Identification of Exchange Rate Pass-Through Effect in Cointegrated VAR: an Application to New EU Member Countries

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Abstract

The exchange rate pass-through is of considerable importance for policymakers in open economies. Based on work of Johansen (2002) this paper discusses the conditions for identification of equilibrium pass-through effect in cointegration framework and shows that this leads to important economic implications. An empirical application to four new EU member countries: the Czech Republic, Hungary, Poland and Slovenia is also provided. The results show that the equilibrium pass-through effect can be actually identified and is generally very high. Moreover, we can clearly rank the countries under analysis according to the size of pass-through effect in relation to the degree of accommodation of monetary policy. The benefits of exchange rate flexibility are consequently rather low, thus providing an argument against delaying entry to the Euro area in new EU countries.

Povzetek

Prehajanje deviznega tečaja v cene je velikega pomena za ekonomsko politiko v odprtih gospodarstvih. Ta članek analizira pogoje za identifikacijo ravnovesnega učinka prehajanja v kointegriranem VAR modelu in pokaže, da ima to pomembne ekonomske implikacije. Izvedena je tudi empirična analiza štirih novih članic EU: Češke, Madžarske, Poljske in Slovenije. Rezultati kažejo, da je učinek prehajanja mogoče identificirati in je v splošnem zelo visok. Poleg tega lahko države jasno rangiramo glede na učinek prehajanja v povezavi s stopnjo akomodativnosti denarne politike. Posledično so koristi fleksibilnega deviznega tečaja relativno nizke, kar predstavlja argument v prid hitremu vstopu v EMU.

JEL codes: E42, E52, E58, C32

Keywords: EMU accession, pass-through effect, identification, cointegration analysis, policy accommodation

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1 Introduction

Exchange rate pass-through is defined as the change in prices caused by the change in the nominal exchange rate. The pass-through effect operates broadly through three basic channels: (1) a direct effect through prices of imported goods in the CPI; (2) an effect through prices of imported intermediate goods; and (3) an effect through price setting and expectations that include also the expected responses of monetary policy (Garcia and Restrepo, 2001). Estimation of pass-through effect leads to important policy implications in the choice of exchange rate regimes. At the same, it represents an important challenge in empirical work.

The presence of a variety of channels affecting the pass-through makes its empirical estimation difficult. Several recent studies have estimated the size of pass-through in large set of cross-country data. The present study differs substantially from previous literature in the estimation methodology employed. I use the cointegrated vector autoregression model that is able to capture complex dynamics in the data and at the same time account for long-run equilibrium relations. In addition, I invoke theoretical results from Johansen (2002) to address the issue of identification of pass-through effect, which has not been achieved in previous studies using cointegration analysis (e.g. Kim, 1998). Thus, this is not the first paper to use cointegration analysis to estimate exchange rate pass-through, but the first to solve the identification problem within this framework. Because the problem of identification arises any type of estimation that does not a priori rule out a stable long-run relationship between nominal exchange rate and prices, I argue that previous analyses did not properly measure the coefficient of pass-through. With some exceptions, the literature generally finds low estimates of pass-through effect. The estimates presented in this paper are in general higher. The difference may arise dues to lack of identification and focus on transitory changes in the nominal exchange rate only.

In the empirical application this paper focuses on four countries that have entered the European Union in 2004: the Czech Republic, Hungary, Poland and Slovenia (CEEC-4 hereafter). They will eventually adopt the euro as they have no opt-out clause. Therefore, the main open question about exchange rate policy for new members is the speed of entry into the Euro area. The chosen strategy crucially depends on the role policy-makers attribute to the exchange rate for macroeconomic performance. This, in turn, depends on the impact that nominal exchange rate fluctuations have on domestic inflation, in other words, on the magnitude of the pass-through effect. Potential advantages of delaying entry in the Eurozone declines if movements in the nominal exchange rate mostly affect domestic inflation rather than the real exchange rate.

The paper concentrates on CEEC-4 for the following reasons. The main reason is that from a structural point of view these countries form a more homogeneous set than other emerging markets. They are all closely linked to the European Union and are small open economies with a highly diversified trade and industrial structures. Finally, from a policy point of view they are all engaged in a process that will eventually lead to adopting the Euro.

