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by John Beirne [2]

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In 2009 all ECB publications feature a motif taken from the €200 banknote.

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2 Corresponding author: European Central Bank, Kaiserstrasse 29, D-60311, Frankfurt am Main, Germany; e-mail: John.Beirne@ecb.europa.eu and Brunel University, Economics and Finance Department, Uxbridge, West London, UB8 3PH, United Kingdom; e-mail: John.Beirne@brunel.ac.uk

3 European Central Bank, EU Countries Division, Kaiserstrasse 29, D-60311, Frankfurt am Main, Germany; e-mail: Martin.Bijsterbosch@ecb.europa.eu

Address
Kaiserstrasse 29
60311 Frankfurt am Main, Germany

Postal address
Postfach 16 03 19
60066 Frankfurt am Main, Germany

Telephone
+49 69 1344 0

Website
http://www.ecb.europa.eu

Fax
+49 69 1344 6000

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CONTENTS

Abstract 4
Non-technical summary 5
I Introduction 6
II Exchange rate pass-through: earlier studies and applications to central and eastern Europe 7
III Inflation and exchange rate dynamics in the central and eastern European economies 11
IV Methodology and data 12
V Empirical results 15
VI Conclusions 20
References 22
Appendix 26
European Central Bank Working Paper Series 36

Abstract

This paper provides estimates of the exchange rate pass-through (ERPT) to consumer prices for nine central and eastern European EU Member States. Using a five-variate cointegrated VAR (vector autoregression) for each country and impulse responses derived from the VECM (vector error correction model), we show that ERPT to consumer prices averages about 0.6 using the cointegrated VAR and 0.5 using the impulse responses. We also find that the ERPT seems to be higher for countries that have adopted some form of fixed exchange rate regime. These results are robust to alternative normalisation of the VAR and alternative ordering of the impulse responses.

Keywords: exchange rate pass-through, inflation, central and eastern Europe

JEL Classification: E31, F31

Non-technical summary

The estimation of exchange rate pass-through (ERPT) sheds light on the extent to which fluctuations in nominal exchange rates affect inflation. Whereas there is a vast literature on exchange rate pass-through in advanced economies, there is a more limited number of studies that focus on catching-up economies in central and eastern Europe. Knowledge of the link between nominal exchange rates and inflation in these economies is important as it may shed light on the degree to which inflation convergence vis-à-vis the euro area is sustainable after the conversion rate to the euro is irrevocably fixed (and the dampening impact of the nominal trend appreciation on inflation disappears). Knowledge of the exchange rate pass-through dynamic also has important monetary policy implications. For example, the extent and timing of the pass-through is important for forecasting inflation and thus for monetary policy decision-making.

This paper examines the degree of ERPT to domestic prices in nine central and eastern European EU Member States. The methodological framework used is a cointegrated VAR (vector autoregression) with five variables (the nominal effective exchange rate, consumer prices, producer prices, oil prices and industrial production). This approach provides a coherent means by which to deal with the inherent non-stationarity of the variables of interest in a simultaneous framework. In addition, it enables retention of the important information contained in 'levels' variables, which is particularly relevant for catching-up economies. The paper also uses impulse responses derived from the VECM (vector error correction model).

Our results show that ERPT coefficients for domestic consumer prices average about 0.6 using the cointegrated VAR and 0.5 using the impulse responses. These ERPT estimates are higher than those typically found using VARs in first differences, but in line with other cointegrated VAR based studies of this issue. An explanation for this difference is that cointegrated VARs, as well as retaining 'levels' information, also capture the responsiveness of inflation to exchange rate movements in a long-run equilibrium context. Knowledge of this 'long-run' pass-through is essential in economies that experience a trend appreciation. We also find notable differences across countries with fixed exchange rate regimes compared to those with more flexible regimes, with the former exhibiting larger pass-through degrees than the latter. The results of the cointegration analysis suggest that the size of ERPT is somewhat larger than in the impulse response analysis, which may be due to the longer time horizon in the former. As the adjustment process is not fully completed during the considered time horizon in the impulse response analysis, the long-run effects found in the cointegration analysis should indeed be somewhat higher.

The findings may indicate a stronger link between nominal variables, and thus a higher ERPT to domestic consumer prices, in particular where some form of fixed exchange rate regime is in place. For countries with flexible exchange rate regimes aiming at euro adoption, a high degree of exchange rate pass-through indicates that nominal exchange rate appreciations are likely to lower inflationary pressures, which could help to reduce inflation to below the Maastricht reference value for inflation. As this nominal trend appreciation ceases to exist after the adoption of the euro, the reduction of inflation engineered in this way may not necessarily result in a sustainable convergence of inflation. More broadly, the findings may have clear implications for monetary policy as regards the timing and scale by which exchange rate fluctuations impact upon prices.

I. Introduction

The study of exchange rate pass-through (ERPT) in central and eastern European (post-)transition economies is a relatively new strand to the literature on this issue. One of the reasons for this was perhaps due to the insufficient data spans available and data unavailability for many transition economies. This would, of course, render any econometric work either impossible or not sufficiently robust.

The estimation of ERPT for central and eastern European economies helps to determine the extent to which fluctuations in the nominal exchange rates affect inflation in these countries. Nominal exchange rate fluctuations can have a short-term character, but they could also be of a structural nature associated with the real appreciation trend that so-called catching-up economies tend to experience (due to e.g. Balassa-Samuelson effects). Knowledge of the link between nominal exchange rates and inflation in catching-up countries may shed light on the degree to which inflation convergence vis-à-vis the euro area is sustainable after the conversion rate to the euro is irrevocably fixed (and the dampening impact of the nominal appreciation on inflation disappears).

In terms of the broader context, knowledge of the exchange rate pass-through dynamic has important monetary policy implications. For example, the level of the pass-through provides an indicator of macroeconomic transmission. More specifically, the extent and timing of the pass-through is important in terms of forecasting inflation and thus for monetary policy decision-making. This is particularly important in an inflation-targeting framework.

The added value of this study is two-fold. First, it provides new up-to-date estimates of ERPT for the economies of nine central and eastern European EU Member States.[3] This is important as there is relatively little empirical work done on assessing ERPT for these economies, certainly compared to the vast amount of work done in relation to advanced economies.[4] In addition, the existing studies on the central and eastern European region cover time periods that are increasingly outdated. Second, the methodological framework (cointegrated VAR) is more sophisticated than in most previous studies on ERPT in central and eastern Europe. The cointegration approach provides a coherent means by which to deal with the inherent non-stationarity of the variables of interest in a simultaneous framework. In addition, it enables retention of the important information contained in 'levels' variables, which is particularly relevant for catching-up economies.

This paper is structured as follows. Section II provides an overview of some earlier studies on ERPT, focusing on central and eastern Europe. Section III briefly discusses some stylised facts regarding inflation and exchange rate developments in this group of countries during the past decade. Section IV describes the data and the methodology used and Section V contains the main empirical results. Finally, Section VI concludes.

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[3] The countries covered in this paper are: Bulgaria (BG), the Czech Republic (CZ), Estonia (EE), Hungary (HU), Latvia (LV), Lithuania (LT), Poland (PL), Romania (RO) and Slovakia (SK).

[4] Studies that have been done on transition economies tend to have relatively short data spans. This clearly hinders the robustness of results found. For example, Darvas (2001) examines ERPT for the Czech Republic, Hungary, Poland and Slovenia over the period 1993 to 2000; while Korhonen and Wachtel (2005) examine the issue for a sample of CIS countries over the period 1999 to 2004 (see Table 1).

II. Exchange Rate Pass-Through: Earlier Studies and Applications to Central and Eastern Europe

The issue of ERPT to prices has emerged as a strand of the exchange rate literature over the past thirty years or so, notably since the breakdown of the Bretton Woods system. Exchange rate pass-through reflects the extent to which exchange rate changes are passed on to the local currency prices of traded goods.[5] Goldberg and Knetter (1997; p. 1248) define ERPT as "*the percentage change in local currency import prices resulting from a one percent change in the exchange rate between the exporting and importing countries*". These changes in import prices can subsequently be passed on into producer and consumer prices, thereby affecting the general price level in the economy. In this paper, the focus is on the impact of exchange rate fluctuations on changes in consumer prices.

Incomplete pass-through of exchange rate fluctuations into prices – a typical finding in the literature – can arise from various factors. A seminal contribution in this regard is Dornbusch (1987), who applies industrial organisation models to explain differences in ERPT. In these models, incomplete pass-through can arise from firms that operate in a market characterised by imperfect competition and adjust their mark-up in response to an exchange rate shock. Krugman (1987) refers to such exchange rate induced mark-up adjustments as 'pricing-to-market' strategies. Pricing-to-market behaviour occurs when exporters base their foreign currency export prices on competitive conditions in their foreign markets by allowing profit margins, rather than foreign currency prices, to fluctuate in response to exchange rate fluctuations.

Other explanations of the incomplete nature of exchange rate pass-through relate to wage or price stickiness stemming from staggered price adjustments (e.g. Taylor, 1980), menu costs (Mankiw, 1985) or trade costs such as transport costs or barriers to trade (Obstfeld and Rogoff, 2000). These market frictions imply that the size and the speed of pass-through should decline along the distribution chain, i.e. import prices respond more strongly and rapidly to exchange rate shocks than producer or consumer prices. This is indeed a typical finding in the empirical literature and we incorporate this feature also in our model of ERPT (see Section IV).

