex 1, **Table A.1.6**). The unit root hypothesis and thus the luck of cointegration between pairs of nominal and real exchange rates could not be rejected this time.

[36] Perron (1997), in contrast to Andrews and Zivot (1992) derive the limiting distribution of the sequential test without trimming.

[37] Note that the statistical significance of dummy coefficients could not be confirmed for Slovakia and the Czech Republic (in the later case only for the real exchange rate) confirming the robustness of the stationarity results obtained on the basis of the DF-GLS and MZ-GLS tests.

[38] Since the Schwart criterion sets large maximum number of lags, it is likely that some models were overparametrisised.

Kwiatkowski, Phillips, Schmidt and Shin (KPPS) Unit Root Test

Since Kim and Maddala suggest confirmatory analysis by using test which set a stationarity hypothesis as a null, given the ambiguous results obtained by the previous tests, KPSS test for variables in question were performed (i.e., first differences of nominal and real exchange rate series in Hungary, Latvia, Poland and Slovenia, all the selected OMSs, and levels of price ratios for Hungary and Slovenia are tested).[39] This time all the differenced exchange rates' series - except for the nominal exchange rate in Hungary and real exchange rate in Latvia - were I(0). Going back to price ratio series, the stationarity hypothesis could not be accepted at the 1, 5 and 10 per cent level either for Hungary or for Slovenia (see Annex 1, Table A.1.1-3)

Concluding Remarks

In summary, based on the results from the unit root tests, all the series of interest enter univariate and bivariate estimations in first differences. Additionally, since all the exchange rate series of countries previously proposed for the structural VAR analysis are stationary in first differences (with some uncertainty assign to the Hungarian nominal exchange rate), and since 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, they are all included in the SVAR modelling.

Even if unit root tests for the differenced nominal exchange rate in Hungary seem not to be robust, it is argued that because the BQ decomposition restricts nominal exchange rate shocks to only have temporary effect on the real exchange rate (with the possible long run effect on the nominal rate), the possible non-stationarity of the Hungarian nominal exchange rate should have a limited impact on the performed analysis. Given the lack of robustness of the unit root tests performed on the Latvian differenced real exchange rate series, the results of the univariate variance analysis for Latvia should be treated with caution and perhaps should be modelled rather using Generalized Autoregressive Conditional Heteroscedasticity (GARCH) techniques.

Part 3: Econometric Methodology

The econometric methodology employed in the study is explained below. As established in Part I, the analytical approach used to address the questions raised in this paper consists of two parts: univariate and bivariate variance analysis. These are discussed in turn.

[39] KSPP test was performed only for the countries for which DF-GLS and MZ tests could not confirmed stationarity of the data in first differences **and** for which the break dummy coefficient was not statistically significant.

3.1 Univariate Variance Analysis: Technical Aspects

Nominal Exchange Rate Volatility.

Within the univariate variance analysis nominal exchange rate volatility differential was calculated in terms of the difference (i.e., distance from the Club Med average) between the three-year average nominal exchange rate volatility calculated for the respective NMSs country vis-à-vis the euro, and the average volatility of Club Med countries vis-à-vis the euro over the period 1993 to 1995 period, as well as the 3-year period preceding the introduction of the common currency (i.e., 1996-1998).[40] Given nonstationarity of the data, nominal exchange rate volatility in the NMSs was calculated over the 3-year period (September 2002 to August 2005) and was defined (in line with Gurjarati (2003)) as a mean of the squared deviation of the first difference of the logged nominal exchange rate from its mean:

$$vol_t = (1/n)\Sigma(\Delta ner_t - \Delta \overline{ner}_t)^2 \qquad (4)$$
$$t = 1...35,$$

where Δner - nominal exchange rate changes

The volatility of exchange rates for the Club Med countries over the period 1993-1995 and 1996-1998 was calculated in the same way as the volatility for the NMSs. Given that the New Member States in almost all cases want to join the ERMII (the Czech Republic, Hungary, Poland) or EMU (Estonia, Latvia, Lithuania and the Slovak Republic)[41] in around 3 years, the proposed 3-year period for volatility estimation represents, approximately, the same stage in the accession process as the period 1993-1995 or 1996-1998 for the Club Med countries, respectively.[42]

Real Exchange Rate Volatility.

Empirical analysis of real exchange rate movements involves estimating the *unexpected* (i.e., conditional) real exchange rate variances between the respective New Member States and the European Monetary Union members treated as a group (i.e., real exchange rates were deflated by the ratio of prices between a particular NMS and the euro area HICP inflation). As spelled out in

[40] Given that Greece joined the eurozone only in 2001, the 3-year period of 1998-2000 was chosen instead.
[41] The Slovak Republic represents a boarder case since it joined the ERMII very recently (see footnote 15).
[42] The 2-year window for the ERM participation for past candidates covers the period from March 1996 to February 1998.

Part I, this approach draws on Vaubel and is similar to that of von Hagen and Neumann and Gros and Hobza.[43] Again, given the unit root process 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 estimated by OLS by regressing seasonally adjusted real exchange rate changes on their own lags:

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

where $\Delta rer_{i,t}$ - relative change in a real exchange rate

Residuals $u_{i,t}$ obtained from these regressions represent conditional real exchange rate shocks (see von Hagen and Neumann). Next, the standard deviations of these shocks (i.e., real convergence) are measured:

$$sd = [\mathrm{var}(u_{i,t})]^{1/2} \quad (6)$$

Because the unexpected component in this autoregressive model (i.e., residuals from the estimated model) is itself a generated regressor (i.e., a deviation from the mean), an attempt to instrument the conditional standard deviations was made. However, the performed Hausman specification error test did not support this method of estimation. To obtain white-noise errors $u_{i,t}$ from the estimated models (i.e., eq. (5)), where desirable, the dummy variables were used. Necessarily, this lowered computed standard errors (and hence the proposed measure of real convergence). However, events responsible for lack of normality (i.e., financial crises, random exchange rate movements, contagion effects from other markets) are generally outliers and are unlikely to repeat themselves in the future in any systematic manner. To some extend, therefore, this also corrects for the negative bias in measuring real convergence (i.e., bias due to asymmetries arising from speculative pressures or irresponsible central bankers).

In order to check whether volatility changes in real exchange rates are significant (i.e., test for variance equality between sub-samples), various statistical tests were performed. Von Hagen and Neumann propose White's tests for heteroskedasticity. In this study an ARCH test was additionally carried out, as financial market data often follow an Autoregressive Conditional Heteroscedasticity (ARCH/GARCH) process. Where it was the case, the presented standard errors come from estimating such processes by the Maximum Likelihood method (GARCH (1,1)

[43] Gros and Hobza, however, look at observed rather than unexpected exchange rate variability.

turned to be a sufficient specification). Finally, unlike von Hagen and Neumann (1994), this study does not allow the interactive dummies on the lag terms from the autoregressive models (eq. (5)) in order to take into account structural breaks. This is because it would be inappropriate to pool regressions for which variances are believed to be different (since stability tests based on dummy variables or pooled regressions explicitly assume equal variances). Therefore, in order to obtain conditional variances, separate regressions were estimated for each sub-sample (1993-1995, 1996-1998, 1999-2004/5). To facilitate assessments of the magnitude of these real exchange rate variances (i.e., to decide when the variance should be considered large and when - small), estimates of the observed real exchange rate volatility of the selected current EMU members are also provided; they where obtained by applying the same methodology to the two sub-samples (i.e., 1993-1995 and 1996-1999).

As argued in Part I, 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 the real exchange rate movements. Of course, 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. In order to distinguish between real and nominal shocks, the methodology described above is applied to the different frequency data (i.e., monthly and quarterly). A crucial assumption (in line with von Hagen and Neumann) is made that high-frequency real exchange rate changes mostly reflect nominal shocks, and low-frequency real exchange rate changes are principally due to real shocks. This distinction is also 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

This Section sets out with the technical aspects of the structural VAR model considered in this paper. Given the set of variables of interest $y_t = (rer_t, ner_t)'$, where rer_t and ner_t stand for real and nominal exchange rates respectively, the model can be written in the following form:

$$By_t = \Gamma_0 + \Gamma(L)y_{t-1} + \varepsilon_t \qquad (7)$$

where 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. In keeping with the economic model under consideration, 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. It is further assumed that rer_t and ner_t are characterised by single unit root processes and are not cointegrated. In this case, the equation of interest should be specified in first differences, i.e., $\Delta y_t = (\Delta rer_t, \Delta ner_t)'$. Given that there are more parameters than equations to be estimated, the inference starts from estimating the flowing reduced form VAR model by OLS:[44]

$$\Delta y_t = C_0 + C(L)\Delta y_{t-1} + e_t \tag{8}$$

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:

$$\Omega = B^{-1}D(B^{-1})' \tag{9}$$

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 $\Theta(1)$. These long-run multipliers can be obtained by estimating the following moving average representation of Δy_t:

$$\Delta y_t = \mu + \Phi(L)e_t \tag{10}$$

where $\Phi(L) = (I - C(L)L)^{-1}$, $\quad \mu = (I - C(1))^{-1}C_0$

Since $e_t = B^{-1}\varepsilon_t$ this can be written in terms of structural shocks:

$$\Delta y_t = \mu + \Theta(L)\varepsilon_t \tag{11}$$

where $\Theta(L) = \sum_{k=0}^{\infty}\Theta_k L^k = \Phi(L)B^{-1} = B^{-1} + \Phi_1 B^{-1}L + \ldots$, that is $\Theta_k = \Phi_k B^{-1}$ for k=0,1…

The elements of Θ_k (i.e., θ_{ij}^k) give impulse responses of Δy_t to changes in ε_t.

[44] 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.

Now, since y_t is non stationary, the **long-run** impacts of shocks ε_t on the level of y_t are given by:

$$\lim_{s\to\infty} \frac{\partial y_{t+s}}{\partial \varepsilon_t} = \theta_{1i}(1) = \sum_{s=0}^{\infty} \theta_{1i}^{s} \quad \text{, for i =1,2} \tag{12}$$

Because it was assumed that ε_{2t} has no long-run effect on a real exchange rate, i.e., $\lim_{s\to\infty}\left[\frac{\partial rer_{t+s}}{\partial \varepsilon_{2t}}\right] = \theta_{12}(1) = \sum_{s=0}^{\infty} \theta_{12}^{s} = 0$, $\Theta(1)$ can be obtained as a lower triangular:

$$\Theta(1) = \begin{bmatrix} \theta_{11}(1) & 0 \\ \theta_{21}(1) & \theta_{22}(1) \end{bmatrix} \tag{13}$$

It is easy to see that $\Theta(1)$ equals $(I - C(1))^{-1} B^{-1}$, which can be re-written as:

$$B^{-1} = [(I - C(1))\,\Theta(1)] \tag{14}$$

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

$$\Omega = (I - C(1))\,\Theta(1) D((I - C(1))\,\Theta(1))' \tag{15}$$

which can be further rephrased to:

$$[(I - C(1))^{-1}]\ \Omega\ [(I - C(1))']^{-1} = \Theta(1) D\, \Theta(1)' \tag{16}$$

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 (which *de facto* means that structural shocks ε_t have unit variances) 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*, $\varepsilon_t = (\varepsilon_{1t,} \varepsilon_{2t})'$:

$$\lim_{s\to\infty} \begin{bmatrix} rer_{t+s} \\ ner_{t+s} \end{bmatrix} = \begin{bmatrix} \theta_{11}(1) & 0 \\ \theta_{21}(1) & \theta_{22}(1) \end{bmatrix} \begin{bmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \end{bmatrix} \tag{17}$$

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). The long run derivative is defined as:

$$\lim_{s\to\infty}\left[\delta ner_{t+s} / \delta\varepsilon_{2t}\right] = \theta_{22}(1) = 1 \tag{18}$$

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.

To allow for structural decomposition of the forecast error variance it is sufficient to assume that a nominal shock has no long-run effect of on a real exchange rate, i.e., $\theta_{12}(1) = 0$. With this restriction imposed, it becomes possible to determine the relative importance of the nominal and real shocks to the evolution of nominal and real exchange rates - one of the central objectives of this paper. To construct the forecast error variance decomposition, one needs to determine how much variability in forecasting y_t at time $t+s$ based on information at time t is due to variability in the structural shocks ε_t between times t and $t+s$. It can be shown that since the variance of the forecast error for the real exchange rate may be decomposed as follows:

$$\text{var}(y_{1t+s} - \hat{y}_{1t+s|t}) = \sigma_1^2(s) = \sigma_1^2[(\theta_{11}^0)^2 + \ldots + (\theta_{11}^{s-1})^2] + \sigma_2^2[(\theta_{12}^0)^2 + \ldots + (\theta_{12}^{s-1})^2] \tag{19}$$

the variability of the real exchange rate due to real shocks is equal to:

$$\rho_{1,1}(s) = \frac{\sigma_1^2[(\theta_{11}^0)^2 + \ldots + (\theta_{11}^{s-1})]^2}{\sigma_1^2(s)} \tag{20}$$

and the variability of the real exchange rate due to nominal shocks is equal to:

$$\rho_{1,2}(s) = \frac{\sigma_2^2[(\theta_{12}^0)^2 + \ldots + (\theta_{12}^{s-1})]^2}{\sigma_1^2(s)} \tag{21}$$

Analogously, the same decomposition can be applied to the forecast error for the nominal exchange rate.