The empirical analysis indicates that pass-through to CPI inflation is high in the four new EU members examined, although important differences emerge, which can be associated with differences in exchange rate regimes. This fact leads to a clear policy implications: in any disinflation experiment the central role should be given to the path of the nominal exchange rate. Even abstracting from the issue of propagation of exogenous shocks originating in international financial markets, flexible exchange rates are not an effective instrument for absorbing asymmetric real shocks (Masten, 2002). Large pass-through to import price is an incentive for policy-makers to attempt ex-post to drive the exchange rate in a way that improves external competitiveness. Large pass-through to CPI, however, reduces the expenditure switching effect and benefits of exchange rate flexibility. A policy of real exchange rate targeting creates persistent inflationary pressures that can be broken down by credibly adopting a non-accommodating exchange rate policy. Luckily, candidate countries have the point of arrival, the euro, already set. In answer to the question about the speed of entry to the Euro area posed at the beginning this paper suggest that there are no significant advantages to delaying entry.

The paper proceeds as follows. Section 2 discusse the choices of exchange rate regimes and inflation and exchange rate dynamics in CEE-4. Section 3 contains the main empirical analysis of the paper, focusing on the identification of pass-through effect in a cointegrated VAR. Section 4 concludes.

2 Inflation and Exchange Rate Regimes in CEEC-4

CEEC-4 started the transition process with two-digit inflation levels that gradually declined. The gradual decline in inflation has been accompanied by a sizable appreciation of the real exchange rate in all CEEC-4, and indeed in all transition economies.[1] A component of this trend appreciation can be considered an equilibrium phenomenon, in line with the Balassa-Samuelson effect and relative price convergence (on the latter see Čihak and Holub, 2002). Reduction to single-digit inflation was much faster in Slovenia and the Czech Republic, countries less affected by large stocks of debt and the attendant need to finance large debt service payments. However, inflation rates seem to be more stubborn in Slovenia and Hungary than in the Czech Republic and Poland. In the period 1998-2002, inflation hovered around 8-9 % in Slovenia and Hungary, with signs of decline after the second half of 2002, in a period of economic slowdown. The sharp decline in inflation in the Czech Republic and Poland perhaps reflects an overshooting of the inflation target. This was the result of an excessively tight monetary policy that negatively affected the economy during a period of general economic slowdown in Europe. The most puzzling was inflation behavior in Slovenia. Although Slovenia apparently had the best fundamentals of CEEC-4 and intially had also a favorably low inflation, it was the worst inflation performer towards the end of the period examined (see Figure 1). As shown below important reasons for such developments lie in differences in exchange rate regimes and degrees of policy accommodation.

During the period examined (1993-2002) CEEC-4 displayed a high degree of heterogeneity in their exchange rate policies. Poland and the Czech Republic gradually moved to floating regimes with inflation targeting, while Slovenia and Hungary operated in managed float systems, where the exchange rate was in principle flexible, but heaviliy controlled.[2] From a theoretical point of view, it is easy to establish a connection between exchange rate policy and the size of pass-through. For instance, in a simple New Keynesian framework in which firms in the non-tradable sector are price setters and internalize the government policy rule, exchange rate policies based on a predictable path of the exchange rate are likely to imply high pass-through. Empirically, such connection has been suggested by Calvo and Reinhart (2002) that argue that one of the main reasons for the much higher size of pass-through in emerging markets is due to their different

[1] As seen from Figure 2, average real exchange rate growth was negative, thus implying real appreciation.

[2] Poland adopted a flexible exchange rate system with inflation targeting in 2000, while Hungary switched to an inflation targeting regime in October 2001. It is questinable, however, whether de iure inflation targeting regime in Hungary is operational also de facto.

exchange rate policy i.e. managed floating and real exchange rate targeting. In such regimes exchange rate movements tend to be more persistent and predictable. That higher persistence of shocks implies higher pass-through of costs to prices was emphasized by Taylor (2000). Along simililar line of reasoning Coricelli et al. (2004) propose a simple theoretical framework to show that real exchange rate targeting results in perfect exchange rate pass-through.