A number of different strands in the recent empirical literature on ERPT can be identified that focus on different aspects, such as the time-varying nature of ERPT, the determinants of pass-through and non-linearities and/or asymmetries.[6] As regards the first, a growing body of research has documented a decline in pass-through to domestic prices. Sekine (2006), for example, finds that ERPT has fallen over time in six major industrial countries (both between exchange rate fluctuations to import prices as well as between import price movements and consumer prices). Other studies that deal with this aspect include McCarthy (2007), Amstad and Fischer (2005) and Campa and Goldberg (2005). Related to the time-varying nature of ERPT,

[5] A one-to-one response of import prices to exchange rates is deemed as 'full' or 'complete' ERPT, while a less than one-to-one response is known as 'incomplete' or 'partial' exchange rate pass-through. The extent and speed of pass-through is dependent upon a range of factors including expectations of the duration of the depreciation, the cost of adjusting prices and local demand conditions. A range of previous studies are in broad consensus of incomplete pass-through of about 60%, e.g. Goldberg and Knetter (1997) for the US; Campa and Goldberg (2002) across a range of OECD countries. Specific reasons for incomplete pass-through include imperfect competition or 'pricing to market' whereby foreign producers adjust their mark-up to ensure a stable market share in the domestic economy (e.g. Dornbusch, 1987; Krugman, 1987). Menu costs can also contribute to this (Gosh and Wolf, 2001).

[6] See Wolden Bache (2007) for a more extensive recent overview of the theoretical and empirical literature on exchange rate pass-through.

a second strand of the literature concentrates on the determinants of ERPT. Using new open economy macroeconomic models, Engel (2002) discusses several factors that affect ERPT, such as the degree of price flexibility, the importance of producer currency pricing versus local currency pricing, shipping costs and the share of non-traded goods. Taylor (2000) in particular has put forward the hypothesis that the link between exchange rate fluctuations and prices depends positively on the rate of inflation. Also Choudri and Hakura (2001) note that ERPT seems to be endogenous to alternative exchange rate regimes and smaller when inflation is low. A third, relatively recent strand in the empirical literature focuses on the possibility that the pass-through may not be linear and/or symmetric. Looking at advanced economies, some authors, such as Bussière (2007), find that non-linearities and asymmetries cannot be ignored, although their magnitude seems to differ noticeably across countries.

Whereas there is a vast literature on exchange rate pass-through in advanced economies, there is only a limited number of studies that focus on central and eastern Europe (although there is a growing literature on ERPT in emerging markets). Most studies on central and eastern Europe are country case studies and there are very few systematic cross-country comparisons.[7]

Table 1 provides an overview of recent cross-country studies that cover the central and eastern European region. Three points stand out from the table. First, as regards the methodology the empirical literature seems to adopt either single equation or VAR-based approaches, with the more recent studies using the latter approach. A key reason for doing so is that a simultaneous equation approach does not suffer from misspecification problems that affect single-equation estimates (associated in particular with the likely endogeneity between the variables of interest). Second, the existing empirical studies cover time periods that are increasingly outdated. Given the structural changes taking place in the economies in the region and the likely time-varying nature of ERPT, up-to-date estimates are of key importance for policy makers. Third and finally, the existing studies on the region tend to focus on a narrow group of countries (mainly central Europe), whereas cross-country studies with a wider geographical coverage of the region are very rare.

[7] For example, Dabusinskas (2003) uses a single equation approach and estimates a zero ERPT to consumer prices for Estonia over the period 1995 to 2003; Gueorguiev (2003) uses a VAR in first differences for Romania over the period 1997 to 2002, finding an ERPT to consumer prices of around 35%; and Vyskrabka (2007) also uses a VAR in first differences for Slovakia, estimating a pass-through to consumer prices of around 13% for the period 1999 to 2006.

Table 1 Cross-country studies on ERPT in Central and Eastern European Economies

Author(s)	Methodology	Sample period & data frequency	Country coverage (CEE Member States) and ERPT to consumer prices estimates (in parentheses)
Darvas (2001)	single equation (time varying ECM)	1993-2000 (Q)	CZ (15%), HU (40%), PL (20%), SI (40%)
Mihaljek & Klau (2001)	single equation	1990/94-2000 (Q)	CZ (6%), HU (54%), PL (45%). 13 emerging markets in all.
Bitans (2004)	VAR	1993-2003 (M)	BG (84%), CZ (22%), EE (54%), HU (25%), LV (42%), LT (70%), PL (43%), RO (48%), SK (35%), SI (54%). 13 CEE countries in all.
Campa & Goldberg (2005)	single equation	(1975)-2003 (Q)	CZ (61%), HU (85%), PL (99%). 23 OECD countries in all. Note: figures relate to ERPT to import prices.
Frankel, Parsley & Wei (2005)	single equation (ECM)	1990-2001 (A)	CZ, HU, PL, RO. 76 countries. Note: panel based approach across large panels of rich and developing countries; no individual country estimates provided.
Korhonen & Wachtel (2005)	VAR	1999-2004 (M)	CZ (3%), HU (6%), PL (9%), RO (113%), SK (5%), SI (18%). 27 transition and developing economies in all.
Coricelli, Jazbec & Masten (2006)	cointegrated VAR	1993-2002 (M)	CZ (46%), HU (97%), PL (80%), SI (101%)
Ca'Zorzi, Hahn & Sanchez (2007)	VAR	1988/93-2003 (Q)	CZ (77%), HU (91%), PL (56%). 12 emerging markets in all.
María-Dolores (2008)	single equation (with and without ECM)	2000-2006 (M)	CZ (11%), HU (32%), LV (24%), PL (12%), SK (31%), SI (28%). 7 CEE countries.

A survey of the literature on ERPT in central and eastern Europe yields the following observations. There is some evidence to suggest that previous studies of ERPT in central and eastern European economies come up with a range of results as regards the size of the pass-through. As outlined in Coricelli et al (2006a), studies of one particular country can produce very different pass-through estimates. For example, Dabusinkas (2003) estimates the ERPT to the CPI to be zero in the case of Estonia. However, Bitans (2004) finds that the average pass-through is 54% for Estonia. Using a single equation technique, Mihaljek and Klau (2001) estimate ERPT of 6% for the Czech Republic, 45% for Poland, and 54% for Hungary. A drawback of the approach, however, was the neglect of potential endogeneity issues, as well as volatile exchange rate and inflation conditions over the time period used. Darvas (2001) uses a time-varying based single equation approach, which allows for regime shifts. He estimates ERPT of 15% for the Czech Republic, 20% for Poland, and 40% for Hungary and Slovenia. Bitans (2004) uses a VAR in first differences across a number of CEE Member States. As can be seen from Table 1, the ERPT estimates vary across countries. Using a cointegrated VAR approach, Coricelli et al (2006b) estimate full pass-through for Slovenia and Hungary, 80% for Poland, and 46% for the Czech Republic. The ERPT estimates from previous studies would suggest a high degree of sensitivity of results to selected empirical methodologies and data periods. This clearly inhibits the meaningfulness that can be derived from making direct comparisons of ERPT estimates across different papers.[8]

[8] A point worth making, however, is that the long-run nature of cointegration should suggest that pass-through rates are higher using cointegration-based approaches compared to VARs in first differences.

Another explanation for the range of results in earlier studies could be that the degree of ERPT may depend on factors that have changed during the transition process. For example, Korhonen and Wachtel (2005) note that for countries at a lower stage of economic development, the responsiveness of the CPI to changes in the exchange rate tends to be higher. Studies focusing on the earlier stages of transition, such as Ross (1998) and Kuijs (2002), find indeed that evidence of a low ERPT does not hold for transition economies, due to factors such as a lack of credibility of the relevant national central banks and the extent to which domestic firms are price-takers on global markets. Bitans (2004) concludes that falling inflation rates are closely linked to the decline in pass-through estimates since the late 1990s, although ERPT in central and eastern Europe remains on average larger than for industrial countries. Ca'Zorzi et al (2007) confirm the existence of a positive relationship between the degree of ERPT and inflation for a broader group of emerging economies. Such a decline in pass-through has significant monetary policy implications in that exchange rate fluctuations have smaller effects on prices than previously thought.

The nature of the exchange rate regime in place also appears to have an influence on ERPT. Where a floating exchange rate is in place (with inflation expectations anchored via inflation targeting), there may be a disconnect between the exchange rate and prices (mainly non-tradables).[9] Coricelli et al (2006a) note that where some form of fixed exchange rate regime is in place, any pre-announced currency devaluation provides a nominal anchor for expectations. A stronger link between the exchange rate and prices is evident in this scenario as exchange rate changes may signal price changes. More recent research by Coricelli et al (2006b) suggests that also differences in the degree of monetary policy accommodation can contribute to different pass-through estimates across countries. For example, the latter study finds a very large pass-through from exchange rate depreciation to CPI for Slovenia and Hungary (where accommodative monetary and exchange rate policy was pursued), while a very small impact is found for Poland and the Czech Republic (where real exchange rate stabilization was not a key policy objective).

In terms of the different approaches used in the literature (see Table 1), it is clear that the structural VAR is the most common, whereby the impulse responses of prices are estimated following a structural exchange rate shock. One of the problems with this approach, however, is that exchange rates can change not only due to a shock, but also as a result of policy shifts. Moreover, Coricelli et al (2006b) has noted that the failure of structural VARs to account for policy shifts (e.g. in relation to the exchange rate or inflation) could result in biased and underestimated ERPT estimates. This is of notable importance for the central and eastern European countries. Single equation models can help to overcome the problem that the exchange rate can change for reasons other than stochastic shocks. However, the main drawback of both single equation models and structural VARs is that they fail to recognise cointegration. Given the theoretical co-movement of prices and exchange rates in the long-run, one might expect that cointegration should be taken into account.

As stated, our analysis of ERPT is based on cointegration methods. Differences in approaches employed in previous studies plus a relative dearth of cointegration-based studies on ERPT for the economies examined in this paper makes comparison with other previous work difficult. To the knowledge of the authors, only the Coricelli et al (2006b) paper carries out such an analysis. However, in that paper only the Czech Republic, Hungary, Poland, and Slovenia were considered and the time period was

[9] See Coricelli, Jazbec and Masten (2006b), Darvas (2001), and Kara et al (2005) for further details.

from 1993 to 2002. Given the market reforms in place at the beginning of this period and the differences in macroeconomic performance and stability compared to the sample in our analysis (which runs from 1998 to 2008), we have avoided making direct comparisons with the results of the Coricelli at al (2006a) paper. Nonetheless, we have broadly borne in mind these results in assessing our findings.