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 nominal exchange rate fluctuations (on a monthly basis) and real exchange rate fluctuations (on a monthly and quarterly basis) are estimated and contrasted with magnitudes of such fluctuations in the selected Old Member States.

Monthly Volatility Changes.

Nominal Exchange Rate Variability. As a reminder, the magnitudes in **Table 4.1.1** reflect the differences between the each of the NMSs (September 2002 and August 2005), and the average of the nominal exchange rate volatility calculated for the group of selected OMS over the two different 3-year time periods, i.e., before the participation of these countries in the ERMII (1993-1995) and before the EMU membership (1996-1998).[45]

Table 4.1.1 Nominal exchange rate volatility: distance from the Club Med average

Average	Czech Rep.	Estonia	Hungary	Latvia	Lithuania	Poland	Slovak Rep.	Slovenia
1996-1998	0.24	-0.75	1.83	0.36	-0.75	2.96	0.13	-0.73
1993-1995	-1.07	-2.06	0.52	-0.96	-2.06	1.64	-1.19	-2.05

Source: Author calculation based on IMF IFS and ECB data.

In terms of the distances to the Club Med average between the years 1996-1998, in all cases except for Estonia, Lithuania and Slovenia, the differences in volatility are positive, indicating that NMSs have more volatile nominal exchange rates than the Club Med countries. Not surprisingly, Estonia and Lithuania, both participating in the ERMII mechanism and maintaining euro-based currency board arrangements (CBA), have more stable nominal exchange rates relative to the Club Med average. The stability of the Slovenian currency is clear, despite its *de jure* managed float regime (or *de facto* crawling band around euro) up to June 2004 and ERMII

[45] Since the results presented in this Section are an updated version of what was published in Blaszkiewicz-Schwartzman and Wozniak, they may slightly differ. The differences may steam from the different method of seasonal adjustment adopted in this paper, the availability of the HICP series going back to 1993 (previously missing data were approximated by the eurozone PPI index) as well as euro based exchange rates (previously cross exchange rates had to be calculated) at the ECB webside. Nevertheless, the discrepancies are not substantial and thus confirm the robustness of the previously obtained results.

standard fluctuation margins +/- 15 per cent since then.[46] Among countries which still have volatile exchange rates, the country with the smaller distance from the Club Med average is the Slovak Republic, followed by the Czech Republic and Latvia. Poland and Hungary exhibit the most volatile nominal exchange rates.[47] However, as argued, given that the two-year window for the ERMII participation for past candidates covers the period from March 1996 to February 1998, it can be argued that for comparative analysis, at least for the NMSs which do not participate in ERMII yet, the early 1990s should rather be used as a reference point. Assuming that they will join ERMII as soon as possible, the period 2002-2004/5 represents approximately the same stage in the accession process as 1993-1995 for the Club Med countries.

Comparing the average exchange rate volatility for the each of the NMSs with the average volatility of Club Med countries between 1993 and 1995 instead of between 1996 and 1998 (i.e., the second line in **Table 4.1.1**) brings a strikingly different results: Now, only Poland has a significantly more volatile nominal exchange rate (i.e., has a distance greater than one); the Czech and Slovak Republic exhibit smaller volatility than the Club Med average (i.e., a distance smaller than zero) - the distance for Hungary dropped more than three times.

In conclusion, although the New Member States still have volatile exchange rates when compared with the Club Med average of 1996-1998, this does not appear to be the case when compared with the 1993-1995 average. Clearly, there are three countries which outscore and two countries which underscore the rest of the sample as well as the Club Med average (irrespective of the comparable point in time). These are, respectively, Estonia, Lithuania and Slovenia; and Hungary and Poland. However, the stability of Lithuanian and Estonian nominal exchange rate is not surprising given their currency board arrangements.[48]

Real Exchange Rate Variability. The results of the estimations of the conditional real exchange rate volatility in the NMSs as well as selected members of the EMU (on a monthly and quarterly basis) are collected in **Table 4.1.2** and **Table 4.1.3**

[46] It should be stressed however, that before joining ERMII, the Bank of Slovenia engineered a continuous depreciation of the exchange rate. Since its participation in ERM II, the tolar has been trading close to its central rate of 239.640 tolar per euro (EC Convergence Report, 2004).
[47] The 2003 devaluation of the central parity of the Hungarian forint contributed significantly to its volatility. In 2004 the forint has been again more volatile in the context of increasing inflation, large fiscal deficit and uncertainty related to the future policy course of the Hungarian Central Bank.
[48] The presented differences change in a less favorable direction when the New Member Sates are compared with the OMS group including France and Germany in addition to the Club Med (the results are available upon request). Still, Estonia, Latvia, Lithuania, the Slovak Republic and Slovenia have smaller volatility than the OMS group.

Table 4.1.2 Short-run (monthly data) volatility (NMSs)

VOLATILITY CHANGES		I	II	III	White Heteroskedasticity		ARCH	
		93-95	**96-98**	**99-04/05**	**93-98**	**96-04/04**	**93-98**	**99-04/05**
Czech Rep.	No*/Yes*	0.66	1.90	1.19	**0.02**	**0.01**	0.22	0.13
Estonia	Yes/Yes*	0.75	0.46	0.23	0.80	**0.00**	0.25	**0.03**
Hungary	Yes*/Yes	1.66	1.07	0.98	**0.02**	0.92	0.99	0.91
Latvia	Yes*/No	2.48	0.91	1.44	0.49	0.41	**0.04**	0.69
Lithuania	Yes*/No*	2.75	1.80	1.88	0.99	**0.03**	0.41	**0.00**
Poland	Yes*/No*	1.90	1.78	2.01	0.96	**0.04**	**0.04**	0.28
Slovak Rep.	No/No	0.77	1.05	1.36	0.40	0.83	0.45	0.64
Slovenia	Yes*/Yes*	0.73	0.58	0.39	**0.01**	**0.02**	0.60	**0.09**
Average		***1.46***	***1.19***	***1.18***				

Note: Yes-convergence, No-divergence, i.e., we observe a decrease/increase in standard deviation of real exchange rates between the two tested sub-samples (I-II and II-III); * - Statistically significant changes in standard deviation of real exchange rates between the two sub-samples (based on White Heteroskedasticity and Autoregressive Conditional Heteroscedastic (ARCH) errors tests and 10% significance levels). If the null hypothesis is rejected then errors are heteroskedastic, i.e., the changes in conditional RER variances between sub-samples are statistically significant. Columns from 6 to 9 report p-values of conducted statistical tests.

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

Table 4.1.2 summarises estimates of conditional standard deviations (STDs) of monthly real exchange rate shocks for the NMSs. Among them, there are 3 countries for which standard deviations of real exchange rate shocks exhibit a consistent and decreasing trend. This is the case for Estonia, Hungary and Slovenia. Latvia, Lithuania 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 more volatile. In the Czech Republic there is a clear evidence of stabilizing policies between 1999 and 2004/05. As for the Slovak Republic, there has been a continuous increase in real exchange rate volatility throughout the whole estimating period.

The assessment of the role of different exchange rate regimes for stabilization purposes, as in the case of nominal exchange rate volatility, shows that the fact that the NMSs adopted a broad range of different regimes seems not to matter for real exchange rate stability: The hypothesis that less flexible regimes contribute to more stable real exchange rates was not confirmed by the data (as illustrated by comparisons of Poland and Lithuania, Hungary and the Slovak Republic, Poland and the Czech Republic). This fact can be interpreted as showing that nominal exchange rate flexibility is not necessary to accommodate real exchange rate shocks.

Table 4.1.3 Short-run (monthly data) volatility (OMSs)

VOLATILITY CHANGES		**I**	**II**	**White Hetero**	**ARCH**
		93-95	**96-98**	**93-98**	**93-98**
Germany	Yes*	0.58	0.31	0.31	**0.12**
France	Yes*	0.57	0.32	**mean-eq**	**0.02**
Italy	Yes*	1.27	0.67	0.30	**0.11**
Greece	No*	0.47	0.56	0.61	**0.01**
Portugal	Yes*	0.93	0.47	**0.03**	**0.01**
Spain	Yes*	1.38	0.41	0.64	**0.02**
Average		***0.87***	***0.46***		
Average (ClubMed)		***1.01***	***0.53***		

Note: Yes-convergence, No-divergence, i.e., we observe a decrease/increase in standard deviation of real exchange rates between the two tested sub-samples (I-II and II-III); * - Statistically significant changes in standard deviation of real exchange rates between the two sub-samples (based on White Heteroskedasticity and Autoregressive Conditional Heteroscedastic (ARCH) errors tests and 10% significance levels). If the null hypothesis is rejected then errors are heteroskedastic, i.e., the changes in conditional RER variances between sub-samples are statistically significant. Columns from 5 to 6 report P-values of conducted statistical tests.

Source: Author calculation based on IMF IFS and ECB data

The detailed results for the selected EMU Member States show that all countries except for Greece intensified their effort in lowering real exchange rate volatility at the onset of the euro introduction (a pattern which is not observed for all the NMSs). Moreover, in all countries the reduction in volatility was found to be statistically significant. When the average magnitude of the Club Med real exchange rate shocks in the early 1990s, as well as in years preceding the creation of the eurozone, is compared with the NMSs average, the results show that, on average, the NMSs real exchange rate volatility is over two times higher (2.25) than the real exchange rate volatility of Club Med countries in years preceding the EMU membership (i.e., 1996 to 1998), and 1.2 times higher than the variance of Club Med countries in the early 1990s. It should be however stressed that for countries like Estonia and Slovenia, real exchange rate volatility is smaller or almost the same as it was for the core EMU Members States, irrespective of the time period considered.

Quarterly Volatility Changes.

Table 4.1.4 Long-run (quarterly data normalized to monthly) volatility (NMSs)

VOLATILITY CHANGES		I	II	III	White Heteroskedasticity		ARCH	
		93-95	96-98	99-04/05	93-98	99-04/05	93-98	99-04/05
Czech Rep.	No/No	0.63	1.33	1.49	0.28	0.45	0.23	0.82
Estonia	Yes*/Yes*	1.16	0.52	0.24	**0.00**	**0.12**	0.87	0.70
Hungary	Yes/Yes	1.48	1.23	1.10	mean_eq	mean_eq	0.52	0.39
Latvia	Yes*/No	2.02	1.43	1.55	**0.04**	0.48	0.96	0.32
Lithuania	Yes*/Yes	3.60	2.59	1.61	0.67	0.63	**0.07**	0.48
Poland	No*/No	0.79	1.36	2.73	**0.08**	mean_eq	0.76	0.82
Slovak Rep.	No/No	0.39	1.25	1.61	0.67	0.59	0.46	0.92
Slovenia	No/Yes*	0.44	0.91	0.31	0.32	**0.00**	0.25	0.38
Average		***1.31***	***1.33***	***1.33***				

Note: Yes-convergence, No-divergence, i.e., we observe a decrease/increase in standard deviation of realexchange rates between the two tested sub-samples (I-II and II-III); * - Statistically significant changes instandard deviation of real exchange rates between the two sub-samples (based on White Heteroskedasticity and Autoregressive Conditional Heteroscedastic (ARCH) errors tests and 10% significance levels). If the null hypothesis is rejected then errors are heteroskedastic, i.e., the changes in conditional RER variances between sub-samples are statistically significant. Columns 5 and 6 report p-values of conducted statistical tests.

Source: Author calculation based on IMF IFS and ECB data

Table 4.1.4 presents the estimates of conditional STDs of relative real exchange rate changes obtained for lower frequency (quarterly) data. As postulated, since the real variability of exchange rates is influenced by nominal variability, by working with different frequencies the attempt is made to eliminate the problem of nominal variability in real exchange rate movements. This distinction also serves as a basis for evaluating the differences between asymmetric real and nominal real exchange rate shocks.

Looking at the results for the NMSs and comparing them with the average result for the Club Med, almost the same general conclusion is drawn as for high frequency data. The average stance of NMSs between 1999 and 2004/05 is closer to that of the Club Med between 1993-1995 than between 1996 and 1998. The relative magnitudes of real exchange rate shocks in these two sub-samples were, respectively, 1.2 and 2.5 times higher. Since the results for the early 1990s are the same as for the high frequency data, it could be concluded that neither the degree of nominal nor real shocks is more important for NMSs than it was the case with average shocks for the Club Med countries.