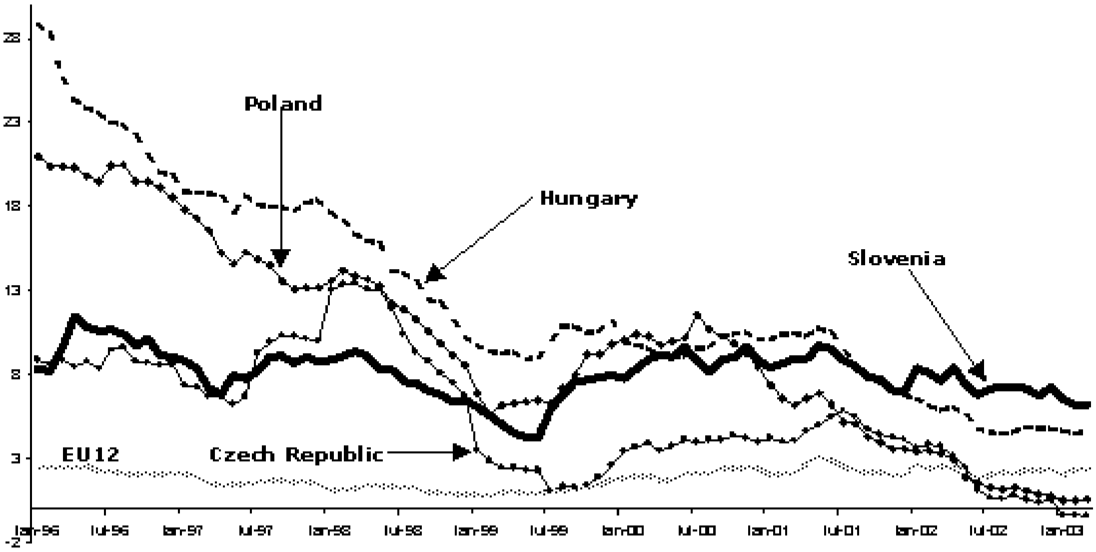

Figure 1: Inflation rates in CEEC-4

In advanced market economies, pure floating regimes tend to dominate. In such a case, fluctuations of the nominal exchange rate absorb the various shocks hitting the economy, while monetary policy is driven by the objective of achieving an inflation target. It thus follows that fluctuations in nominal exchange rate need not have a large effect on domestic inflation. By contrast, in many emerging markets exchange rate policy is accommodative, i.e. tries to neutralize the effects of adverse shocks on the real exchange rate. When taking the form of a reaction function to perceived disequilibria in the real exchange rate, a systematic component is induced into the dynamics of the nominal exchange rate. Such a policy is likely to be incorporated into the pricing decisions of economic actors. For this reason, a strong correlation between exchange rate movements and inflation rates can be observed in managed float regimes. By contrast, episodic changes of the exchange rate, even if large, tend to have little effect on pricing decisions and thus on inflation. Moreover, a showed by Uribe (2003), a real exchange rate rule results in indeterminacy, which implies also inability to control inflation.

A conjecture that different exchange rate regimes had a fundamental impact on domestic inflation can be very relevant for the case of CEEC-4. In particular, it should not be surprising to find a high or even perfect pass-through from exchange rate growth to domestic inflation in Slovenia and Hungary and a smaller impact in the Czech Republic and Poland. Even though it has never been officially declared, Slovenia apparently targeted the real exchange rate throughout the period, trying to maintain external competitiveness. The real exchange rate rule in