III. Inflation and Exchange Rate Dynamics in the Central and Eastern European Economies

HICP inflation in most of the central and eastern European countries in this paper initially declined from relatively high levels at the end of the 1990s (see Figure A1 in the Appendix). Between 2003 and 2005, however, inflation started to rise again in most countries in the region, particularly in Bulgaria, Estonia, Latvia and Lithuania, reaching a peak in 2008 (the end of the reference period in this paper). In Romania and Slovakia, inflation followed a broad downward trend, although recently this trend seems to have turned around.

Over the period since 1998, a trend appreciation of the real exchange rate is evident across all of the countries considered in this study. Where exchange rates were fixed (Bulgaria, Estonia, Latvia and Lithuania), the appreciation came as a result of inflation, while where a more flexible exchange rate regime was in place (the Czech Republic, Hungary, Poland, Romania and Slovakia), the appreciation was due to a combination of nominal appreciation and inflation.

In addition to the real appreciation process, there also seems to be a link between inflation and nominal exchange rate fluctuations. Table 2 provides an overview of the simple (contemporaneous) correlation between annualised monthly changes in inflation and exchange rates since 1999. The expected negatively signed correlation coefficient is evident for all of the countries except the Czech Republic and Poland. The level of the coefficient is notably high in the case of Romania, and reasonably high in Bulgaria, Latvia, and Slovakia.

Table 2 Inflation and Exchange Rate Dynamics, 1999 to 2008

Country	Correlation between Δs in nominal e and inflation	Mean			Standard Deviation		
		e^{nom}	HICP	e^{real}	e^{nom}	HICP	e^{real}
Bulgaria	-0.43	0.44	6.49	-5.56	1.93	3.38	4.15
Czech Republic	0.21	3.66	2.54	1.11	5.08	1.80	4.94
Estonia	-0.03	0.34	4.32	-3.77	2.38	2.30	3.18
Hungary	-0.21	-0.03	6.80	-6.32	5.96	2.54	6.39
Latvia	-0.38	-0.53	5.19	-5.31	4.73	3.80	5.76
Lithuania	-0.05	3.07	2.28	0.89	4.18	2.88	5.80
Poland	0.09	1.31	3.90	-2.44	8.62	3.03	8.46
Romania	-0.86	-10.87	21.28	-23.99	15.62	15.77	20.94
Slovakia	-0.35	1.91	6.35	-3.98	6.39	3.83	7.84

Source: ECB and authors' calculations

As documented in earlier studies, ERPT in central and eastern European economies may differ according to the nature of the exchange rate regime in place. The country charts in Figure A1 in the Appendix seem to confirm this to some extent. In particular for Bulgaria and Estonia, nominal exchange rate fluctuations and inflation exhibit a comparatively aligned co-movement. For countries with more flexible exchange rate regimes, this relationship seems to be less strong on the basis of the charts. These economies, by contrast, display a high correlation between fluctuations in the real and nominal exchange rates.

IV. Methodology and Data

To examine the pass-through of exchange rate fluctuations into consumer prices, our model incorporates features of a distribution chain pricing framework that controls for the impact of supply and demand shocks (similar to McCarthy, 2007; and Hüfner and Schröder, 2002). Given the lack of import price data with the required (monthly) frequency, the distribution chain in our model consists of producer and consumer prices (as a proxy for import prices, we considered using euro area consumer prices, although these are shown to be I(2)). The model thus allows exchange rate fluctuations to affect consumer prices both directly and indirectly via producer prices. Our model implicitly assumes that there is full exchange rate pass-through to import prices (see also Bitans, 2004). This seems a reasonable assumption for central and eastern Europe as the bulk of foreign trade is invoiced in foreign currencies, implying that producer currency pricing may be more relevant than local currency pricing.[10] In addition to exchange rate fluctuations, inflation at each stage of the chain (i.e. consumer and producer price changes) is assumed to be affected by supply shocks and demand shocks. Supply shocks are represented by the oil price in local currency, whereas demand shocks are proxied by industrial production.

Our empirical methodology is based on a cointegrated VAR framework (using the Johansen cointegration procedure).[11] The cointegration approach provides a coherent means by which to deal with the inherent non-stationarity of the variables of interest in a simultaneous framework. In addition, it enables retention of the important information contained in 'levels' variables.[12] Particularly for assessing the sustainability of convergence in central and eastern European countries, it is crucial to distinguish between permanent and transitory changes in the nominal exchange rate, which we are able to do in a cointegrated VAR framework.

While some studies have been done using this type of approach, the nature of the data and variables used in the VAR systems help to differentiate this paper from previous work done.[13] Thus, the VARs constructed for each of the countries are comprised of the following system of variables:

$$x_t = (hicp_t, ppi_t, oil_t, e_t, y_t) \qquad (1)$$

Monthly data were collected for the nine central and eastern European EU Member States in this paper across the following five variables: domestic consumer prices ($hicp_t$), (total) producer prices (ppi_t), oil prices (oil_t), the nominal effective exchange

[10] The prevalence of producer currency pricing suggests that exchange-rate pass-through to import prices would be high.

[11] While there are certain circumstances where a stationary variable can enter a system of otherwise I(1) variables in a cointegration framework, the framework is also more generally valid when each of the variables in the system is integrated of the same order.

[12] This 'levels' information is lost in more traditional first-difference based VARs. This can result in the impulse response functions lacking statistical significance in the long-run (Billmeier and Bonato, 2002).

[13] Examples would include Kenny and McGettigan (1998), Kim (1998), and Menon (1995).

rate (e_t) and industrial production (y_t). The time period runs from January 1995 to April 2008. The start dates differ somewhat across countries depending on data availability (see Table A1 in the Appendix). While a monthly frequency is necessary to ensure a sufficient number of observations, Coricelli et al (2004) also note that monthly data are more informative than lower frequency data in studies such as this as price dynamics are not averaged out.[14]

As a first step, we check the non-stationarity of the data. In order to test this each of the variables are tested for unit roots using the traditional ADF test, but to ensure robustness the order of integration of the variables using other such as the DF-GLS test or the KPSS test is also carried out (which is structured under a different null hypothesis, that of stationarity). Results of the unit root tests of the variables reveal that the majority of the variables to have been generated via an I(1) process (see Table A2 in the Appendix). Stationarity after first-differencing is found in at least two of the three tests undertaken for the vast majority of variables. Three variables are found to be generated via an I(2) process however: Czech HICP, and possibly the HICP for Hungary and Poland. Instead of the HICP, a CPI series was collected for the Czech Republic, Hungary and Poland (all of which are I(1)). In constructing the unit root tests, the variables in levels were tested in the presence of both an intercept and trend. The subsequent tests of first differences included only an intercept given the lack of trending behaviour in the first-differences series.

Subsequently, we apply cointegration tests for each country to check whether long-term relationships exist between the variables. The Johansen test is used to assess whether or not cointegration exists between the system of variables. In order to describe this, supposing that the vector x_t is such that $x'_t = (hicp_t, ppi_t, oil_t, e_t, y_t)'$, consider firstly the following VAR(k) model:

$$x_t = A_1 x_{t-1} + \ldots.. + A_k x_{t-k} + \mu + \psi D_t + \varepsilon_t \tag{2}$$

Equation (2) can be denoted as a VEC (vector error correction) equation as follows (in first-differenced form):

$$\Delta x_t = \Gamma_1 \Delta x_{t-1} + \ldots + \Gamma_{k-1} \Delta x_{t-k+1} + \Pi x_{t-1} + \mu + \psi D_t + \varepsilon_t \tag{3}$$

where $\varepsilon_t \sim Niid(0,\Sigma)$ for $t=1,\ldots,n$; μ is a constant term; D_t is a vector of nonstochastic variables (centred seasonal dummies and intervention dummies); Σ is the variance-covariance matrix of the disturbances; and $\Gamma_i = I - A_1 - \ldots. - A_k$ $(i = 1,\ldots., k-1)$, and $\Pi = -(I - A_1 - \ldots - A_k)$

By notating the system in this way, information is provided on the long-run and short-run relationships, i.e. an indication is provided of how the system responds in both the long-run and the short-run to changes in the x_t. Short-run information is given by the estimates of Γ_i, while long-run information is provided by estimates of Π. The matrix Π can be decomposed as $\Pi = \alpha\beta$`, where the matrix α represents the speed of adjustment to equilibrium, and β represents the cointegrating vectors. Equation (3) can also be augmented to include a constant term to capture trending behaviour, and

[14] Coricelli et al (2004) also make the important point that a monthly frequency in a relatively short data period enables the use of control variables, and therefore a fully-fledged cointegration analysis.

where appropriate a time trend to account for 'catch-up' effects, as well as dummy variables to capture seasonal effects or regime shifts.[15]

Using the Johansen cointegration procedure, two specific test statistics are provided; one relating to the trace test and the other to the maximum eigenvalue test. Both tests yield the number of cointegrating vectors in the system, i.e. the cointegration rank. When the appropriate model has been identified for the system in terms of lag length and residual diagnostics, the coefficients on the β matrix reveal the long-run dynamic while the coefficients on the α matrix reveal the drivers towards the long-run equilibrium. In order to determine the pass-through effect to consumer prices, the results are normalised on consumer prices. To assess the pass-through to producer prices, the system (i.e. the unrestricted cointegrating vector) is normalised on producer prices. The coefficients on the variables indicate the degree of pass-through.[16] Our primary concern is to assess the ERPT to consumer prices. Re-normalisation of the long-run vector is also implemented, however, in order to demonstrate the robustness and consistency of the results found.