Table 4.1.5 Long-run (quarterly data normalized to monthly) volatility (OMSs)

VOLATILITY CHANGES		I	II	White Hetero	ARCH
		93-95	**96-98**	**93-98**	**93-98**
Germany	Yes	0.51	0.19	mean_eq	**0.66**
France	Yes*	0.50	0.24	**0.03**	**0.11**
Italy	Yes*	1.87	0.30	mean_eq	**0.03**
Greece	No	0.28	1.23	0.59	0.39
Portugal	Yes	1.03	0.37	0.16	0.86
Spain	Yes*	1.42	0.25	**0.05**	**0.11**
Average		***0.94***	***0.43***		
Average (ClubMed)		***1.15***	***0.54***		

Note: Yes-convergence, No-divergence, i.e., we observe a decrease/increase in standard deviation of real exchange rates between the two tested sub-samples (I-II and II-III); * - Statistically significant changes in standard deviation of real exchange rates between the two sub-samples (based on White Heteroskedasticity and Autoregressive Conditional Heteroscedastic (ARCH) errors tests and 10% significance levels). If the null hypothesis is rejected then errors are heteroskedastic, i.e., the changes in conditional RER variances between sub-samples are statistically significant. Columns 5 and 6 report P-values of conducted statistical tests.

Source: Author calculation based on IMF IFS and ECB data

In almost all cases, in the third sub-sample, the magnitude of individual quarterly real exchange rate variances for the NMSs is slightly higher than of monthly changes.[49] 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 Estonia and Slovenia real exchange volatility for the NMSs is higher than it was for the selected OMSs, it is clear that **asymmetric real shocks are still an important source of real exchange rate volatility in these countries**. Also, it is hard to say whether the reported long-run volatility decline in some countries was due to policy changes, or to common shocks hitting those countries. Compared with monthly changes, the decline was significant only for Estonia, and Slovenia between the second and the third sub-sample, and for Estonia, Latvia and Lithuania between the first and the second sub-sample. But even then, the size of the shock in Poland in the third sub-sample is 2.4 times larger than the Club Med average of 1993-95; for the Czech Republic, Latvia, Lithuania and the Slovak Republic, on average, it is 1.4 times bigger. When Latvia, Lithuania and the Slovak Republic are compared with the mid-1990s Club Med average, real volatility becomes almost 3 times larger! Finally, Estonia and Slovenia exhibit much more stable real exchange rates than the Club Med average they are compared with, irrespective of the 3-year period they are compared with.

[49]This result is somewhat in contrast with that of Gros and Hobza who however deal with different span of the data (i.e., 1999-2001) and estimate simple instead of unexpected standard deviations.

Turning to individual cases of selected Member States, as in the case of monthly shocks, it is only Greece that failed to lower the unexpected real exchange rate fluctuations between the two sub-samples leading to the EMU membership.[50] The variance reduction was not statistically significant for Germany or Portugal. The long-run volatility for Club Med countries was slightly higher if compared with the short-run volatility for years 1996 to 1998, but marginally lower between 1993 and 1995 (for the NMSs, in general, the long run volatility was on average slightly higher than short-run volatility).

4.2 Bivariate Variance Analysis.

In what follows, the specification of VAR models is described and tested for adequacy. 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; i.e., the reduced form VARs described by equation (8) are estimated for each country considered in the bivariate analysis. The VAR lag order for each country is chosen based on the Akaike Information Criterion (AIC). The estimated lag length $\hat{p}$ is choose for the value of p that minimises AIC(p) with the maximum number of lags of p_{max}=6. 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 conduced 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.

[50] In the case of Greece, unlike in nominal exchange rate volatility estimations, the magnitude of real shocks was calculated over the years 1993-1998.

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)). The White's heteroskedasticity tests are primarily chooses 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 2, Table A.2.1-2). [51]

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 the case of Czech Republic, Hungary, Slovakia and Slovenia (for the selected OMS, two countries – Greece and Italy – do not pass normality tests). As indicated by the Whie's tests, in the cases of Czech Republic, Slovakia, Slovenia and Italy, 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 to small to confirm asymptotical normality.

Structural Changes.

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

[51] Since the results of Portmanteau tests did not differ from the results of LM tests, only the former are presented in Table 3.2.1. The results of the LM tests can be obtained however from the author upon request.

and "exponential" tests of Andrews and Ploberger (1994) (i.e., SupF, ExpF, AveF tests).[52] 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 breakdate 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.[53]

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 use 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,

---

[52] 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

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

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.[54]

Tests 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 only statistically significant for Hungary; for the rest of the countries, there is no evidence of shifts in the mean growth rates (Annex 2, Table A.2.3). The results obtained from the test due to Hansen are different (Annex 2, Table A.2.4). According to the p-values, there is some evidence of coefficient instability in the Czech Republic, Poland and Slovenia in the NMSs group (in Poland and Slovenia a possible break is only evident in the nominal exchange rates' equations); and Germany and Portugal in the OMS group. However, in Poland and Slovenia the potential structural break is not confirmed once the tests are robust to the presence of heteroskedasticity. The fact that estimated breakdates 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 per cent level for both VAR equations.[55] 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. Moving to the Old Member States, in Germany only CUSUM squared indicated some significance of the structural break; in Portugal, both tests strongly confirmed the presence of the structural break.

As documented by Diebold and Chenn, 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, Germany and Portugal (i.e., the dummy variable, respectively, equals one from 1999M3, 1995M4, 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. No break was assumed for Slovenia.

[54] Given that results of the unit root tests are 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).
[55] The results can be obtained from the author upon request.

Once estimated with structural breaks taken into account, the results of misspecification tests did not change much (Annex 2, Table A.2.3-4). Only in the case of Portugal did the estimated errors turn out to be normal and homoskedastic. In other cases (the Czech Republic, the Slovak Republic and Slovenia), despite various attempts, normal and homoskedastic errors could not be obtained.[56] 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.[57]

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. 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 $bias = E[C - \hat{C}]$ 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;

[56] In cases of 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.
[57] Gauss programming language was used to obtain results presented in this section.

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.[58]

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 construted 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.[59] 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 counties where better specification of VAR models could not be achieved, in order to obtain the White robust variance estimate for the errors, e_t, each of the equations in a reduced form VAR system is estimated by

[58] See Kilian (1998a), p.220 for details.

[59] 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.

WLS instead of OLS (in order to obtain accurate confidence intervals, WLS VAR estimation also replaces OLS in the above described bootstrapping). Because heteroskedasticity is of the unknown form, the following steps are taken:

1) Each of the reduced form VAR equations are estimated by the OLS in order to retrieve the residuals $\hat{e}_t$;
2) Squared residuals in the logarithm form, $\ln \hat{e}_t^2$, are regressed on a constant and the lagged values of nominal and real exchange rates ($\Delta y_t(L) = (\Delta rer_t(L), \Delta ner_t(L))$;
3) Fitted values (i.e., the estimates of the variance) are obtained;
4) Weights are obtained by the inverse of the exponent of the fitted vales;
5) Dependent and independent variables are then multiplied by the square root of those weights;[60]
6) Each VAR equations are then re-estimated by the OLS; weighted residuals are obtained and the mean is subtracted to ensure that the transformed residuals will have zero mean;
7) Finally, based on the weighted residuals the covariance-variance matrix of the VAR model, Ω, is calculated and used for variance decompositions.

In result of the WLS estimation for the Czech Republic and the Slovak Republic, the estimated VAR residuals in where heteroskedasticity previously detected turned to be i.i.d.. Unfortunately, in the case of Slovenia the appropriate weights could not be found. Given that the Hansen structural break test detected a potential break in Slovenia in 1996M9, which could not be confirmed once the heteroskedasticity adjusted bootstrap methods were applied, the sample starting from 1996M9 was tried. 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 1996M9 to 2005M8.

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

[60] To be precise, WLS estimation is performed by first dividing the weighted series by its mean and then by multiplying each observation in the data by the scaled weighted series. The scaling of the weighted series is a normalization that has no effect on parameter results, but makes weighted residuals more comparable to the unweighted residuals (i.e., the variance-covariance matrix computed using weighted residuals is comparable in magnitude to that computed using unweighted residuals).

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 the Czech Republic,[61] Hungary and the Slovak Republic, over 90 per cent of shocks to the real exchange rate are real in nature. Poland and Slovenia are somewhat different, with nominal shocks playing a quite substantial role in the variation of their real exchange rates (in the first month, the extent is 50 per cent for Poland and 20 per cent for Slovenia).

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	95.8	75.5-100	4.2	0.0-24.5	71.9	32.9-100	28.1	0.0-67.1
6-month	98.3	90.0-100	1.7	0.0-10.0	85.2	62.7-98.2	14.8	1.8-37.3
12-month	98.9	93.7-100	1.1	0.0-6.3	88.6	70.8-97.9	11.4	2.1-29.2
24-month	99.4	96.5-100	0.6	0.0-3.5	90.6	75.2-98.1	9.4	1.9-24.8
60-month	99.8	98.5-100	0.2	0.0-1.5	91.9	78.0-98.4	8.1	1.6-22.0
HUNGARY								
1-month	97.8	76.8-100	2.2	0.0-23.2	97.2	72.7-100	2.8	0.0-27.3
6-month	99.5	87.5-100	0.5	0.0-12.5	75.6	46.4-97.2	24.4	2.8-53.6
12-month	99.7	93.1-100	0.3	0.0-6.9	62.4	32.8-92.3	37.6	7.7-67.2
24-month	99.9	96.6-100	0.1	0.0-3.4	53.5	24.4-89.1	46.5	10.9-75.6
60-month	99.9	98.7-100	0.1	0.0-1.3	48.6	18.4-86.4	51.5	13.6-81.6
POLAND								
1-month	50.3	19.4-100.0	49.7	0.0-80.6	36.2	9.8-100.0	63.8	0.0-90.2
6-month	68.5	45.5-100.0	31.5	0.0-54.5	47.1	20.5-99.4	52.9	0.6-79.5
12-month	80.0	62.7-100.0	20.0	0.0-37.3	45.5	23.5-98.2	54.5	1.8-76.5
24-month	89.1	80.3-100.0	10.9	0.0-19.7	42.7	24.2-97.5	57.3	2.5-75.8
60-month	96.0	93.4-100.0	4.0	0.0-6.6	39.1	21.9-97.5	60.9	2.5-78.1
SLOVAK REP.								
1-month	98.3	78.7-100	1.7	0.0-21.3	59.2	21.7-100.0	40.8	0.0-78.3
6-month	98.4	88.3-100	1.6	0.0-11.7	71.5	36.4-95.5	28.5	4.5-63.68
12-month	99.2	93.8-100	0.8	0.0-6.2	77.7	46.7-96.1	22.3	3.9-53.3
24-month	99.6	96.9-100	0.4	0.0-3.1	80.6	51.1-96.4	19.4	3.6-48.9
60-month	99.8	98.8-100	0.2	0.0-1.2	82.4	53.2-97.3	17.6	2.7-46.8
SLOVENIA								
1-month	75.4	36.3-100	24.6	0.0-63.7	6.5	0.0-71.3	93.5	28.7-100.0
6-month	69.6	34.5-100	30.4	0.0-65.5	4.1	0.1-67.3	95.9	32.7-99.9
12-month	75.1	50.0-100	24.9	0.0-50.0	3.0	0.3-57.2	97.0	42.8-99.7
24-month	84.4	70.6-100	15.6	0.0-29.4	3.5	0.2-55.8	95.5	44.2-99.8
60-month	93.5	88.9-100	6.5	0.0-11.1	5.2	0.1-57.6	94.8	42.4-99.9

[61] As indicated in **Box 1**, two sample spans were examined for the Czech Republic. However, this did not change final results.

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

Nevertheless, after a year, the significance of the nominal shock drops significantly, with real shocks explaining around 80 per cent of the forecast variance error of real exchange rates in both countries. 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. On the issue of competitiveness, the fact that there is hardly any presence of nominal shock in real exchange rate movements in all the countries but Poland and Slovenia, suggests that competitiveness in these countries can only be enhanced by improvements in productivity or permanent changes in the fiscal stance. Monetary policy can still be used to impact change relative prices in Poland and Slovenia.

Variance decomposition of nominal exchange rates is more heterogeneous. Nominal shocks overwhelmingly dominate the variation of the Slovenian tolar (over 90 per cent of movements are due to this type of shocks), and are an important part of the volatility of the Polish zloty (60 per cent irrespective of the forecast horizon), the Hungarian forint (despite a very minimal initial impact, after a year it increases to 40 per cent and higher) and to some extent the Slovakian koruna (which remains at the 20 per cent level after a year). In the Czech Republic, the impact is limited and transitory, i.e., dropping to 15 per cent already in the sixth month.