Slovenia was probably internalized by price setters, thus becoming a persistent source of inflation and a plausible explanation for the unfavorable behavior of inflation documented in Figure 1. This reasoning is confirmed also in Figure 2. Consistenly with our prior about low degree of policy accommodation, Czech data reveal the smallest comovements between inflation and nominal depreciation rate. This is also clearly the case for Poland after 1999, when they moved away from real exchange rate targeting. Consistently with lower correlation between inflation and nominal depreciation, nominal depreciation more strongly transfers into changes in the real exchange rate. The situation is different in Hungary and especially Slovenia. In both countries we can visually observe a higher correlation between nominal depreciation and inflation.[3] Inspection of average nominal depreciation rates in CEEC-4 also offers some interesting insights. It is especially notable for the Czech Republic that its nominal exchange rate growth never deviated permanently from zero and it was on average the by far lowest, which considerably contributed to its most favorable inflationary performance in the CEEC-4 group. While Poland and Hungary stopped maintaining positive nominal depreciation rates in 2000 and second half of 2001 respectively, Slovenia maintained a positive depreciation rate, which is entirely policy induced and a sign of strongest policy accommodation. Furthermore, the two countries that used more actively the exchange rate as a policy tool, Hungary and Slovenia, displayed a smaller volatility in the real exchange rate, suggesting the presence of some form of real exchange rate targeting and thus a higher degree of policy accommodation.[4]

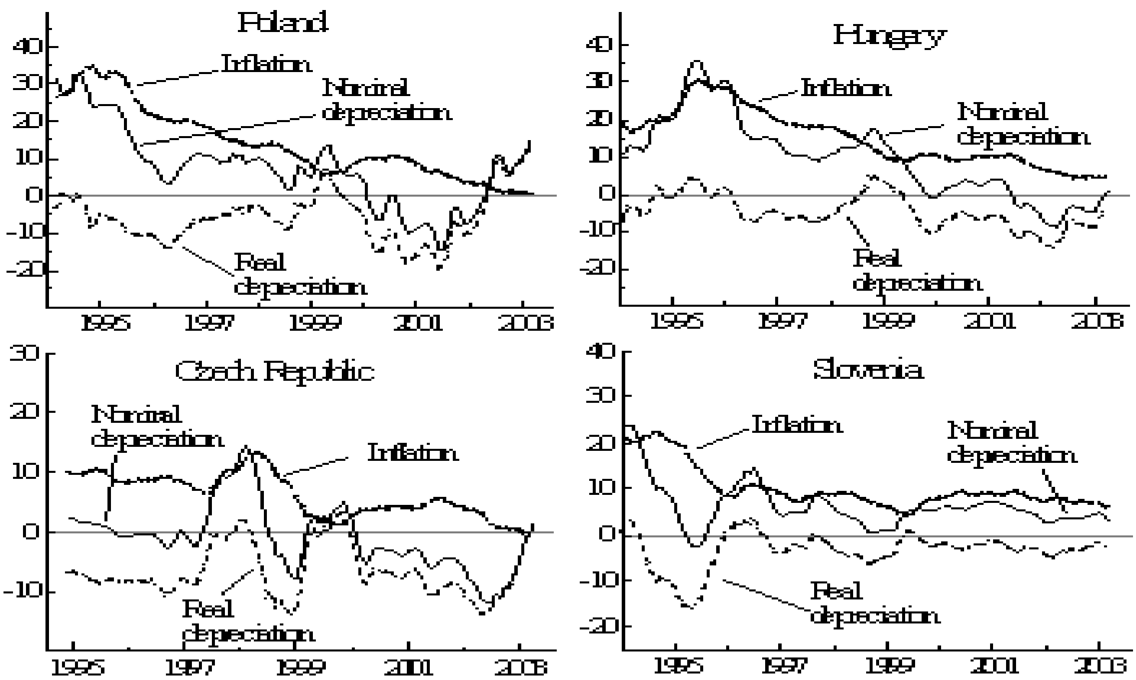

Figure 2: Inflation and nominal and real depreciation (3-month moving averages)

[3] I do not report simple correlation coefficients between inflation and nominal depreciation because both variables result to be nonstationary in CEEC-4 and correlation could be spurious.

[4] Standard deviations of real exchange rate changes during the period 1994 - 2002 in the Czech Republic, Poland, Hungary and Slovenia are 6.6, 6.1, 4.9 and 4.5 respectively.