As well as providing results based on the unrestricted parameters, a number of restrictions are also imposed on the long-run parameters in order to examine specific hypotheses, as follows:

H1: Full ERPT to domestic prices with zero constraints on other long-run parameters (i.e. test of whether {1 0 0 1 0} holds).

H2: Full ERPT to domestic prices with other parameters unrestricted (i.e. test of whether $\{1\ \eta\ \upsilon\ 1\ \kappa\}$ holds).

H3: Zero ERPT to domestic prices with zero constraints on other long-run parameters (i.e. test of whether {1 0 0 0 0} holds).

H4: Zero ERPT to domestic prices with other parameters unrestricted (i.e. test of whether $\{1\ \eta\ \upsilon\ 0\ \kappa\}$ holds).

The particular advantage of using the VAR-based approach for this purpose is that it also enables joint hypotheses to be tested due to the simultaneous nature of the estimation. As a form of robustness check on the cointegration results, the second empirical technique employed is based on impulse response function analysis. To a certain extent, the choice of impulse response function analysis appropriate is a function of the time series properties of the variables. Assuming that the variables are cointegrated, then the impulse responses can be derived from two types of VAR. The first is a VAR in levels. The second is a VECM that retains information attained from any cointegrating relationships found. While the unrestricted VAR should produce results that are consistent, the approach is generally felt to be inefficient since information on the long-run (i.e. the cointegration) is not taken into account. By contrast, where there exists no cointegration, impulse responses would be derived by a VAR in first differences.[17]

[15] Burke and Hunter (2007) show that the model with intercept is, in general, always used as the base model.
[16] Of course, our focus is on the coefficient of the exchange rate variable.
[17] In order to derive impulse responses across either a VAR or VECM, it is necessary to impose a set of identifying restrictions. These can be based on the contemporaneous effects of shocks, or by imposing long-run restrictions. The Cholesky approach is based on the former.

Thus we will assess the extent and timing of pass-through using the traditional orthogonalised Cholesky decomposition of the residual variance-covariance matrix. This standard type of approach subjects the cointegrated VAR to an orthogonalised shock in one of the variables and the response of the system is assessed.[18] Moreover, the recursive structure embodied in the approach (implying that the variables in the system do not react contemporaneously to shocks imposed on the exchange rate) means that it is important to ensure a correct ordering scheme. We will track the evolution of the pass-through at various time horizons. While we acknowledge the potential drawback of deriving impulse responses directly from the VECM, we feel that it is more appropriate than alternatives.[19] This is primarily due to the fact that the approach is a natural progression from the cointegration analysis, and enables the full set of information from the cointegration analysis to be utilised.

V. Empirical Results

Cointegration Results

Analysis of the standardised residuals from the unrestricted VARs for each country revealed a number of outliers (Table A.3 in Appendix). These were dealt with using intervention dummy variables for the relevant period. These intervention dummies coincide with specific one-off economic events that need to be controlled for in a modelling context. The outliers identified econometrically are explained in terms of their economic rationale, such as changes in administered prices and/or indirect taxes or exchange rate fluctuations related to regime shifts (see the economic explanations of the dummies in Table A.3). In addition, centred seasonal dummies were included to account for seasonality. The lag structure for each model was based on assessment of the AIC in conjunction with ensuring a lack of any residual problems in the unrestricted VAR. Table A.4 details the misspecification tests carried out across each system of variables, indicating no signs of autoregressive behaviour, non-normality, ARCH or heteroskedasticity.[20] The cointegration rank test results (shown in Table A.6) indicate that there exists some variation across the countries as regards the number of cointegrating relationships. The null hypothesis of no cointegration was rejected for all nine countries, with a cointegration rank identified of between one and three. Subsequent tests of identification indicate uniqueness of the cointegrating vectors within the various systems. Of course, where multiple cointegration vectors are found, a need exists to address the issue of identification. Since the Johansen procedure only provides information on the uniqueness of the cointegration space, it is also necessary for $r>1$ to examine the uniqueness of each cointegrating vector. Tests of generically identified structural hypotheses were carried out to examine whether any linear combination of the stationary vectors is also itself stationary. Following Pesaran and Shin (1994), $k=r^2$ restrictions were imposed (so for $r=3$, 9 restrictions and for $r=2$, 4 restrictions), indicating that the vectors are indeed unique.[21]

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[18] See Lutkepohl and Reimers (1992) for more details.

[19] As well as the sensitivity to the ordering of the variables in the system, it also needs to be borne in mind that confidence intervals for the impulse responses from VECMs have not been proven to be asymptotically robust in finite samples.

[20] There are some signs of non-normality across a small proportion of the models. However, as described by Gonzalo (1994) and Cheung and Lai (1993), the trace statistic is robust to non-normality where the non-normality is due to excess kurtosis rather than skewness (see Table A.5).

[21] The following restrictions were tested for r=3 systems: $\beta_{11}=1$, $\beta_{22}=1$, $\beta_{33}=1$, $\beta_{21}= \beta_{32}= \beta_{13}=-1$, $\beta_{31}= \beta_{12}= \beta_{23}=0$. For r=2 systems, the following were tested: $\beta_{11}=1$, $\beta_{21}=0$, $\beta_{12}=0$, $\beta_{22}=1$. At conventional significance levels, the results confirmed unique identification for the cointegrating vectors within the systems.

Even so, as noted by Stephens (2004), the cointegrating vectors are identified only up to some arbitrary normalisation, and as a result the main focus should be on relative signs and magnitudes of the coefficients. In the subsequent cointegration analysis, we have decided to focus on the first cointegrating vector. As noted by Johansen and Juselius (1992), the first cointegrating vector has the highest eigenvalue and is thus "most associated with the stationary part of the model".[22] Moreover, Maddala and Kim (1998) have suggested that the Johansen test can be biased to finding a rank that is too high where unrestricted VAR residuals exhibit signs of non-normality for example. The fact that the vectors are in any case unique within the various systems means that there does not appear to be any spillover effects across the vectors within the cointegration space. Table 3 provides a summary of the number of cointegrating vectors identified across each country, as well as the optimal lag length.

Table 3 Summary of CVAR Characteristics

Country	Rank	VAR lags
Bulgaria	1	7
Czech Republic	3	8
Estonia	3	7
Hungary	2	4
Latvia	2	7
Lithuania	1	7
Poland	3	7
Romania	3	10
Slovakia	3	8

In all of the cases with the exception of the Czech Republic, the most appropriate model appears to be that which includes a trend in the cointegrating equation and permits the intercept to enter both the cointegration space and the VAR, i.e. unrestricted intercept and restricted trend.[23] For the Czech Republic, an intercept and trend term is restricted to the cointegration space. In cointegration analysis, the use of unrestricted intercepts and restricted trends is consistent with data that exhibit some form of trending behaviour. Over the time period in question for the economies considered, the process of catch-up is in line with this proposition.[24] As well as this economic argument, the choice of the deterministic component for each model is also justified econometrically in terms of the information provided by the unit root tests and the residual diagnostics of the VARs in unrestricted form (see Tables A.2 and A.4). Figure A.2 provides an indication of the stability of the cointegrating vectors identified. Over a 36-month window, recursive estimates of the eigenvalues are generally constant.

[22] MacDonald and Marsh (1997) also focus their analysis of PPP for Germany, the UK and Japan on the first significant cointegrating vector.
[23] The inclusion of the trend variable on grounds of model specification may also suggest that Balassa-Samuelson effects for the economies in question may have been apparent for the data period under investigation.
[24] This holds for all economies except the Czech Republic, where perhaps catch-up has not been as prevalent in relation to the other economies considered.

Table 4 Long-run Matrix: Coefficients of First Cointegrating Vector

Country	CV normalised on Domestic Consumer Prices					
	hicp	ppi	oil	e	y	t
Bulgaria	1.000	0.503* (0.238)	0.027* (0.012)	*-0.698** *(0.298)*	0.372 (0.339)	0.003* (0.001)
Czech Republic	1.000	0.443* (0.218)	0.094* (0.014)	*-0.505** *(0.165)*	-0.761 (0.667)	0.000 (0.001)
Estonia	1.000	0.973* (0.153)	0.032* (0.016)	*-0.925** *(0.193)*	0.076 (0.073)	0.004* (0.000)
Hungary	1.000	1.087* (0.229)	0.001 (0.029)	*-0.634** *(0.209)*	-0.251* (0.091)	0.001* (0.000)
Latvia	1.000	0.384* (0.083)	1.060* (0.023)	*-0.969** *(0.123)*	-0.048 (0.150)	0.002* (0.000)
Lithuania	1.000	0.340* (0.127)	0.492* (0.046)	*-0.440** *(0.115)*	0.087 (0.079)	-0.001* (0.000)
Poland	1.000	1.006* (0.479)	0.015 (0.068)	*-0.469** *(0.203)*	-0.232 (0.315)	0.002* (0.001)
Romania	1.000	0.490* (0.205	0.079* (0.029)	*-0.436** *(0.162)*	-0.290 (0.204)	0.001* (0.001)
Slovakia	1.000	0.942* (0.457)	0.064* (0.037)	*-0.370** *(0.142)*	0.443 (0.384)	0.000 (0.000)

(Note: Standard errors in parentheses, * denotes significance at the 5% level or below)

The unrestricted long-run parameters for each system in Table 4 comprise those present in the first (most statistically significant) cointegrating vector. Prior to analysing the coefficients, a first step is to examine the appropriateness of the normalisation imposed on domestic consumer prices. Following Boswijk (1996), a test of the imposition of a zero restriction on domestic consumer prices in the long-run matrix is rejected in all cases (see Table A.7). This helps to provide confidence in the normalisation scheme.