Turning to the results for the OMSs,[62] there is no doubt that real shocks are responsible for real and nominal exchange rate movements in all the countries except for Greece and Portugal in the case of a nominal exchange rate forecast error variance. For Greece, the nominal shock is persistent and amounts to over 40 per cent of the nominal exchange rate volatility in the 1-year horizon; in Portugal it drops to 15 per cent in the same forecast horizon. Interestingly, despite the fact that all countries included in the OMS group except for Germany adopted some form of exchange rate regime *de facto* pegged 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 below). 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

[62] The results for the OMS group, variance decomposition and impulse responses, can be found in the Annex 3.

overall showed that a positive nominal shock leads to currency depreciation (i.e., D(2,2) is always positive, see **Table 4.2.2** below).

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

	L_BAND	D(2,2)	U_BAND
Czech Rep.	0.15	**0.30**	0.46
Hungary	1.03	**2.32**	2.89
Poland	0.52	**1.96**	2.36
Slovak Rep.	0.22	**0.47**	0.63
Slovenia	0.18	**0.99**	1.12

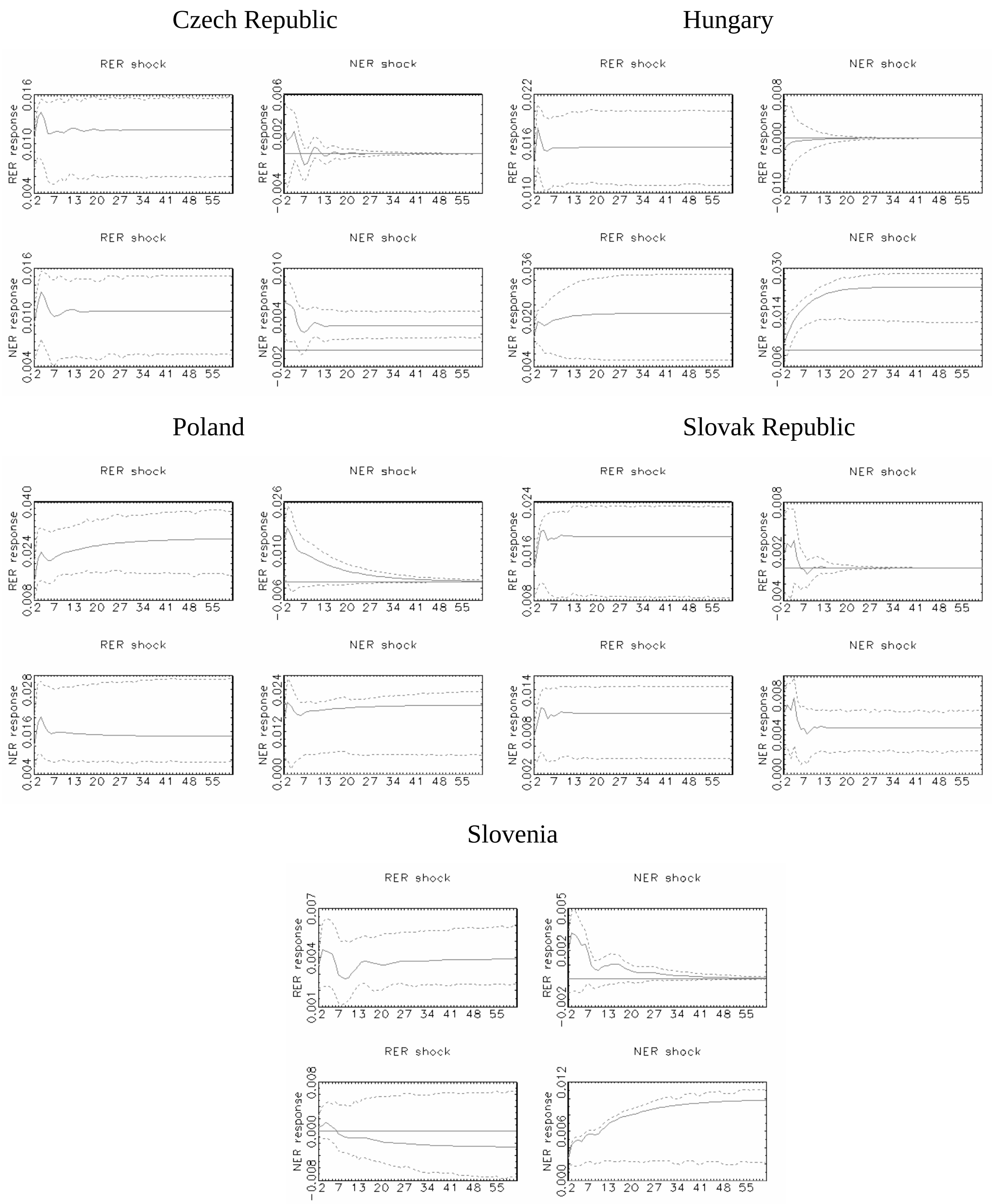
Note: Columns L_Band and U_Band stand for the lower and upper band, respectively, of 90 per cent bootstrapped confidence intervals.

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

This effect is less than proportional in the case of the Czech Republic and the Slovak Republic with Slovenia being a border case.[63] Despite the fact that the overshooting effect is present in Poland and Hungary, the bootstrapped 90 per cent confidence bands are rather wide. Finally, a positive real shock, in all countries, causes long-run nominal **a**nd real exchange rate depreciation. This brings some evidence against the 'exchange rate disconnect' theory. The remainder of this section deals with individual country cases:

[63] Even if D(2,2) equals one, the lower band of the confidence internal goes well below this magnitude.

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 ECB data.

Czech Republic

In the Czech Republic, following a real shock, nominal and real exchange rates jump by almost the same amount (in the long-run, responses to the real shock seem to be proportional). After the initial jump, both rates return to their long-run values within six months. Given that these responses are almost identical, it could be concluded that permanent changes to the real exchange rate (i.e., caused by a real shock) occur mainly through the nominal exchange rate changes and not through changes to a relative price (thus the nominal and real exchange rates should be highly correlated). This would not be possible in the EMU.

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, real exchange rate increases, but this effect is minimal 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. A formal over-identification test (i.e., D(2,2) =1) confirms this: D(2,2) is positive, but significantly less than one according to 90 per cent confidence bands. This is to say that in the long-run, a positive nominal shock causes the currency to depreciate, but the effect is less than proportional. And 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.

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>. The tested coefficient (i.e., D(2,2) =1) is more than two and lies within the 90 per cent confidence bands.

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 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.

In response to a nominal shock a real exchange rate overshoots its long-run value. The impact is long lasting, i.e., it only disappears after 24 months. 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 true for the nominal exchange rate. Despite the evidence of moderate 'overshooting' (i.e., smaller than in Hungary), the response of the Polish zloty to a real shock is greater. Moreover, 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 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. It should be noticed that – in a short-run - the real exchange rate depreciates much more in response to the real shock (i.e., the effect is more than proportional) than the nominal exchange rate does (i.e., the effect is less than proportional). Again, as in the case of 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 jump in response to nominal shocks affecting them, the jump is less than proportional (this is confirmed by the 90 per cent confidence bands).

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. However, 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 domination of the nominal shock in the nominal exchange rate movement is not surprising given the discretionary exchange rate policy in Slovenia (see footnote 46).

Overall, the scale of impulse responses in Slovenia, with the exception of the response of the nominal rate to the nominal shock, is minimal. Taking into account the fact that the magnitude of real asymmetries is at the level estimated for Germany (and below the Club Med average) and the fact that this number is perhaps still overestimated (variance decomposition showed that within the first three months the contribution of the nominal shock to the nominal rate is around 30 per cent), it would seem that Slovenia could join the EMU at very little cost.

Old Member States

In the selected Old Member Sates, similar to the five NMSs, real shocks cause nominal and real exchange rates to depreciate. The long-run values are reached almost immediately. However, these responses are of much smaller magnitude when compared with the five NMSs. Relatively higher jumps in real and nominal exchange rate in response to the real shock in Italy and Spain perhaps reflect greater volatility of real rates during 1993 and 1995 (see Annex 3, Figure 3). There is no evidence of overshooting. The scale of real and nominal exchange rate responses to nominal shocks is minimal.

Robustness Checks.

Since the Faust and Leeper 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 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.[64]

Part 5: Conclusions

This paper estimates the degree and character of nominal and real exchange rate volatility in EMU accession countries and looks at the shock-stabilising role of nominal exchange rate. As discussed in Part I, the degree of real and nominal exchange rate volatility is a crucial element of the decision of whether to enter into the ERMII/EMU. Stability of the nominal exchange rate is directly linked to Maastricht criteria. Stability of the real exchange rate indicates that there is little need for other adjustment mechanisms and/or that other adjustment mechanisms are very effective in accommodating asymmetric real shocks. Furthermore, the degree of existing real asymmetries in real exchange rates can serve as a measure of real convergence between a particular EMU candidate country and the eurozone. Given that the magnitude of real exchange rate variability indicates a country underlying conditions, it is argued that such definition is superior to the more frequently used definition, which is based on differences in per capita income levels. Even if the real exchange rate volatility approach is not new, and goes back to Vaubel in the context of common currency areas, to the author's best knowledge, it has not yet been applied as a measure of real convergence. Apart from the size of real and nominal exchange rate volatility and factors causing it, this paper also addresses the issue of the shock-accommodating role of a nominal exchange rate. Were nominal rates not responding to those shocks that are responsible for the majority of real exchange rate movements, there would be less of a role for monetary and exchange rate independence. Nevertheless, only when the degree of real convergence is known and high, can it be said that there is no cost to the loss of monetary and exchange rate independence.

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 variance decomposition has shown, real exchange rate movements in the Czech Republic, Hungary, and the Slovak Republic can be explained by real shocks at least in 95 per cent already at the 1-month forecast horizon. In the case of Poland and Slovenia, even if nominal shocks still contribute to real exchange rate movements, at the 3-month forecast horizon, nominal

[64] The results of robustness checks for the OMS group can be requested from the author upon request.

shocks are responsible for no more than 30 per cent of real exchange rate volatility. These results are robust to the Faust and Leeper critique.

The domination of real shocks in real exchange rate movements - important for a number of reasons to be discussed - is a key element leading to the positive assessment of the accuracy of the univariate *real* exchange rate variance analysis (at least for countries for which the VAR analysis was possible). Given a moderate impact of nominal shocks to real exchange rate variability at the 3-month forecast horizon, it is not implausible to assume that quarterly volatility changes represent the 'true' magnitude of real convergence in these countries (perhaps in the case of Poland and Slovenia, quarterly real exchange rate volatility should be scaled down by approximately 30 per cent). Likewise, because the variance decomposition for the Club Med countries as well as France and Germany showed that real shocks also dominated real exchange rates in these countries, the benchmark magnitudes for real convergence can be treated with confidence.

Now, based on the univariate variance analysis one can conclude that variances of real and nominal exchange rates in the Czech Republic, Hungary, Poland, and the Slovak Republic resemble the stance of the Club Med countries in the early, rather than, the mid-1990s – the period which is considered to be comparable in terms of their ERMII/EMU preparations. Given that Estonia, Latvia, Lithuania, and Slovenia are planning to join the EMU in 2008 at the latest, the reference Club Med volatility period should be set at the mid-1990s instead. From this group of the early EMU entry planners, clearly Estonia and Slovenia stand out – the level of real convergence achieved by those countries is already higher than the Club Med average (or France and Germany) and is statistically significant between the samples. The results for Slovenia were also confirmed by the SVAR analysis, where the scale of short-run impulse responses was found to be very small.

As the estimates show, real convergence, measured by the size of real asymmetric shocks for investigated Old Member States[65] during the ERMII 'trial' period averages 0.3. Because the exchange rate volatility in these countries during the ERMII period was substantially (and significantly in the relatively high volatility cases) minimised compared to the early 1990s, in order to lessen costs of the EMU membership, the same should be expected from the New Member States (i.e., the average volatility of NMSs, when compared with the volatility of the selected OMSs in the mid-1990s, was 2.5 higher). However, given the striking dominance of real shocks in real exchange rate forecast error variance (at all forecast horizons) in the five NMSs for

[65] This number excludes Greece, as it was not admitted to the EMU together with the rest of countries; once Greece was included, the number increases to 0.5 (see **Table 4.1.4**, Section 14.1).

which the SVAR analysis was possible, and lower levels of income in these countries, it seems unlikely that this process is going to be a quick one. Also, based on the estimates, it seems that the targets set by Latvia, Lithuania, and the Slovak Republic to join the eurozone in 2008, 2007, and 2009 respectively, are probably ambitious, and may not be in the best interest of these countries unless real asymmetries are minimised.