In sum we can say that it is important to take into account whether exchange rate management is used as a key tool for achieving a preset inflation target, or to accommodate shocks to the real exchange rate. As argued above, while in the former the exchange rate does not add to inflationary pressures, it does so in the latter. Even though CEEC-4 operated in a de iure flexible exchange rate regimes it is in Hungary and Slovenia where exchange rate flexibility contributed to inflationary pressure to a considerably higher extent. Even though the country sample is small, there is evidence consistent with a strong relationship between exchange rate policy and regime and the size of pass-through. Following this line of reasoning, the estimates of exchange rate pass-through in CEEC-4 suggest that regimes with a more accommodative stance of exchange rate policy generate a higher pass-through effect. Although accommodative exchange rate rules could stabilize the real exchange rate, it is questionable whether such a policy is welfare improving as it generates costs associated with higher average inflation.

3 Estimation and Identification of Pass-Through Effect

This section presents the empirical analysis of pass-through effect of exchange rate growth to CPI inflation within a cointegration framework. Compared to existing studies of pass-through the analysis introduces two novelties. First, within I(1) cointegration analysis the paper offers a formal discussion of identification of pass-through effect conditional on cointegration rank. It is shown that reported estimates of equilibrium pass-through effect are actually identified and can thus be interpreted as elasticities in structural sense. The issue of identification of pass-through effect has not yet been addressed in the existing literature and as is turns out this is of crucial importance when conclusions from empirical analysis are used for policy implications. Second, I discussed the necessary contemporaneous correlation of short-run changes that yields the empirically observed equilibrium pass-through effect. All these features, combined with actual estimates, lead to important implications for the choice of exchange rate regime.

The choice of variable for the empirical analysis has been motivated by the following reasoning. First, besides the nominal exchange rate (differenced) and inflation, a measure of real output is inserted for two reasons, both stemming from a general perception that in the long run prices and nominal exchange rates move together, except for long-run changes in real exchange rate. The first reason follows from strongly present tendecy of real exchange rates in CEEC-4 to appreciate (see Figure 2). This trend appreciation can be associated to the productivity improvement due to the Balassa-Samuelson effect and adjustment of relative prices. Without controlling for output dynamics the estimate of pass-through would necessarily be larger than unity, which is also what Campa and Goldberg (2002) report for some countries. The second reason for including output follows from noting that in a properly defined inter-temporal model, nominal exchange rate movements may affect the equilibrium real exchange rate if this is affected by the rate of inflation. Indeed, it is conceivable that inflation may affect growth and the level of output in economies with severe imperfections in financial markets. The second additional variable considered in the empirical model is interest rate spread because in economies open to capital movements, like the CEEC-4, exchange rate changes affect the economy via the interest rate parity. In summary, the empirical application considers the system (4.1) in which the effects of changes in nominal exchange rates on inflation are analyzed taking into account the endogenous dynamics in both output and interest rate differentials:

$$X_t = (y_t, \Delta e_t, \pi_t - \pi_t^*, i_t - i_t^*) \tag{4.1}$$

y_t, denotes index of total industrial production, Δe_t the growth of nominal exchange rate, $\pi_t - \pi_t^*$ the inflation differential with respect to Germany, and $i_t - i_t^*$ the nominal interest rate differential with respect to 3-month Euribor.[5] Data are monthly and cover the period 1993 - 2002 (exact estimation samples are reported in Table 1). Transformations of variables (levels or differences) follow from I(2) cointegration analysis in Coricelli, Jazbec and Masten (2003). Domestic and foreign CPI inflation rate enter as a homogeneous relation because the relation between nominal exchange rate growth and inflation differential is what we are primarily interested in. The coefficient to the inflation rate differential can be directly related to the pass-through effect of nominal exchange rate changes to domestic inflation. In addition imposing homogeneity restriction reduces the dimension of the system, which enables a valid statistical inference with the data sample available for CEECs. This was also the reason to include the nominal interest rates as a spread. Producer price inflation is not included because the equilibrium pass-through effect cannot be identified in such a case (see corollary to Proposition 1 below).[6]