The long-run parameters are well-founded based not only on well-specified models but also the absence of signs of weak exogeneity (see Table 5 for the α matrix loading factors). In unrestricted form, it is clear that the signs of the parameters appear in most cases to accord with priors. For example, a positive coefficient is observed on both producer prices and oil prices, while a negative coefficient is observed for the exchange rate series[25]. Thus, a rise in PPI and oil prices is associated with a rise in domestic consumer prices, while a depreciation of the domestic currency is associated with a rise in consumer prices. There appears to be some inconsistency regarding the sign on industrial production. In almost all cases, however, this parameter lacks statistical significance. The trend term is significant in all cases except the Czech Republic and Slovakia, suggesting Balassa-Samuelson effects have had a significant role on domestic prices for the majority of the transition economies. As a form of robustness check on these long-run coefficients, an alternative normalisation scheme was carried out on the PPI. The results from this (set out in Table A.8) appear to confirm that the ERPT estimates are not sensitive to a change in the normalising variable. The Boswijk normalisation test is also carried out for this normalisation scheme, the results of which also validate PPI as a suitable normalising variable (see Table A.9). Moreover, the results normalised on PPI confirm the typical finding in the literature that ERPT to producer prices is higher than that to consumer prices across all countries examined (see Section II).

[25] A negative coefficient on the NEER implies a depreciation.

The degree of ERPT appears to be most prevalent in the Bulgaria, Estonia, and Latvia, where almost a 1:1 relationship can be observed. For example, a 1% fall in the NEER (i.e. a depreciation) for Latvia increases domestic consumer prices by 0.97%. From the ERPT estimates in Table 4, the average across all countries is 0.605. Averaging across the fixed exchange rate countries (i.e. Bulgaria, Estonia, Latvia and Lithuania) yields a pass-through to domestic prices of 0.758. Across the more flexible exchange rate regime economies (the Czech Republic, Hungary, Poland, Romania, Slovakia), the average ERPT is 0.483. Lower pass-through estimates appear to be evident where inflation has become more subdued over time (e.g. Czech Republic). The nature of the exchange rate regime in place may have had a strong role to play in contributing to low inflation. A fixed regime should imply a strong relationship between the exchange rate and nominal variables (e.g. prices), and therefore a high ERPT. On the other hand, a more flexible regime should be associated with a lower extent of ERPT as the link between the exchange rate and prices weakens. This would appear to be borne out in the empirical work undertaken.

Table 5 Loading Factors

Country	hicp	ppi	oil	e	y
Bulgaria	-0.088* (0.028)	0.304* (0.089)	-0.381 (0.942)	-0.127* (0.042)	0.532* (0.264)
Czech Republic	-0.131* (0.057)	-0.112* (0.043)	-0.855 (0.956)	0.687* (0.131)	0.361 (0.200)
Estonia	-0.044* (0.029)	0.097* (0.029)	-0.028 (0.516)	0.129* (0.033)	0.374* (0.151)
Hungary	0.045* (0.021)	-0.095* (0.030)	0.665* (0.328)	0.259* (0.049)	0.226* (0.092)
Latvia	0.093* (0.021)	-0.086* (0.022)	1.237* (0.300)	0.128* (0.030)	0.059 (0.077)
Lithuania	0.060* (0.026)	0.302* (0.111)	0.478 (0.516)	0.159* (0.078)	0.743* (0.356)
Poland	-0.037* (0.005)	0.036* (0.005)	0.243* (0.095)	0.082* (0.024)	0.010 (0.028)
Romania	0.099* (0.026)	0.113* (0.036)	0.517* (0.135)	0.093* (0.049)	-0.073* (0.041)
Slovakia	-0.078* (0.031)	0.066* (0.023)	1.073* (0.334)	0.107* (0.050)	-0.263* (0.076)

(Note: Standard errors in parentheses, * denotes significance at the 5% level or below)

The loading factors reveal the speed with which the long-run equilibrium is achieved. A lack of significance on these parameters indicates the presence of weak exogeneity. This means that the variable does not respond to or correct for deviations to the long-run equilibrium. Across half of the countries, oil prices appear to be weakly exogenous, perhaps signifying a strong role played by oil prices in affecting the price, exchange rate and broader economic growth dynamics of the CEECs. The industrial production variable is statistically significant for the majority of countries considered in the α matrix. Thus, there is an important role played by this variable in reacting to restore the long-run equilibrium as the speeds of adjustment are particularly high across many countries. Thus, despite the lack of significance for industrial production in the long-run matrix, it appears to have a key role to play in driving the short-run structure.

Table 6 Restrictions on long-run parameters to examine full and zero pass-through of exchange rate on domestic prices (λ^2)

Country	Full Pass-through		Zero Pass-through	
	H1	H2	H3	H4
Bulgaria	32.83 (0.00)	0.35 (0.56)	33.08 (0.00)	8.33 (0.02)
Czech Republic	17.01 (0.00)	11.93 (0.01)	17.65 (0.00)	22.27 (0.00)
Estonia	30.57 (0.00)	0.43 (0.50)	28.59 (0.00)	12.00 (0.01)
Hungary	10.48 (0.02)	7.43 (0.03)	20.28 (0.00)	33.93 (0.00)
Latvia	18.10 (0.01)	5.13 (0.05)	12.98 (0.01)	23.97 (0.00)
Lithuania	56.60 (0.00)	0.19 (0.67)	53.31 (0.00)	13.39 (0.00)
Poland	36.63 (0.00)	40.62 (0.00)	10.82 (0.02)	40.58 (0.00)
Romania	44.78 (0.00)	12.82 (0.01)	10.31 (0.02)	19.25 (0.00)
Slovakia	21.30 (0.00)	28.36 (0.00)	9.19 (0.03)	19.23 (0.00)

Notes: Restrictions based on Likelihood Ratio tests with a chi-squared distribution, with the number of degrees of freedom equal to the number of restrictions imposed; p-values in parentheses.

Across all countries, it is clear that the H3 and H4 are rejected, implying that EPRT is not zero for any of the CEECs. H1 is also rejected for all countries, indicating that full ERPT is rejected when other variables in the system are constrained to have no effect on domestic consumer prices. When the other variables in the system are left unrestricted, however, it is apparent that full pass-through takes place for Bulgaria, Estonia, Latvia and Lithuania.

Impulse Response Function Analysis

This empirical approach comprises assessing the response of the domestic HICP to shocks imposed on the exchange rate. Using the traditional orthogonalised impulse response function analysis (a standard Cholesky decomposition), the framework is based on the following variable ordering:

OIL → NEER → IND PROD → PPI → HICP

This ordering is consistent with a number of other studies that have used these or similar variables in deriving impulse responses. For example, an influential paper by McCarthy (2007) uses this type of ordering, as do Ca'Zorzi et al (2007).[26] The rationale for the ordering chosen is based on a progression from the variable that is most exogenous to that which is less exogenous. Oil prices, as the most exogenous variable, are ordered first in the scheme therefore, while domestic consumer prices are ordered as the last variable in the scheme. While one could have arguments for a range of alternative ordering schemes, we feel that the scheme used in our analysis is both consistent with the literature and reflective of a natural shock progression. For example, in a system of 5 variables, there are 120 possible ordering schemes, and estimation thus requires selecting a preferred ordering based on economic theory. Moreover, we follow the suggestion of Sims (1980) by providing an alternative ordering scheme as a form of *prima facie* robustness check. Table 7 provides estimates of the impulse responses at 6, 12, 24, and 48 month time horizons. We only

[26] Given that we are aware of the possible sensitivity of the Cholesky approach to the ordering of the variables and have carried out the analysis the following alternative ordering scheme: *OIL → IND PROD → NEER → PPI → HICP*. These results (reported in Table A.10) do not change the broad pattern and magnitude of the results reported in Table 6, in what we believe to be the more natural ordering of the variables.

report the estimates for the response of HICP to an orthogonalised 1% shock imposed on the NEER. Also reported for ease of comparison are the ERPT estimates from the earlier cointegration analysis.

Table 7 Summary of ERPT Estimates

Country	Response of HICP to 1% NEER Shock				Cointegration
	6 months	*12 months*	*24 months*	*48 months*	*Long-Run*
Bulgaria	0.203	0.213	0.320	0.360	0.698
Czech Republic	0.246	0.384	0.414	0.434	0.505
Estonia	0.060	0.159	0.572	0.598	0.925
Hungary	0.089	0.239	0.367	0.396	0.634
Latvia	0.356	0.439	0.509	0.619	0.969
Lithuania	0.147	0.211	0.335	0.460	0.440
Poland	0.267	0.360	0.397	0.400	0.469
Romania	0.135	0.177	0.230	0.340	0.436
Slovakia	0.046	0.176	0.389	0.391	0.370

In terms of comparisons with the pass-through estimates from the earlier cointegration analysis, it is apparent that the estimates from the impulse response function analysis are somewhat lower. This may be due to the longer time horizon in the cointegration analysis. As the adjustment process is not fully completed during the considered time horizon in the impulse response analysis, the long-run effects found in the cointegration analysis should indeed be somewhat higher. The full extent of the pass-through appears to largely take place within 24 months, although there continues to be some minor effects up until 48 months. The impulse response estimates at 48 months are extremely close to the cointegration estimates for five of the countries: the Czech Republic, Lithuania, Poland, Romania and Slovakia. For the remaining countries, the impulse response estimates are roughly about two-thirds of the cointegration estimates. A similar picture emerges as that with the cointegration analysis as regards the scale of the ERPT estimates for fixed versus floating exchange rate regime economies. After 48 months, the average pass-through to prices from VECM impulse responses is 0.445. For the fixed regime countries, the average ERPT is 0.509. For the more flexible regime countries, average pass-through is estimated at 0.392.