Throughout this paper, it is stressed that the real exchange rate volatility criterion does not depend on the actual exchange rate regime in place, and that the only assumption made is that price stability is desirable, and that therefore, in order to avoid changes in the real exchange rate that entail inflation or deflation relatively to the eurozone, it may be advantageous to maintain monetary and exchange rate independence. Therefore, the issue at hand is whether nominal exchange rates in the NMSs play a shock-stabilising role or not. Unfortunately, as discussed, the univariate variance analysis is not well designed to address this issue. The structural VAR analysis, however, allows to conclude 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 nominal exchange rate at the onset of a real shock, and thus <u>does play</u> a shock-stabilising role and represents a **c**ost of the EMU membership. Even if in the case of Poland, nominal shocks are responsible for 30 per cent of the real exchange rate forecast error variance in the short run, it is also true that almost 50 per cent of the initial responses of the nominal exchange is to due to a real shock. The presence of a nominal shock in real exchange rate movements in Slovenia is perhaps related to the fact that Slovenian authorities used to perform regular interventions in order to keep real exchange rate constant. Moreover, given the limited volatility of the real and nominal exchange rates in the case of Slovenia, the presence of the nominal shocks in the real and nominal exchange rate movements can hardly be an argument against the EMU membership. Rather it points to a low degree of real asymmetries, or the existence of other flexible mechanisms.

Turning to results obtained for the nominal exchange rate movements, it is shown that the distances between the NMSs' volatilities and the Club Med average calculated for the years 1993-1995 are positive only for Poland and Hungary, pointing to their more volatile exchange rates. When compared with the 1996-1998 average, positive volatility distances were also observed in the Czech and Slovak Republic, as well as in Latvia. Stability of Estonian and Lithuanian exchange rates, to a great extent, reflexes their currency board regimes. The results of the B&Q variance decomposition show that the nominal component in nominal exchange rate movements in the five countries covered by SVAR modelling is not insignificant (perhaps to a lesser extent in the Czech Republic). Therefore, it could be argued that the gains from speedy EMU membership involve elimination of these fluctuations. Nevertheless, given the necessity to fulfil Maastricht criteria first – including ERMII participation – this argument is no longer unquestionable. For instance, even if the OMSs faced a similar degree of nominal volatility at the onset of their ERMII participation, this volatility was mainly driven by real factors).[66] Given the specifics and constrains of the ERMII, this piece of evidence, combined with the degree of real asymmetry should not be ignored, as perhaps there is a risk assigned to premature ERMII participation (i.e., due to increased capital flows or consumption boom).[67] Finally, significant nominal exchange rate responses to real shocks in all five NMSs but Slovenia need to be further investigated and perhaps taken into account in assessing nominal exchange rate stability at the end of the ERMII period.[68]

Summing up, (i) given 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, it is not quite clear what the immediate benefits are from giving up monetary and exchange rate independence by all the NMSs, except for Estonia and Slovenia.[69] Additionally, since these economies have not reached their equilibrium levels yet, and given the presence of nominal shocks in the nominal exchange rate movements, to minimise costs of loosing monetary and exchange rate independence, it may be advisable to adopt a more gradual path for ERMII/ EMU entry, at the same time enhancing structural reforms.[70]

As Rogoff (2002) says, even if new open macroeconomic models dominate the international macroeconomic literature in recent years, the Dornbusch's model still offers the quick answer to how monetary policy may impact the exchange rate. In this respect, a few observations on equilibrium exchange rate modelling in the NMSs can be made. First, since in the Czech Republic, Hungary and the Slovak Republic, the permanent component explains more than 90 per cent of the forecast variance error, with virtually no evidence of nominal exchange rate overshooting (Hungary is an exception here), the Dornbusch's disequilibrium model hardly fits the data in these countries. Moreover, since according to the Stockman's (1987) equilibrium

[66] Except for Greece to a certain degree, but one has to bear in mind that Greece was not admitted to the EMU together with the rest of the countries and that its central parity had to be revaluated in the onset of the membership.
[67] One could additionally argue that because nominal shocks do not move real exchange rates in the Czech Republic, the Slovak Republic and Hungary, they are effective tools in stabilising shocks.
[68] Given the broad classification of shocks, at this stage, without bigger identification of shocks, it is impossible to investigate the split between the supply and demand shocks.
[69] The conclusion about Estonia is based only on the univariate variance analysis due to the currency board regime adopted by this country.
[70] Even if one insisted that the differences between the Club Med countries and the NMSs are not all that significant, as shown by Angeloni, Aucremanne and Ciccarelli (2006), it is also not clear whether or not the EMU membership was beneficiary for these countries (i.e., they cannot find and evidence that EMU entry has visible effect on both price setting and inflation persistence).

model, real and nominal rates can simultaneously respond to real shocks, and since the permanent component also dominates nominal exchange rates movements in the Czech Republic, and is not insignificant in Hungary and the Slovak Republic, the equilibrium hypothesis should be further explored.[71] In the case of Poland, a large temporary component in the real exchange rate, and a proportional effect of a nominal shock on a nominal exchange rate, bring some support for the disequilibrium model of the exchange rate. Based on the bivariate SVAR analysis, neither of these models is really suitable for Slovenia.

[71] In the Stockman's model high correlation between real and nominal rates is explained by the simultaneous response of real and nominal exchange rate to a real shock. The concept of equilibrium refers here to the assumption that markets clear through price adjustments, so equilibrium models typically assume that all prices are fully flexible.

REFERENCES

Andrews, D. W. (1993), Tests for Parameter Instability and Structural Change with Unknown Change Points, Econometrica 61 (4), 821–856.

Andrews, D. W., W. Ploberger (1994), Optimal Tests When a Nuisance Parameter is Present Only Under the Alternative, Econometrica 62 (6), 1383–1414.

Bai, J., Lumsdaine, R. L., J.H Stock (1998), Testing For and Dating Common Breaks in Multivariate Time Series, Review of Economic Studies 65, 359-343.

Banerjee, A., Lumsdaine, R.L., J.H. Stock (1992), Recursive and Sequential Tests of the Unit Root and Trend Break Hypothesis. Journal of Business and Economic Statistics, 10, 271-288.

Baxter, M., A. Stockman (1989), Business Cycles and the Exchange-Rate Regime: Some International Evidence, Journal of Monetary Economics, Vol. 23 (May), pp. 377–400.

Bayoumi, T., B. Eichengreen (1993), Shocking Aspects of European Monetary Integration, in F. Torres and F. Giavazzi (eds.), Adjustment and Growth in European Monetary Union, Cambridge University Press, Cambridge.

Bayoumi, T., B. Eichengreen (1996), Operationalising the Theory of Optimal Currency Areas, CEPR DP 1484.

Berkowitz, J., L. Kilian (1996), Recent Developments in Bootstrapping Time Series, Federal Reserve Board, Finance and Economics Discussion Series No. 45.

Berkowitz, J., L. Kilian (1997), Recent Developments in Bootstrapping Time Series, Board of Governors of the Federal Reserve System, Finance and Economics, Discussion Series 94/95.

Blanchard, O. (2001), Country Adjustments Within Euroland: Lessons After Two Years, in Alesina, A, Blanchard, O., Gali, J., Giavazzi, F., and H. Uhlig (eds), Defining a Macroeconomic Framework for the Euro Area: Monitoring the ECB 3, London: CEPR.

Blanchard, O. J., D. Quah (1989), The Dynamic Effects of Aggregate Demand and Supply Disturbances, American Economic Review 79 (4), 655–673.

Blaszkiewicz, M., Kowalski, P., Rawdanowicz, L., and P. Wozniak, Harrod-Balassa-Samuelson Effect in Selected Countries of Central and Eastern Europe, CASE Reports No. 57

Blaszkiewicz-Schwartzman, M., and P. Wozniak (2006), Do the New Member States Fit the Optimum-Currency-Area Criteria in The Eastern Enlargement of the Eurozone, Dabrowski, Marek, Rostowski, Jacek (Eds.), Springer 2005

Borghijs A., L. Kuijs (2004), Exchange rates in the CECs: a blessing or a curse?, International Monetary Fund Working Paper No. 04/02.

Brada, J.C. (1998), Introduction: Exchange Rates, Capital Flows, and Commercial Polices in Transition Economies, Journal of Comparative Economics, 26, 613-620.

Buiter, Willem H, 2000. Optimal Currency Areas: Why Does The Exchange Rate Regime Matter?, CEPR Discussion Papers 2366, CEPR Discussion Papers.

Canzoneri, M., J. Valles and J. Vinals (1996), Do Exchange Rates Move to Address International Macroeconomic Imbalances?, CEPR Discussion Papers no. 1498,

Ching, S., M. B. Devereux (2000), Risk Sharing and the Theory of Optimal Currency Areas: a Re-examination of Mundell 1973, Hong Kong Institute for Monetary Research, HKIMR Working Paper No. 8/2000.

Ching, S., M. B. Devereux (2003), Mundell Revisited: a Simple Approach to the Costs and Benefits of a Single Currency Area, Review of International Economics, 2003 11 (4), pp. 674-691.

Christiano, L.J. (1992), Searching for a Break in GNP, Journal of Business and Economic Statistics, 10, 237-250.

Clarida, R. and J. Gali, (1994), Sources of Real Exchange Rate Fluctuations: How Important are Nominal Shocks?, Carnegie-Rochester Series on Public Policy 41, pp. 1-56.

Clark, P. and R. MacDonald, (1999), Exchange Rates and Economic Fundamentals: A Methodological Comparison of BEERs and FERRs, in MacDonald R., and J. L Stein eds. (1999), Equilibrium Exchange Rates, Kluwer Academic.

Convergence Report 2004, European Central Bank, http://www.ecb.int/pub/pdf/conrep/cr2004en.pdf

Dabrowski, M. (2005), A Strategy for EMU Enlargement, Center For Social and Economic Research, Studies and Analysis No. 290.

Davidson, R. and J. G. MacKinnon (1993), Estimation and Inference in Econometrics, New York: Oxford University Press.

De Grauwe P., F. Mongelli (2005), Endogeneities of Optimal Currency Areas. What Brings Currency Sharing a Single Currency Closer Together? European Central Bank, Working Paper No. 468.

Devereux, M.B.,Engel, C. (2002), Exchange Rate Pass-through, Exchange Rate Volatility and Exchange Rate Disconnect, National Bureau of Economic Research Working Paper No. 8858.

Dibooglu, S. and A. Kutan (2001), Sources of Real Exchange Rate Fluctuations in Transition Economies: The case of Poland and Hungry, Journal of Comparative Economics, 29, 257 – 275.

Diebold, F. X., C. Chen (1996), Testing Structural Stability with Endogenous Breakpoint: a Size Comparison of Analytic and Bootstrap Procedures, Journal of Econometrics 70: 221 241.

Dornbusch, R. (1976), Expectations and Exchange Rate Dynamics, Journal of Political Economy, No. 84, pp. 1161-1176.