In the estimation I use cointegrated VAR analysis. Other authors used different approaches. Campa and Goldberg (2002) who estimate a simple single-equation model for 25 OECD countries over the period 1975 to 1999 and measure the pass-through effect (to import prices in their case) with the coefficient on the nominal exchange rate. Goldfajn and Werlang (2000) study the relationship between exchange rate depreciations and inflation for 71 countries in the period 1980 to 1998. Choudhri and Hakura (2001) extend the study of Goldfajn and Werlang (2000) and try to establish the role of the exchange rate regime in determining the extent of pass-through in 71 countries in the period 1979 to 2000. Single equation approach is used also by Darvas (2001) for the group of Acceding Countries.

The second methodological approach is the use of structural VAR without cointegration analysis, of which McCarthy (2000) is a very notable example. It is a systems approach that captures consistently the endogenous dynamic relation between the nominal exchange rate, prices and other conditioning variables and is thus preferable to single equation estimation. In a structural VAR pass-through is measured by means of impulse responses of different price series to an exchange rate shock.[7] The problem with this approach is that it gives only a partial estimate of pass-through. As shown by Corsetti and Dedola (2003) we can observe as many measures of (short-run) pass-through as there are identified structural shocks. Concentrating on price responses after only exogenous exchange rate shocks neglects other sources of stochastic variation in the nominal exchange rate and prices. In addition, exchange rate changes need not occur only as consequences of stochastic shocks, but they can also reflect systematic changes in policy like the change in inflation target. All such changes are not accounted for in a typical SVAR

[5] 3-month Fibor prior to 1999.

[6] It follows from the discussion of identification below that inclusion of two inflation rates whose price components overlap in the same cointegration relation would imply a zero equilibrium pass-through effect of permanent changes in depreciation rate into one of the two inflation rates. This is what happens in the analysis of Billmeier and Bonato (2002). In such a situation the equilibrium pass-through effect is clearly not identified. By including only the CPI inflation rate I avoid this problem.

[7] A different use of SVAR analysis is found in Choudri, Faruqee and Hakura (2002). Their empirically observed impulse responses of prices to an exchange rate shock are used not to measure pass-though directly but as a benchmark for simulated responses obtained from calibrated theorectical model under differen assumptions about nominal rigidities in the economy.

analysis. Estimates of exchange rate pass-through could in such a case be severely biased and underestimated. Present analysis accounts for all sources of variation in nominal depreciation rate.

Estimation of pass-through effect with cointegrated vector autoregression model (CVAR) has a number of important advantages. Price series are commonly integrated at least of order one, which calls for an explicit test for cointegration. From an economic point of view, neglecting cointegration is surprising since long-run co-movement of prices and exchange rate is borne out by theory. Neglecting cointegration when it is genuinely present means neglecting the intrinsic meaning of equilibrium long-run relationship between the nominal exchange rate and prices. In order to maintain comparability with typical SVAR-based studies this paper at the end of the analysis estimates also a structural vector error-correction (VEC) model (see Warne, 1993 for a detailed discussion).

CVAR has been used in previous studies (Kim, 1998; Billmeier and Bonato, 2002). As shown below, however, the estimates of the pass-through presented in these two studies are not identified. Lack of identification means that coefficient estimated cannot be interpreted as coefficients of pass-through.

The second important advantage of estimation with a CVAR is that it enables us to distinguish between permanent and transitory shocks. This is a fundamental distinction, since only permanent exchange rate shocks have a non-zero long-run equilibrium pass-through effect. Indeed, it is unlikely that transitory exchange rate shocks induce significant short-run changes in pricing behavior if firms face costs associated with frequent price changes. If the analysis is to be used for policy implications about disinflation policies and the choice of exchange rate regime, tracing the effects of permanent shocks separately becomes even more important. As briefly discussed in section 2, one of the main sources of persistence in exchange rate changes has to do with the presence of real exchange rate targets. Transitory changes in the depreciation rate can be seen as facilitating relative price adjustment in presence of nominal rigidities and thus welfare improving. Permanent changes, on the contrary, go beyond relative price adjustment and can be understood as a consequence of policy accommodation and thus it should not be surprising that pass through perfectly into inflation. Because the interplay between policy accommodation and the size of pass-through effect is of central importance in the paper, I concentrate on the analysis of permanent changes the depreciation rate and inflation and expect to find higher pass-through coefficients in countries following accommodative exchange rate rules.