VI. Conclusions

This paper assesses the degree of ERPT to consumer prices using both a multivariate cointegration approach and impulse responses derived from the VECM for nine central and eastern EU Member States. The particular advantage of the methodology is that it enables a coherent means by which to deal with the non-stationarity inherent in the variables. This approach also ensures that none of the 'levels' information is lost. Using a fully-fledged system of five variables, we show that ERPT to domestic consumer prices averages at about 0.6 using the cointegrated VAR and 0.5 using the impulse responses. These ERPT estimates are higher than those typically found using VARs in first differences, but broadly in line with Coricelli et al (2006b), who use the same methodology. An explanation for this difference is that cointegrated VARs, as well as retaining 'levels' information, also capture the responsiveness of inflation to exchange rate movements in a long-run equilibrium context. Knowledge of this 'long-run' pass-through is essential in converging economies that experience a trend appreciation.[27]

We find notable differences across countries with fixed exchange rate regimes compared to those with more flexible regimes. For example, our cointegration results indicate that ERPT for the four fixed exchange rate regime countries (Bulgaria, Estonia, Latvia, and Lithuania) averages 0.758. Moreover, for each of these countries, a hypothesis test for full pass-through cannot be rejected. The impulse responses yield an average pass-through for these countries of 0.509 after 48 months. For the countries with more flexible regimes (the Czech Republic, Hungary, Poland, Romania and Slovakia), the long-run pass-through averages 0.483, and full pass-through is rejected in all cases. Impulse responses after 48 months indicate an average pass-through of 0.392. The results of the cointegration analysis suggest that the size of ERPT is somewhat larger than in the impulse response analysis, which may be due to the longer time horizon in the cointegration analysis. As the adjustment process is not fully completed during the considered time horizon in the impulse response analysis, the long-run effects found in the cointegration analysis should indeed be somewhat higher. The results are robust to both an alternative normalisation scheme on the cointegrated VAR and an alternative variable ordering for the impulse response function analysis.

The findings may indicate a stronger link between nominal variables, and thus a higher ERPT to domestic consumer prices, in particular where some form of fixed exchange rate regime is in place. For countries with flexible exchange rate regimes aiming at euro adoption, a high degree of exchange rate pass-through indicates that nominal exchange rate appreciations are likely to lower inflationary pressures, which could help to reduce inflation to below the Maastricht reference value for inflation. As this nominal trend appreciation would cease to exist after the adoption of the euro, the reduction of inflation engineered in this way may not necessarily result in a sustainable convergence of inflation. More broadly, the findings may have clear implications for monetary policy as regards the timing and scale by which exchange rate fluctuations impact upon prices.

[27] More specifically, Coricelli et al (2006b) examine ERPT for the Czech Republic, Hungary, Poland, and Slovenia over the period 1993 to 2002. They estimated pass-through rates for the countries cited as follows: CZ: 0.46; HU: 0.97; PO: 0.80; SI: 1.01. These pass-through rates are high relative to previous studies carried out using VARs in first differences. In our study, we have also identified relatively high long-run pass-through rates. However, due to differences in time periods used in our study, as well of the range of countries covered, it is difficult to make direct comparisons with previous work done in this area.

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Appendix: Figure A1 HICP Inflation and Exchange Rate Dynamics

(year-on-year changes in %, HICP inflation – left-hand scale; NEER: nominal-effective exchange rate – right-hand scale; REER: real-effective exchange rate (CPI-based) – right-hand scale. A decrease is an appreciation of the respective currency)

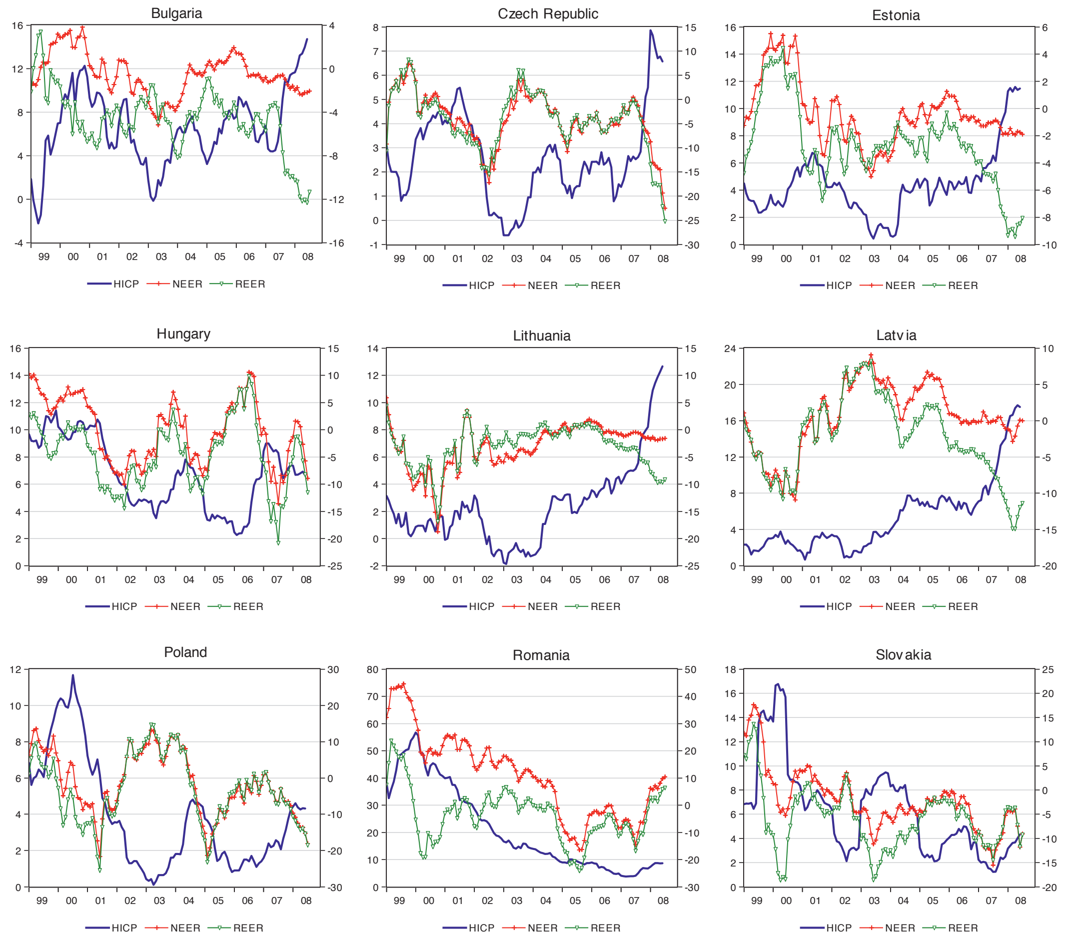


Table A1 Sample Periods by Country

Country	Period	Observations
Bulgaria	2000M01-2008M04	100
Czech Republic	1998M01-2008M04	124
Estonia	1998M01-2008M04	124
Hungary	1998M01-2008M04	124
Latvia	1996M01-2008M04	148
Lithuania	1998M01-2008M04	124
Poland	1996M01-2008M04	148
Romania	1995M01-2008M04	160
Slovakia	1998M01-2008M04	124

Table A2 Unit Root Tests

Country	ADF		DFGLS		KPSS	
	Level	1st Diff.	Level	1st Diff.	Level	1st Diff.
Nominal Effective Exchange Rates						
Bulgaria	-1.90	-8.00*	-1.64	-6.91*	0.15	0.21
Czech Republic	-2.01	-9.47*	-2.16	-2.87*	0.11	0.11
Estonia	-1.96	-8.01*	-1.42	-7.76*	0.19	0.26
Hungary	-3.09	-8.53*	-1.31	-7.47*	0.13	0.34
Latvia	-1.77	-7.67*	-1.01	-6.99*	0.22	0.46***
Lithuania	-1.11	-8.73*	-1.69	-8.32*	0.30	0.20
Poland	-1.81	-7.83*	-1.57	-7.67*	0.18	0.33
Romania	-2.07	-4.29*	-0.56	-3.89*	0.33	0.16
Slovakia	-2.13	-7.74*	-0.40	-6.37*	0.28	0.17
HICP						
Bulgaria	-3.27	-7.59*	-2.53	-6.57*	0.12	0.16
Czech Republic	-2.81	-1.17	-3.08	-1.22	0.14	0.16
Estonia	1.95	-8.34*	-0.19	-2.06**	0.17	0.44***
Euro area	-2.50	-2.33	-1.44	-1.06	0.16	0.46***
Hungary	-2.08	-7.87*	-1.06	-1.01	0.28	0.38***
Latvia	2.95	-7.37*	-1.14	-1.85***	0.33	0.39***
Lithuania	3.84	-7.97*	0.83	-7.76*	0.27	0.25
Poland	-1.86	-6.99*	-0.73	-0.05	0.27	0.65**
Romania	-2.29	-8.13*	-1.07	-2.73*	0.33	0.30
Slovakia	-1.22	-9.83*	-0.73	-9.65*	0.29	0.43***
CPI						
Czech Republic	-3.79	-3.94*	-1.37	-3.79*	0.24	0.52**
Hungary	-2.93	-3.95*	-0.84	-3.26*	0.30	0.48**
Poland	-2.35	-3.76*	-1.69	-2.21*	0.35	0.39***
PPI						
Bulgaria	-0.38	-8.20*	-1.32	-8.24*	0.26	0.29
Czech Republic	-2.09	-6.11*	-2.15	-2.10**	0.07	0.08
Estonia	0.74	-7.80*	-1.01	-7.78*	0.19	0.50**
Hungary	-2.00	-7.94*	-1.42	-7.83*	0.17	0.19
Latvia	2.54	-3.22**	-0.48	-2.69*	0.34	0.19
Lithuania	-1.51	-8.21*	-1.31	-2.48**	0.20	0.32
Poland	-2.54	-6.04*	-0.78	-5.25*	0.29	0.69**
Romania	-1.17	-5.47*	-0.57	-5.32*	0.37	0.27
Slovakia	-2.67	-8.12*	-2.71	-7.34*	0.12	0.07
Industrial Production						
Bulgaria	-2.45	-15.17*	-1.66	-14.46*	0.13	0.15
Czech Republic	-3.15	-16.93*	-0.53	-3.50*	0.31	0.23
Estonia	-4.39	-10.25*	-1.31	-11.32*	0.12	0.26
Hungary	-2.29	-16.29*	-1.49	-1.68***	0.13	0.13
Latvia	-2.19	-14.99*	-1.18	-2.81*	0.24	0.21
Lithuania	-4.37	-13.84*	-3.58	-10.93*	0.18	0.40***
Poland	-1.52	-18.95*	-1.63	-16.43*	0.27	0.17
Romania	-1.36	-14.98*	-1.56	-14.78*	0.29	0.22
Slovakia	-2.58	-12.43*	-0.60	-1.79***	0.15	0.39***
Oil						
Oil price index	-2.18	-10.36	-2.08	-7.83*	0.11	0.10

(Note: The tests were performed on the logs of the series for levels including an intercept and trend. The critical values at 1%, 5%, and 10% levels respectively are: ADF: -4.06, -3.46, -3.15; DFGLS: -3.60, -3.05, -2.76; KPSS: 0.22, 0.15, 0.12. For the first-differences, the tests included only an intercept and were based on the following critical values at the 1%, 5%, and 10% levels respectively: ADF: -3.48, -2.89, -2.58; DFGLS: -2.58, -1.94, -1.61; KPSS: 0.74, 0.46, 0.35. *, **, and *** respectively refer to significance at the 1%, 5%, and 10% levels).