Efron, B. (1979), Bootstrap Methods: Another look at the Jackknife, Annals of Statistics, 7: 1 26.
Egert B., R. Kierzenkowski (2003), Asymmetric Fluctuation Bands in ERM and ERM-II: Lessons from the Past and Future Challenges for EU Acceding Countries, William Davidson Institute Working Paper no. 597.
Eichengreen, B. (1991), Is Europe an Optimum Currency Area?, NBER Working Paper 3579.
Elliott, G., T.J. Rothenberg and J.H. Stock (1996), Efficient Tests for an Autoregressive Unit Root, Econometrica, 64, 813 – 836.
Enders, W., B. Lee (1997), Accounting for Real and Nominal Exchange Rate Movements in the Post-Bretton Woods Period, Journal of International Money and Finance 16(2), 223-254.
Engel, C. (2001), The Responsiveness of Consumer Prices to Exchange Rates and the Implications for Exchange Rate Policy: A Survey of a Few Recent New Open-Economy Macro Models, National Bureau of Economic Research Working Paper No. 8752.
Engel, C. (2001), The Responsiveness of Consumer Prices to Exchange Rates and the Implications for Exchange Rate Policy: A Survey of a Few Recent New Open-Economy Macro Models, National Bureau of Economic Research Working Paper No. 8752.
Engel, C. (2002), Expenditure Switching and Exchange Rate Policy, paper prepared for the National Bureau of Economic Research Macro Annual 2002.
European Commission (2004), EMU After 5 Years, European Economy Special report, no.1/2004.
European Commission, Convergence Report 2004, Special Report No. 2, http://europa.eu.int/comm/economy_finance/publications/european_economy/2004/eesp204en.pdf
European Parliament (1998), Adjustment to Asymmetric Shocks, Directorate-General for Research, ECON 104.
Faust, J. and E. Leeper (1997), When Do Long-run Identifying Restrictions Give Reliable Results?, Journal of Business and Economic Statistics 15 (3), 345–353.
Frankel, J. A., A. K. Rose (1998), The Endogeneity of the Optimum Currency Area Criteria, Economic Journal 108, July 1998.
Gros D., A. Hobza (2003), Exchange Rate Variability as an OCA Criterion: Are the Candidates Ripe for the Euro?, International Centre for Economics Growth European Centre, Working Papers 23, Macroeconomic Studies.
Gurjarati D. N. (2003), Basic Econometrics, Fourth Edition, McGraw-Hill.
Hamilton, J.D., (1994), Time Series Analysis, Princeton University Press.
Hansen, B. E. (2000), Testing for Structural Change in Conditional Models. Journal of Econometrics 97 (1), 93–115.
Husain, A.M., Mody A., K.S. Rogoff (2004), Evolution and Performance of Exchange Rate Regimes, National Bureau of Economic Research Working Paper No. 10673.
Jeanne O., A. K. Rose (2002), Noise Trading And Exchange Rate Regimes, The Quarterly Journal of Economics, MIT Press, vol. 117(2), pages 537-569, May.
Kenen, P. (1969), The Theory of Optimum Currency Areas: An Eclectic View, in R. Mundell and A. Swoboda eds., Monetary Problems in the International Economy, Chicago: University of Chicago Press.
Kilian, L. (1998a), Small Sample Confidence Intervals for Impulse Response Functions. The Review of Economics and Statistics, 80, 218– 230.
Kilian, L. (1998b), Accounting for Lag-Order Uncertainty in Autoregressions: the Endogenous Lag Order Bootstrap Algorithm. Journal of Time Series Analysis, 19, 531–538.
Kilian, L., (2001), Impulse Response Analysis in Vector Autoregressions with Unknown Lag Order, Journal of Forecasting, No. 20. 161-179.
Krugman, P. (1993), Lessons from Massachusetts, in Giavazzi, F. and F. Torres (eds.), Adjustment and Growth in the European Union, Cambridge: Cambridge University Press, 241-261.
Kwiatkowski, D., P.C.B. Phillips, P. Schmidt and Y. Shin (1992), Testing the Null Hypothesis of Stationarity against the Alternative of a Unit Root: How Sure are we that Economic Time Series Have a Unit Root ?, Journal of Econometrics 54, 159 – 178.
Lastrapes, W. D. (1992), Sources of Fluctuations in Real and Nominal Exchange Rates, Review of Economics and Statistics, pp. 530-539.
Lütkepohl, H. (1990), Asymptotic Distributions of Impulse Response Functions and Forecast Error Variance Decomposition of Vector Autoregressive Models, The Review of Economics and Statistics, 72: 116 125.
Madala, G.S., L.M Kim (1998), Unit Roots, Co-integration and Structural Change, Cambridge University Press.
McKinnon, J. (2000), Bootstrap Inference in Econometrics, presented as the Presidential Address at the 2002 Annual Meeting of the Canadian Economics Association.
McKinnon, R. (1963), Optimum Currency Areas, American Economic Review, 53, pp. 717-25.
McKinnon, R. (2000), Mundell, the Euro and Optimum Currency Areas, Journal of Policy Modeling 22(3), pp. 311-324.
Mundell, R. (1961), A Theory of Optimum Currency Areas, American Economic Review, 51, pp. 657- 65.
Mundell, R. (1973), Uncommon Arguments for Common Currencies, [in:] H. Johnson and A. Swoboda (eds.), The Economics of Common Currencies, London: George Allen & Unwin Ltd , pp. 143-173
Mussa, M. (1986), Nominal Exchange Rate Regimes and the Behavior of Real Exchange Rates: Evidence and Implications, Carnegie Rochester Conference Series on Public Policy, Vol. 25, 117-213.

Ng, S., P. Perron (1995), Unit Root Tests in ARMA models with Data Dependent Methods for the Selection of the Truncation Lag, Journal of the American Statistical Association, 90, 268-281.

Ng, S., P. Perron (2001), Lag Length Selection and the Construction of Unit Root Test with Good Size and Power, Econometrica, 69, 1519 – 1554.

Obstfeld, M. (2002), Exchange Rates and Adjustment: Perspectives from the New Open Economy Macroeconomics, National Bureau of Economic Research Working Paper No. 9118.

Perron, P. (1997), Further Evidence on Breaking Trend Functions in Macroeconomic Variables, Journal of Econometrics, 80, 355-385.

Perron, P., (1989), The Great Crash, the Oil Price Shock, and the Unit Root Hypothesis, Econometrica, 57, 1361-1401.

Perron, P., T. Vogelsang (1992), Nonstationarity and Level Shifts with an Application to Purchasing Power Parity, Journal of Business and Economic Statistics, 10, 301-320.

Rose, A. (2000), One money, one market: the effect of common currencies on trade, Economic Policy, April.

Runkle, D. E., (1987), Vector Autoregressions and Reality, Journal of Business & Economic Statistics, American Statistical Association, vol. 5(4), pages 437-42, October.

Sims, C. A., R. Zha (1995), Error Bands for Impulse Responses, Federal Reserve Bank of Atlanta Working Paper: 95 96.

Stockman, A.C. (1987), The Equilibrium Approach to Exchange Rates, Federal Reserve Bank of Richmond Economic Review, 73, 12-30.

Vaubel, R. (1976), Real Exchange-Rate Changes in the European Community - The Empirical Evidence and its Implications for European Currency Unification, Weltwirtschaftliches Archiv 112, pp. 429-470.

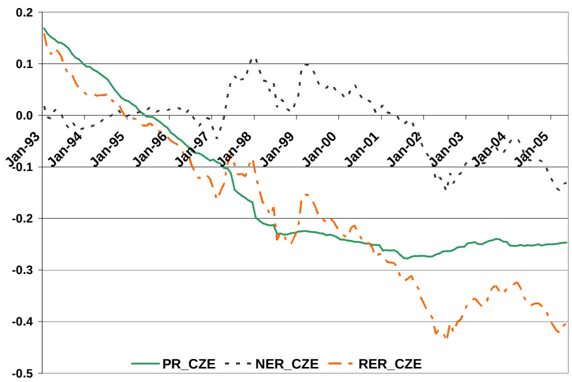
Vaubel, R. (1978), Real Exchange-Rate Changes in the European Community: A New Approach to the Determination of Optimum Currency Areas, Journal of International Economics

Von Hagen, J., M. Neumann (1994), Real Exchange Rates Within and Between Currency Areas: How Far Away is EMU?, The Review of Economic and Statistics, Volume 76, Issue 2 (May, 1994), pp. 236-244.

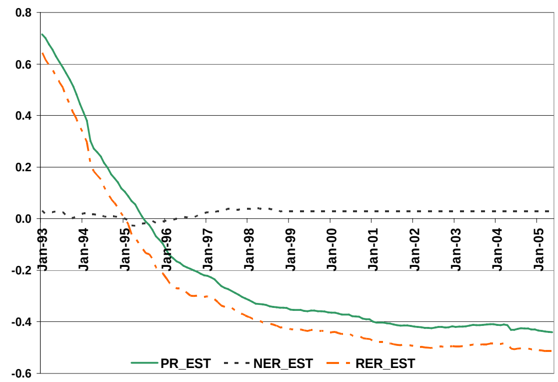
Zivot, E., D. W. K. Andrews (1992), Further Evidence on the Great Crash, the Oil Price Shock and the Unit Root Hypothesis, Journal of Business and Economic Statistics, 10, 251-270.

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

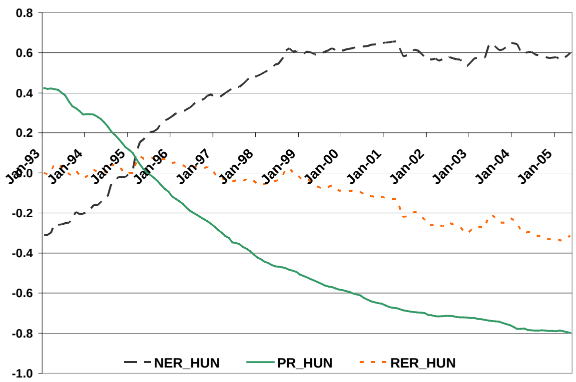
Czech Republic



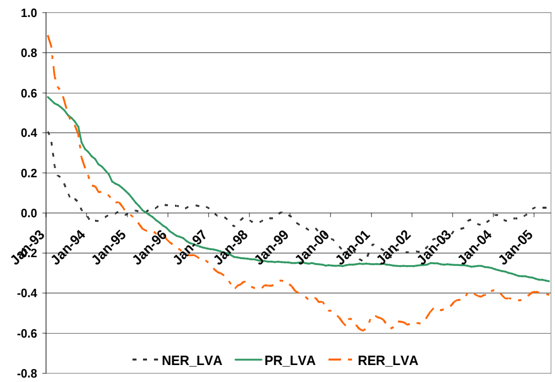
Estonia



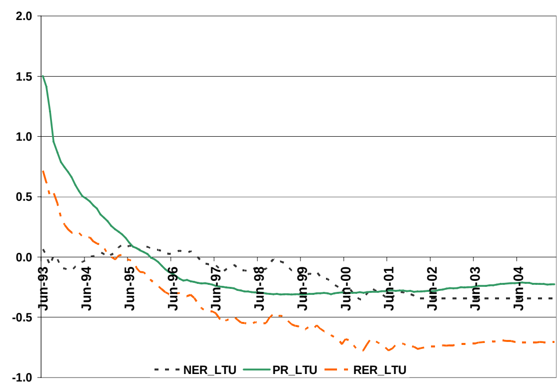
Hungary



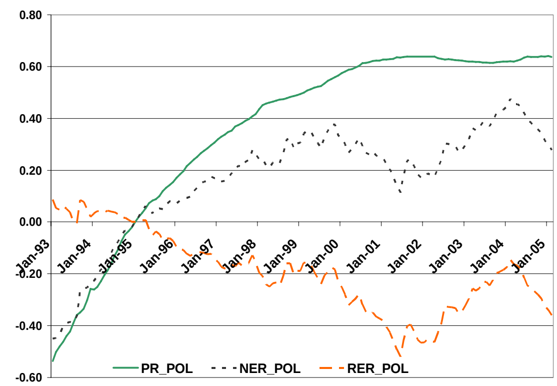
Latvia



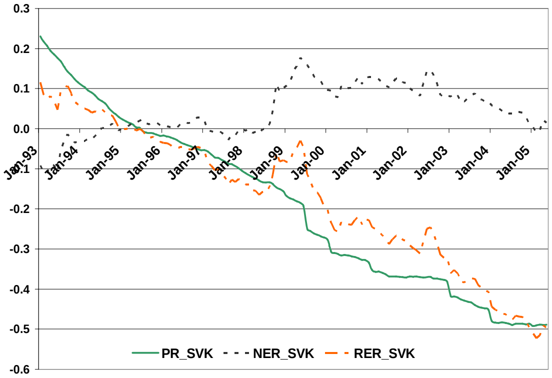
Lithuania



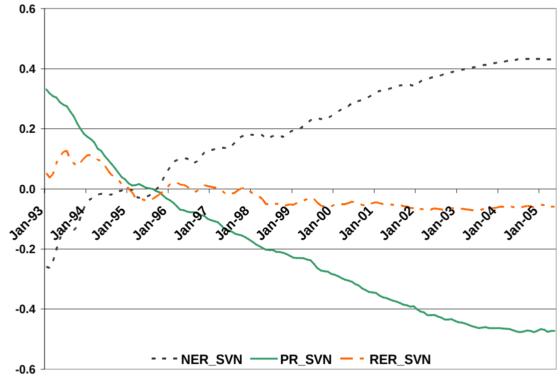
Poland



Slovak Republic



Slovenia



Source: Author's calculation based on data from Eurostat and IMF IFS; Indices were defined so as 1995=1; decrease in the exchange rate index indicates appreciation.