Table A.3 Intervention Dummy Variables

Country	Period
Bulgaria	None
Czech Republic	2000M1
Estonia	2005M5
Hungary	2003M6, 2006M1
Latvia	None
Lithuania	2004M5
Poland	1998M1, 2006M9
Romania	1995M11, 1997M1, 2006M9
Slovakia	1998M10, 1999M7, 2003M1

Economic explanations for the dummies:

CZ	2000M1	Increase in administered prices
EE	2005M5	Decline in consumer prices (mainly lower motor fuel prices)
HU	2003M6	Sharp depreciation of the Hungarian forint following loss of investor confidence
HU	2006M1	Reduction in indirect taxes
LT	2004M5	Increase in administered prices and indirect taxes (mostly related to EU accession)
PL	1998M1	Increase in administered prices
PL	2006M9	Depreciation of the Polish zloty
RO	1995M11	Depreciation of the Romanian leu
RO	1997M1	Liberalisation of exchange rate regime and abolition of price controls
RO	2006M9	No clear economic explanation (possibly lower food prices due to a good harvest)
SK	1998M10	Depreciation of the Slovak koruna following the suspension of the fixed exchange rate regime
SK	1999M7	Increase in administered prices and indirect taxes
SK	2003M1	Increase in administered prices and indirect taxes

Table A.4 Misspecification Tests

Country	Variable	$AR_{(1-7)}$	Normality	$ARCH_{(1-7)}$	Hetero
Bulgaria	**hicp**	1.801[0.124]	1.973 [0.373]	0.272 [0.946]	70.031 [0.544]
	ppi	2.225 [0.061]	1.734 [0.420]	0.355 [0.902]	80.952 [0.220]
	oil	1.470 [0.214]	0.987 [0.611]	0.745 [0.617]	81.388 [0.210]
	e	2.148 [0.069]	2.793 [0.248]	0.338 [0.912]	55.375 [0.927]
	y	1.188 [0.333]	1.760 [0.415]	0.255 [0.954]	68.021 [0.611]
	System	1.265 [0.117]	7.240 [0.703]		1112.6 [0.239]
Czech Republic	**hicp**	1.440 [0.213]	1.625 [0.444]	0.115 [0.997]	86.342 [0.350]
	ppi	1.169 [0.339]	2.070 [0.355]	0.344 [0.928]	75.259 [0.688]
	oil	1.123 [0.279]	2.657 [0.265]	0.880 [0.531]	76.757 [0.643]
	e	1.779 [0.094]	1.249 [0.535]	0.464 [0.854]	81.022 [0.510]
	y	1.920 [0.098]	0.488 [0.784]	0.159 [0.992]	85.092 [0.386]
	System	1.256 [0.122]	6.928 [0.732]		1217.2 [0.597]
Estonia	**hicp**	1.394 [0.230]	2.317 [0.314]	0.327 [0.937]	82.111 [0.476]
	ppi	0.371 [0.915]	0.646 [0.724]	0.509 [0.823]	80.628 [0.522]
	oil	1.339 [0.254]	2.632 [0.268]	0.419 [0.885]	83.731 [0.426]
	e	0.643 [0.718]	0.672 [0.715]	1.819 [0.110]	84.719 [0.397]
	y	1.847 [0.100]	2.441 [0.295]	0.390 [0.903]	79.400 [0.561]
	System	1.257 [0.125]	12.948 [0.227]		1241.7 [0.402]
Hungary	**hicp**	1.304 [0.262]	18.849 [0.001]*	0.632 [0.727]	0.577 [0.954]
	ppi	1.478 [0.189]	13.912 [0.001]*	0.195 [0.986]	0.451 [0.993]
	oil	0.402 [0.898]	5.014 [0.082]	0.635 [0.725]	0.465 [0.991]
	e	1.870 [0.088]	5.146 [0.076]	0.129 [0.996]	0.566 [0.959]
	y	0.258 [0.968]	4.930 [0.085]	0.428 [0.883]	0.433 [0.995]
	System	0.945 [0.648]	51.206 [0.000]*		0.395 [1.000]
Latvia	**hicp**	1.586 [0.151]	0.111 [0.946]	0.283 [0.959]	0.258 [1.000]
	ppi	0.793 [0.595]	0.653 [0.721]	0.442 [0.873]	0.252 [1.000]
	oil	0.426 [0.884]	0.737 [0.692]	0.187 [0.988]	0.204 [1.000]
	e	1.893 [0.080]	5.509 [0.064]	0.621 [0.737]	0.294 [0.999]
	y	1.096 [0.373]	5.223 [0.073]	0.228 [0.978]	0.300 [0.999]
	System	1.169 [0.170]	12.389 [0.260]		0.192 [1.000]
Lithuania	**hicp**	0.902 [0.512]	0.491 [0.782]	0.152 [0.993]	71.353 [0.499]
	ppi	0.785 [0.603]	3.684 [0.159]	0.394 [0.901]	72.983 [0.446]
	oil	1.745 [0.118]	7.939 [0.019]**	0.264 [0.965]	69.806 [0.551]
	e	2.098 [0.059]	0.145 [0.930]	1.288 [0.277]	76.756 [0.329]
	y	0.641 [0.720]	0.533 [0.766]	0.166 [0.991]	73.344 [0.434]
	System	1.190 [0.158]	8.807 [0.551]		1086.4 [0.439]
Poland	**hicp**	0.749 [0.630]	12.882 [0.002]*	2.129 [0.053]	0.139 [0.999]
	ppi	0.703 [0.669]	5.233 [0.073]	0.743 [0.637]	0.044 [1.000]
	oil	1.187 [0.322]	0.084 [0.958]	0.264 [0.966]	0.098 [1.000]
	e	0.595 [0.757]	5.594 [0.061]	0.617 [0.739]	0.053 [1.000]
	y	1.247 [0.289]	3.753 [0.153]	0.162 [0.992]	0.063 [1.000]
	System	0.829 [0.895]	34.614 [0.000]*		1159.5 [0.046]**
Romania	**hicp**	1.284 [0.279]	4.953 [0.084]	0.124 [0.996]	114.87 [0.181]
	ppi	1.439 [0.213]	1.736 [0.420]	0.918 [0.503]	101.52 [0.495]
	oil	0.621 [0.736]	9.686 [0.008]*	0.281 [0.958]	105.92 [0.376]
	e	0.954 [0.475]	5.753 [0.056]	0.375 [0.912]	98.029 [0.593]
	y	0.539 [0.800]	29.647 [0.000]*	0.265 [0.964]	83.187 [0.913]
	System	1.219 [0.131]	57.162 [0.000]*		1429.3 [0.968]
Slovakia	**hicp**	1.695 [0.149]	25.243 [0.000]*	0.063 [0.999]	63.181 [0.939]
	ppi	1.561 [0.187]	12.409 [0.002]*	0.149 [0.993]	89.705 [0.263]
	oil	0.979 [0.465]	0.358 [0.836]	0.112 [0.997]	83.195 [0.442]
	e	1.939 [0.099]	6.035 [0.049]**	0.164 [0.990]	74.708 [0.704]
	y	1.766 [0.133]	3.321 [0.190]	0.129 [0.995]	92.684 [0.197]
	System	1.994 [0.063]	46.549 [0.000]*		1212.3 [0.635]

(Note: * represents statistical significance at the 1% level, while ** represents the 5% level).

Table A.5 Skewness and Excess Kurtosis of Systems with signs of Non-normality

Country	Variables	Skewness	Excess Kurtosis
Hungary	hicp	0.454	5.226
	ppi	0.589	5.821
	oil	-0.434	3.104
	e	-0.086	3.059
	y	-0.148	3.757
Lithuania	hicp	0.061	3.073
	ppi	-0.082	3.407
	oil	-0.462	3.095
	e	-0.107	3.016
	y	-0.143	3.155
Poland	hicp	0.638	5.156
	ppi	0.287	4.171
	oil	-0.049	2.821
	e	0.008	3.731
	y	-0.288	3.577
Romania	hicp	0.531	4.053
	ppi	0.083	3.453
	oil	0.259	4.120
	e	-0.179	4.057
	y	-0.153	5.404
Slovakia	hicp	0.759	6.330
	ppi	0.202	4.701
	oil	-0.077	3.155
	e	-0.006	3.608
	y	0.136	3.543

Table A.6 LR Trace Test Results

H_0: rank=p	λ_{trace}	p-value
Bulgaria		
0	98.75	[0.01]*
1	25.45	[0.15]
Czech Republic		
0	153.75	[0.00]*
1	100.70	[0.00]*
2	58.69	[0.00]*
3	21.39	[0.34]
Estonia		
0	179.38	[0.00]*
1	115.41	[0.00]*
2	65.59	[0.00]*
3	21.27	[0.35]
Hungary		
0	104.18	[0.00]*
1	56.94	[0.01]*
2	25.69	[0.14]
Latvia		
0	120.01	[0.00]*
1	68.38	[0.02]**
2	43.02	[0.13]
Lithuania		
0	91.33	[0.00]*
1	40.53	[0.21]
Poland		
0	189.79	[0.00]*
1	111.11	[0.00]*
2	65.56	[0.00]*
3	25.57	[0.15]
Romania		
0	241.31	[0.00]*
1	129.87	[0.00]*
2	73.12	[0.00]*
3	19.23	[0.49]
Slovakia		
0	181.83	[0.00]*
1	106.51	[0.00]*
2	58.42	[0.00]*
3	19.42	[0.47]

Notes: * denotes rejection of the null hypothesis of no cointegration at the 1% level; ** denotes rejection at the 5% level.

Table A.7 Normalisation Test for Domestic Consumer Prices

Country	LR Test: Zero Restriction on HICP in Long-Run Matrix
Bulgaria	14.979 [0.001]*
Czech Republic	7.584 [0.055]***
Estonia	13.852 [0.003]*
Hungary	8.964 [0.011]**
Latvia	6.123 [0.047]**
Lithuania	7.786 [0.099]***
Poland	23.142 [0.000]*
Romania	22.014 [0.000]*
Slovakia	33.502 [0.000]*

Note: *, **, and *** denotes rejection of the null hypothesis that the zero restriction holds at the 1%, 5%, and 10% levels respectively.

Table A.8 LR Matrix with Alternative Normalisation: Coefficients on First Cointegrating Vector

Country	CV normalised on Domestic Total PPI					
	ppi	hicp	oil	e	y	t
Bulgaria	1.000	0.167 (0.163)	0.138* (0.021)	-0.780* (0.238)	0.002* (0.001)	0.002* (0.000)
Czech Republic	1.000	0.185 (0.341)	0.159* (0.024)	-0.848* (0.183)	0.109 (0.119)	0.000 (0.000)
Estonia	1.000	0.198 (0.192)	0.067* (0.016)	-0.927* (0.126)	0.552* (0.308)	0.003* (0.001)
Hungary	1.000	0.316* (0.046)	0.060* (0.019)	-0.765* (0.085)	-0.128* (0.058)	0.001* (0.000)
Latvia	1.000	0.277* (0.118)	1.132* (0.018)	-1.003* (0.102)	0.184 (0.124)	0.001* (0.000)
Lithuania	1.000	0.421* (0.266)	0.602* (0.030)	-0.575* (0.178)	0.366 (0.218)	0.000 (0.001)
Poland	1.000	0.169 (0.096)	0.028 (0.026)	-0.600* (0.105)	-0.001* (0.000)	0.001* (0.000)
Romania	1.000	0.516* (0.109)	0.181* (0.035)	-0.529* (0.088)	-0.060 (0.153)	-0.003* (0.001)
Slovakia	1.000	0.501* (0.120)	0.137* (0.018)	-0.532* (0.086)	0.280 (0.231)	0.000 (0.000)

(Note: Standard errors in parentheses; * denotes significance at the 5% level or below)

Table A.9 Normalisation Test for Producer Prices

Country	LR Test: Zero Restriction on PPI in Long-Run Matrix
Bulgaria	12.016 [0.003]*
Czech Republic	10.967 [0.012]**
Estonia	19.459 [0.000]*
Hungary	12.725 [0.002]*
Latvia	6.953 [0.031]**
Lithuania	12.871 [0.012]**
Poland	31.221 [0.000]*
Romania	17.371 [0.001]*
Slovakia	45.072 [0.000]*

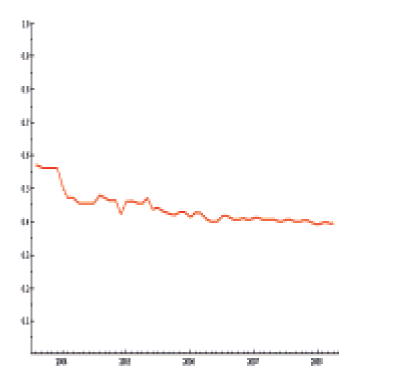
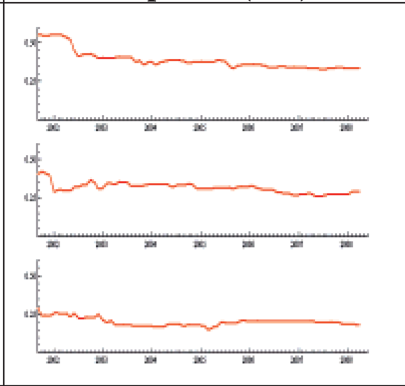
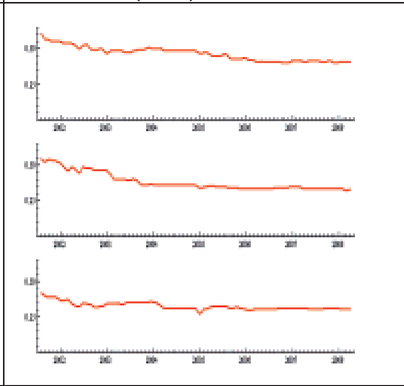
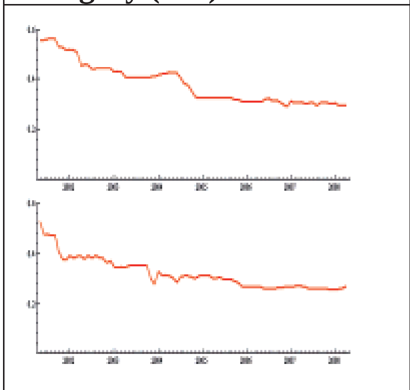
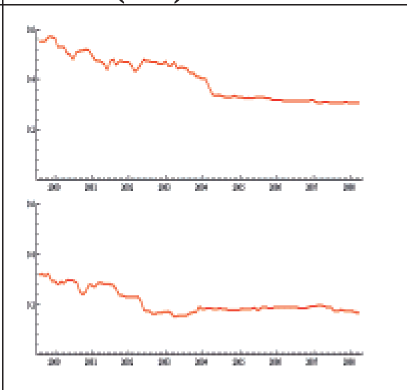
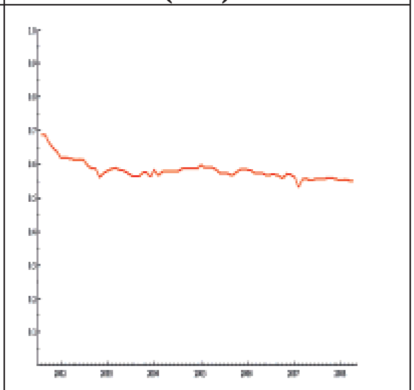
Note: *, **, and *** denotes rejection of the null hypothesis that the zero restriction holds at the 1%, 5%, and 10% levels respectively.

Table A.10 **Alternative Variable Ordering for Impulse Response Function Analysis**

Country	Response of HICP to 1% NEER Shock			
	6 months	*12 months*	*24 months*	*48 months*
Bulgaria	0.114	0.200	0.423	0.468
Czech Republic	0.290	0.444	0.519	0.542
Estonia	0.011	0.123	0.551	0.640
Hungary	0.046	0.105	0.216	0.418
Latvia	0.304	0.391	0.541	0.656
Lithuania	0.047	0.164	0.477	0.544
Poland	0.268	0.312	0.353	0.492
Romania	0.221	0.228	0.359	0.469
Slovakia	0.059	0.144	0.246	0.454

(Note: The variable ordering in this case is as follows: *OIL* → *IND PROD* → *NEER* → *PPI* → *HICP*).

Figure A.2 Recursive Analysis of Eigenvalues, 36 month window

Bulgaria (r=1)	Czech Republic (r=3)	Estonia (r=3)
		
Hungary (r=2)	**Latvia (r=2)**	**Lithuania (r=1)**
		
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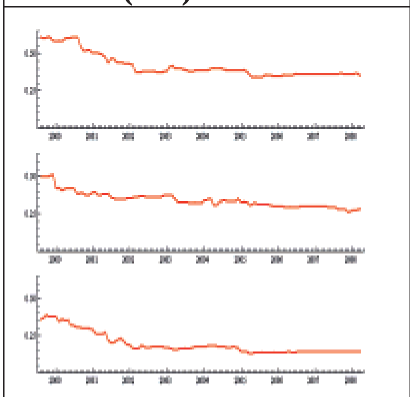	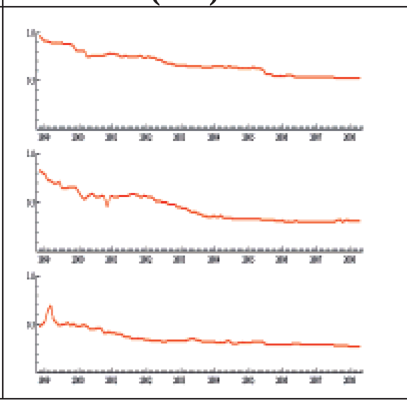	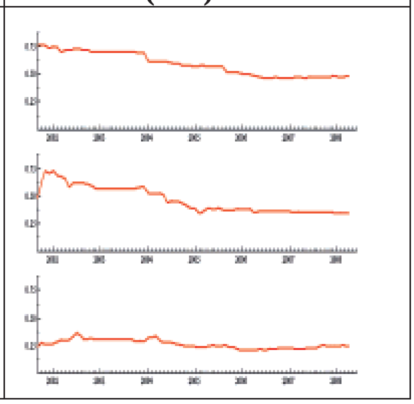

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