Annex 1: Data Integration Properties

Table A.1.1 Unit Root Tests (NMSs)

EXCHANGE RATE \ TEST	DF-GLS		MZt		MZa		KPSS
	L	FD	L	FD	L	FD	FD
NER_CZE	-1.6	-3.2	-1.6	-2.7	-5.8	-14.7	…
RER_CZE	-2.0	-2.7	-1.9	-2.0	-7.7	-8.6	…
NER_EST	-1.8	**0.0**	-1.7	**0.1**	-6.1	**0.1**	0.1
RER_EST	-1.6	**-0.3**	**-1.4**	**-0.3**	-4.7	**-0.5**	**1.1**
NER_HUN	-1.2	**-1.4**	-2.4	**-1.0**	-12.2	**-1.9**	**1.4**
RER_HUN	-1.4	-2.2	-1.7	**-1.5**	-6.5	**-4.3**	0.2
NER_LVA	-0.3	**-0.7**	-0.1	**-0.7**	-0.1	**-1.1**	0.6
RER_LVA	0.0	**-0.4**	**0.4**	**-0.5**	0.3	**-0.7**	**1.0**
NER_LTU	-2.2	**-0.5**	-1.7	**0.1**	-6.3	**0.2**	0.1
RER_LTU	-1.3	**0.4**	**-0.5**	**0.8**	-1.0	**0.4**	**1.1**
NER_POL	-1.6	**-1.1**	-1.6	**-0.6**	-6.2	**-1.1**	0.2
RER_POL	-1.6	**-1.6**	-1.7	**-0.9**	-5.6	**-1.9**	0.1
NER_SVK	-1.4	-3.2	-1.4	-3.2	-4.9	-20.6	…
RER_SVK	-2.5	-6.2	-2.5	-5.0	-12.8	-49.2	…
NER_SVN	-0.9	**-0.9**	-1.1	**-0.8**	-4.2	**-1.7**	0.6
RER_SVN	-2.6	**-1.3**	-2.6	**-0.8**	-14.4	**-1.5**	0.1

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

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

Table A.1.2 Unit Root Tests (selected OMSs)

EXCHANGE RATE \ TEST	DF-GLS		MZt		MZa		KPSS
	L	FD	L	FD	L	FD	FD
NER_FRA	-2.1	**-1.2**	-1.9	**0.0**	-7.5	**0.0**	0.1
RER_FRA	-2.0	**-1.2**	-2.0	**-0.4**	-7.8	**-0.4**	0.1
NER_DEU	-1.2	**-0.7**	-1.5	**-0.6**	-4.6	**-0.9**	0.3
RER_DEU	-1.2	**-0.5**	-1.5	**-0.4**	-4.3	**-0.5**	0.1
NER_ESP	-1.3	-4.8	-1.1	-4.0	-2.6	-31.4	…
RER_ESP	-1.4	-4.0	-1.2	-3.6	-3.1	-26.6	…
NER_GRC	-1.9	-2.3	-1.9	**-1.7**	-7.6	**-5.9**	0.1
RER_GRC	-1.4	**-1.6**	-1.5	**-1.2**	-5.5	**-3.0**	0.2
NER_ITA	-1.4	**-1.5**	-1.6	**-1.1**	-5.1	**-2.6**	0.2
RER_ITA	-1.5	**-1.4**	-1.6	**-1.1**	-5.1	**-2.4**	0.1
NER_PRT	-1.2	-2.9	-0.9	**-1.3**	-2.0	**-4.0**	0.3
RER_PRT	-1.3	-3.4	-1.1	**-1.6**	-2.8	**-5.7**	0.1

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

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

Table A.1.3 Unit Root Tests (price ratio, countries included in SVAR)

TEST / PRICE RATIO	DF-GLS	MZt_MAIC	Mza_MAIC	MZt_MSIC	Mza_MSIC	KPSS
	L	L	L	L	L	
PR_CZE	-1.38	-2.39	-12.28	…	…	…
PR_HUN	-1.08	-4.03	**-33.75**	-4.03	-33.75	0.37
PR_POL	-0.93	-1.66	-6.49	…	…	…
PR_SVK	-0.67	-0.66	-1.75	…	…	…
PR_SVN	-0.74	-2.93	**-19.61**	-2.93	-19.61	0.28
PR_FRA	-0.69	-0.70	-1.39	…	…	…
PR_DEU	-1.29	-1.24	-3.09	…	…	…
PR_GRC	-1.50	-1.44	-4.34	…	…	…
PR_ITA	-0.67	-0.65	-1.33	…	…	…
PR_PRT	-1.24	-1.01	-2.06	…	…	…
PR_ESP	-0.91	-1.21	-3.69	…	…	…

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

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

Table A.1.4 Unit Root Tests with a Break (NMSs)

Perron&Vogelsang (1992)	NER									
FD	_CZE	_EST	_HUN	_LVA	_LTU	_POL	_SVK	_SVN	_HUN(1)	_LVA(1)
Break point	24	25	67	91	24	84.0	25	22	67	88
Dummy_coeff.	-0.01	0.00	-0.01	0.01	-0.01	0.0	0.00	0.00	-0.01	0.01
T-stat	-2.42	2.86	-3.09	**1.38**	-2.19	**-0.6**	**-0.66**	-2.44	-4.00	2.69
Fixed lag	9	4.00	11	6	3	1	3	5	4	1
ADF	-5.11	-6.47	**-3.59**	-4.71	-7.08	-7.9	-6.86	-4.60	-5.51	-8.31
	RER									
Break point	27	13	34	133	62	84	32	25	67	82
Dummy_coeff.	0.00	0.01	-0.01	0.00	0.01	0.00	-0.01	0.00	0.00	0.01
T-stat	**-0.03**	3.95	-2.39	**0.52**	2.69	**0.68**	-2.19	2.71	-1.63	3.78
Fixed lag	9	5.00	9	9	3	1	9	5	1	1
ADF	-4.87	-4.35	**-4.15**	-4.77	-6.56	-7.81	-4.68	-7.00	-9.64	-8.29

Note: NER - nominal exchange rate; RER – real exchange rate; FD – thirst differences. Bolden magnitudes indicate unexpected results at the 5 per cent significance level; i.e., non-stationarity in first differences. Critical Values for T=150 and k=k(t): 10% (-3.86), 5% (-4.26), 1% (-4.97); kmax=int(12*(T/100)^(0.25)). HUN(1) and LVA(1) is estimated for kmax=8. The results for Estonia, due to the presence of autocorrelation otherwise, are presented for kmax=8.

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

Table A.1.5 Unit Root Tests with a Break (selected OMSs)

Perron&Vogelsang (1992)		NER							
FD		_FRA	_DEU	_ESP	_GRC	_ITA	_PRT		
Break point		27	28	11.0	69	25	17		
Dummy_coeff.		0.00	0.00	-0.01	0.00	-0.01	-0.01		
T-stat		**1.09**	2.08	-3.1	**-1.02**	-3.37	-2.32		
Fixed lag		5	1	3	1	3	1		
ADF		-4.26	-5.90	-5.9	-7.80	-8.22	-4.75		
		RER							
		_FRA	_DEU	_ESP	_GRC	_ITA	_PRT	_FRA(1)	_ESP(1)
Break point		38	33	36	39	25	17	12	24
Dummy_coeff.		0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0
T-stat		**1.3**	2.3	1.9	**1.1**	-3.3	-2.0	**-0.6**	**-1.4**
Fixed lag		11	2	6	1	3	1	0	3
ADF		-2.46	-6.05	**-4.10**	-7.99	-8.36	-4.75	-9.02	-4.92

Note: NER - nominal exchange rate; RER – real exchange rate; FD – thirst differences. Bolden magnitudes indicate unexpected results at the 5 per cent significance level; i.e., non-stationarity in first differences, or/and insignificant break coefficients. Critical Values for T=100 and k=k(t): 10% (-3.81), 5% (-4.19), 1% (-4.91); kmax=int(12*(T/100)^(0.25)). FRA(1) and ESP(1) were tested for kmax=8.

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

Table A.1.6 Unit Root Tests with a Break (price ratio, Hungary and Slovenia)

Perron (1992)	PR_HUN	PR_SVN
	L	L
Break point	137	15
dummy_coeff.	0.0	0.0
t-stat	6.8	5.7
TREND_coeff.	0.0	0.0
t-stat	-39.9	-10.4
Fixed lag	7.0	11.0
y(a-1)	0.0	0.0
ADF	-2.3	-1.4

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

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

Annex 2: Model Specification and Checks

Table A.2.1 Misspecification Tests (NMSs)

Country	Lags	Normality	Autocorrelation		White Hetero	
		Doornik-Hansen	Portmanteau	p-val	no cross terms	cross terms
CZE	6	**0.0004**	0.49		0.14	**0.00**
CZE_DUM	6	**0.0025**	0.16		**0.00**	0.33
HUN	2	**0.0000**	0.06		0.60	0.91
POL	3	0.1754	0.42		0.30	0.25
SVK	1	**0.0000**	0.22		**0.01**	**0.02**
SVN_93	6	**0.0069**	0.07		**0.00**	**0.00**
SVN_96	6	**0.0003**	0.49		0.07	0.18

Note: The line CZE_DUM presents results from estimating the model 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 1993M1 to 2005M8 and from 1996M9 to 2005M8, 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 ECB data

Table A.2.2 Misspecification Tests (OMSs)

Country	Lags	Normality	Autocorrelation		White Hetero	
		Doornik-Hansen	Portmanteau	p-val	no cross terms	cross terms
DEU	1	**0.0025**	0.34		0.18	0.13
DEU_DUM	3	0.1149	0.81		0.45	0.40
ESP	3	0.6203	0.09		0.22	**0.03**
FRA	1	0.762	0.17		0.25	0.24
GRC	3	**0.0000**	0.86		0.50	0.68
ITA	3	**0.0001**	0.14		**0.01**	**0.00**
PRT	1	0.6583	0.36		**0.02**	**0.04**
PRT_DUM	1	0.8198	0.18		**0.03**	0.09

Note: The lines DEU_DUM and PRT_DUM present results from estimating the models specified with the dummy variable detected by structural break tests. 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 ECB data

Table A.2.3 Bai, et. al, Structural Break Test

Country	Sample	Lags	Sup-W-15%	Exp-W-15%	Est Break	90% Conf. Int.
CZE	1996:10-2005:8	6	0.79	0.88	1998:2	(1997:11, 1998:5)
HUN	1993:4-2005:8	2	**0.01**	**0.00**	1997:3	(1996:10, 1997:8)
POL	1995:10-2005:8	3	0.14	0.27	1997:12	(1997:9, 1998:3)
SVK	1993:8-2005:8	1	0.11	0.06	1999:3	(1997:11, 2000:7)
SVN	1997:3-2005:8	6	1.00	1.00	2002:2	(2000:2, 2004:2)
FRA	1993:3-1998:12	1	0.09	0.07	1995:3	(1993:12, 1996:6)
DEU	1993:3-1998:12	1	0.21	0.28	1994:11	(1993:11, 1995:11)
ESP	1993:5-1998:11	3	0.18	0.21	1995:1	(1994:9, 1995:5)
GRC	1993:5-2000:12	3	0.28	0.28	1996:8	(1996:4, 1996:12)
ITA	1993:5-1998:12	3	0.56	0.71	1995:2	(1994:6, 1995:10)
PRT	1993:3-1998:12	1	0.67	0.65	1994:2	(0X, 1995:4)

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

Table A.2.4 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**	0.39	**0.00**	**0.00**	0.39	**0.04**	**0.04**	0.26
RER_CZE	6	29	**0.00**	**0.00**	0.36	**0.01**	**0.01**	0.36	**0.05**	**0.05**	0.33
NER_HUN	2	109	0.36	0.24	0.11	0.31	0.25	0.13	0.32	0.24	0.12
RER_HUN	2	66	0.73	0.69	0.70	0.65	0.67	0.67	0.50	0.47	0.47
NER_POL	3	50	0.09	0.08	0.06	0.11	0.12	**0.04**	0.16	0.15	0.06
RER_POL	3	50	0.19	0.15	0.09	0.25	0.21	0.08	0.32	0.28	0.08
NER_SVK	1	25	0.11	0.29	0.42	0.10	0.32	0.43	0.29	0.47	0.47
RER_SVK	1	25	0.53	0.86	0.90	0.54	0.90	0.93	0.51	0.86	0.87
NER_SVN	6	34	**0.01**	**0.01**	0.08	**0.03**	**0.03**	0.10	0.38	0.37	0.12
RER_SVN	6	28	0.12	0.09	0.34	0.17	0.15	0.37	0.51	0.48	0.46
NER_FRA	1	20	0.93	0.84	0.81	0.89	0.86	0.84	0.95	0.92	0.90
RER_FRA	1	20	1.00	0.97	0.98	0.99	0.98	0.98	1.00	0.99	0.99
NER_DEU	1	25	**0.02**	0.06	0.75	**0.04**	0.12	0.78	**0.05**	0.10	0.69
RER_DEU	1	25	**0.01**	**0.02**	0.64	**0.01**	**0.03**	0.69	**0.02**	**0.03**	0.55
NER_ESP	3	23	0.18	0.18	0.48	0.23	0.27	0.52	0.66	0.68	0.59
RER_ESP	3	23	0.21	0.19	0.45	0.24	0.26	0.47	0.67	0.67	0.58
NER_GRC	3	60	0.91	0.96	0.98	0.91	0.97	0.98	0.76	0.80	0.74
RER_GRC	3	67	0.95	0.95	0.93	0.95	0.96	0.93	0.76	0.76	0.65
NER_ITA	3	23	0.18	0.17	0.67	0.25	0.28	0.75	0.45	0.43	0.46
RER_ITA	3	24	0.15	0.15	0.69	0.20	0.24	0.76	0.28	0.28	0.37
NER_PRT	1	10	**0.01**	**0.01**	0.16	**0.01**	**0.01**	0.17	**0.05**	0.07	0.14
RER_PRT	1	10	0.03	0.05	0.27	0.06	0.08	0.29	0.10	0.13	0.22

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

Annex 3: Variance Decomposition and Impulse Responses (OMSs)

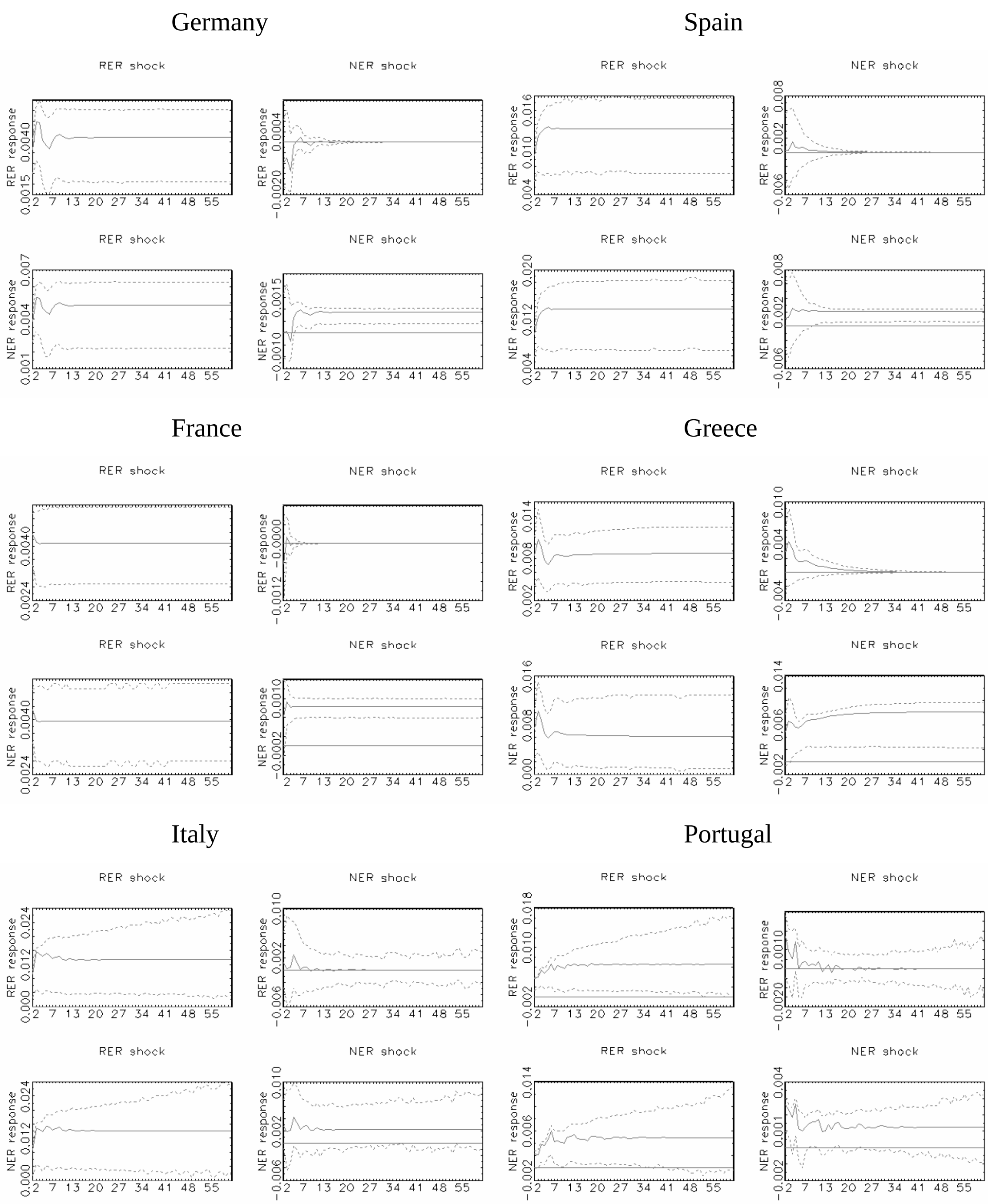
Table A.3.1 Variance Decomposition (OMSs)

Variable	RER				NER			
Variance Decomposition	RER shock		NER shock		RER shock		NER shock	
DEU, dum,ols,95m4								
1-month	95.5	79.5-100	4.5	0.0-20.5	100	82.9-100.0	0.0	0.0-17.1
6-month	97.9	90.7-100	2.1	0.0-9.3	98.3	90.2-99.4	1.7	0.6-9.8
12-month	98.9	95.2-100	1.1	0.0-4.8	97.7	91.9-99.5	2.3	0.5-8.1
24-month	99.5	97.6-100	0.5	0.0-2.4	97.3	92.3-99.5	2.7	0.5-7.7
60-month	99.8	99.0-100	0.2	0.0-1.0	97.0	92.5-99.5	3.0	0.5-7.5
ESP, ols								
1-month	99.9	68.0-100	0.1	0.0-32.0	98.6	66.3-100	1.4	0.0-33.7
6-month	99.6	82.7-100	0.4	0.0-17.3	97.6	77.7-99.9	2.4	0.1-22.3
12-month	99.8	90.6-100	0.2	0.0-9.4	97.6	86.5-99.9	2.4	0.1-13.5
24-month	99.9	95.4-100	0.1	0.0-4.6	97.6	91.5-99.9	2.4	0.1-8.5
60-month	100	98.3-100	0.0	0.0-1.7	97.6	93.8-99.9	2.4	0.1-6.2
FRA, ols								
1-month	99.0	93.4-100	1.0	0.0-6.6	97.9	91.7-100	2.1	0.0-8.3
6-month	99.8	98.3-100	0.2	0.0-1.7	96.1	92.5-98.7	3.9	1.3-7.5
12-month	99.9	99.1-100	0.1	0.0-0.9	95.9	92.4-98.5	4.1	1.5-7.6
24-month	99.9	99.6-100	0.1	0.0-0.4	95.8	92.5-98.6	4.2	1.4-7.5
60-month	100	99.8-100	0.0	0.0-0.2	95.8	92.4-98.4	4.2	1.6-7.6
GRC, ols								
1-month	89.0	56.0-100	11.0	0.0-44.0	70.3	29.6-100	29.7	0.0-70.4
6-month	90.4	68.1-100	9.6	0.0-31.9	63.8	29.2-98.5	36.2	1.5-70.8
12-month	93.9	80.9-100	6.1	0.0-19.1	56.2	23.6-96.0	43.8	4.0-76.4
24-month	96.7	90.5-100	3.3	0.0-9.5	48.3	17.3-95.0	51.7	5.0-82.7
60-month	98.7	96.5-100	1.3	0.0-3.5	41.5	10.9-93.7	58.5	6.3-89.1
ITA, wls								
1-month	98.2	38.0-100	1.8	0.0-62.0	90.3	18.5-100	9.7	0.0-81.5
6-month	99.2	65.1-100	0.8	0.0-34.9	95.7	47.9-99.7	4.3	0.3-52.1
12-month	99.6	78.9-100	0.4	0.0-21.1	96.4	61.8-99.6	3.6	0.4-38.2
24-month	99.8	86.3-100	0.2	0.0-13.7	96.9	69.6-99.8	3.1	0.2-30.4
60-month	99.9	90.5-100	0.1	0.0-9.5	97.2	69.8-99.9	2.8	0.1-30.2
PRT, wls								
1-month	83.1	34.8-100	16.9	0.0-65.2	43.8	0.0-81.3	56.2	18.7-100
6-month	96.2	82.0-99.9	3.8	0.1-18.0	79.2	49.7-98.3	20.8	1.75-50.3
12-month	98.3	90.1-99.8	1.7	0.2-9.9	85.3	60.8-98.8	14.7	1.2-39.2
24-month	99.2	93.9-100	0.8	0.0-6.1	89.8	67.0-99.7	10.2	0.3-33.0
60-month	99.7	95.5-100	0.3	0.0-4.5	92.4	67.2-100.0	7.6	0.0-32.8

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 ECB data.

Figure 3: Impulse Responses (OMSs)

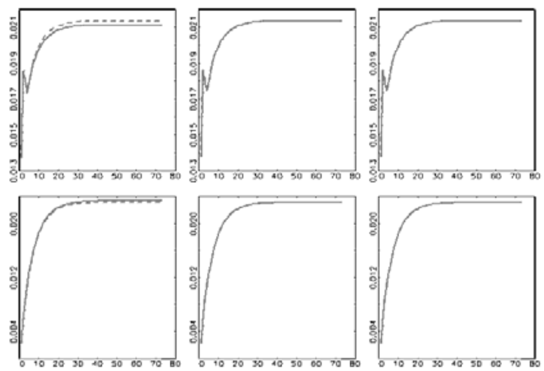


Note: RER – real exchange rate, NER - nominal exchange rate. The top two, out of four, panels presented for each country represent impulse responses of the RER to a unit of real and nominal shocks, respectively; the bottom two panels are impulse responses of the NER to a unit of real and nominal shocks,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

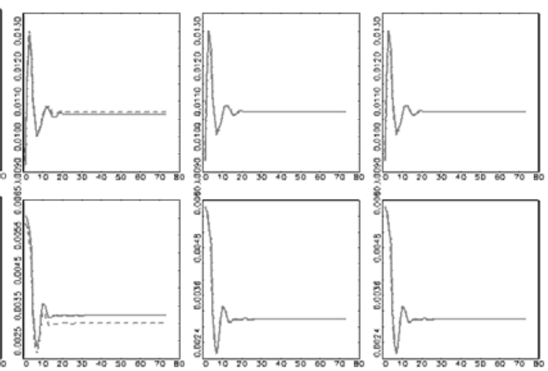
Source: Author's estimations based on IMF IFS and ECB data.

Figure 4: Robustness Checks (NMSs)

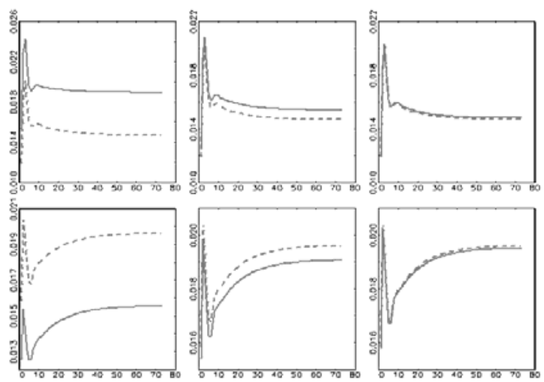
Hungary



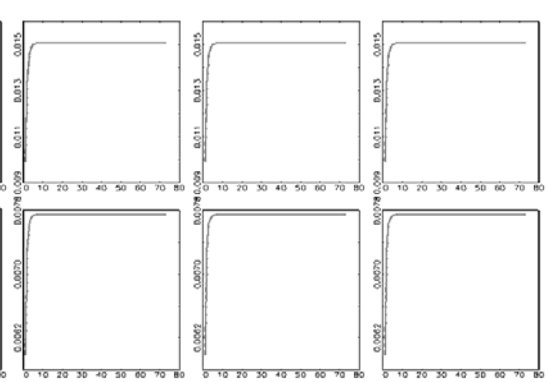
Czech Republic



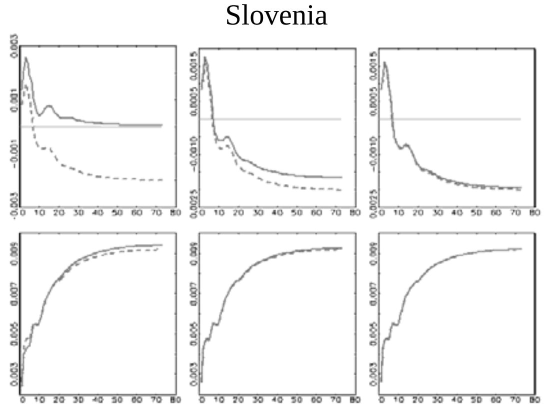
Poland



Slovak Republic



Slovenia



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 ECB data.