3.1 Identification of Pass-Through Effect

In the empirical applicatio the following I(1) system is estimated

$$\Delta X_t = \Pi X_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \Phi D_t + \varepsilon_t \tag{4.2}$$

with a corresponding reduced rank condition $\Pi = \alpha\beta'$ (see Johansen, 1995a for a detailed presentation). The matrix β contains the cointegrating relations and α contains the corresponding loading coefficients. Matrices C and Γ are defined as

$$C = \beta_\perp \left(\alpha'_\perp \Gamma \beta_\perp\right)^{-1} \alpha'_\perp \text{ and } \Gamma = I - \sum_{i=1}^{k-1} \Gamma_i$$

where $\alpha_\perp$ and $\beta_\perp$ are the orthogonal complements to α and β respectively. The cointegration coefficient between depreciation rate and inflation spread is of central interest because we need establish under what conditions it can be interpreted as equilibrium pass-through effect. A cointegration relation containing Δe_t and $\pi_t - \pi_t^*$ can be generically written in regression format as follows

$$\pi_t - \pi_t^* = \lambda_1 \Delta e_t + \lambda_2 \left(i_t - i_t^*\right) + \lambda_3 y_t \tag{4.3}$$

Note that with cointegration rank equal to two we have either $\lambda_2 = 0$ or $\lambda_3 = 0$; with rank three we have $\lambda_2 = \lambda_3 = 0$. In any case we need to check whether λ_1 can be interepreted as equilibrium pass-through effect. In other words, we need to check whether, based on an estimate of the cointegration relation of the type (4.3), we can say that with a long-run change in Δe_t by 1 percentage point inflation spread changes by λ_1 percentage points in equilibrium. This is a question about identification of pass-through effect.

To establish the conditions for identification I invoke results of Johansen (2002). As it turns out identification depends on cointegration rank. From the solution of the error-correction model it follows that the long-run value $X_{\infty/t}$ as a function of current values $(X_t, X_{t-1}, ..., X_{t-k+1})$ is given by

$$X_{\infty/t} = \lim_{h\to\infty} E\left(X_{t+h} \mid X_t, X_{t-1}, ..., X_{t-k+1}\right) = C\left(X_t - \sum_{i=1}^{k-1} \Gamma_i X_{t-i}\right) \tag{4.4}$$

Thus, it follows from the definition of matric C that the long-run changes in endogenous variables are proportional to $\beta_\perp$. A given long-run change $k \in sp\left(\beta_\perp\right)$ can be achieved by either adding k to all current values or by adding to Γk to X_t - a short-run change. It is also assumed in what follows that cointegrating vectors are identified using zero restriction, which justifies the interpretations used below (see Proposition 2 in Johansen, 2002). The following proposition gives a sufficient and necessary condition for identification of the pass-through effect.

Proposition 1 *Equilibrium pass-through effect is identified if and only if the cointegrating rank r is equal to 1 plus the number of variables other than inflation with a non-zero coefficient in $k \in sp\left(\beta_\perp\right)$.*

Proof. X_t is a p dimensional vector of variables. Without loss of generality assume $\pi_t - \pi_t^*$ is placed first and Δe_t second, all other $p-2$ variables are below. Consider a long-run change $k = \left(\mu, 1, \widetilde{k}, \mathbf{0}_{1x(p-2-n)}\right)$, that is, a long-run change in depreciation rate by one percentage point accompanied by a long-run change in inflation differential by μ percentage points, while allowing for a non-zero effect on n variables in sub-vector $\widetilde{k}$. μ measures the equilibrium pass-through effect and is a parameter that needs to be uniquely identified. Note that $k \in sp\left(\beta_\perp\right)$, hence the parameters in k must solve $k'\beta = 0$. β is rxp and it must be identified using zero restriction (see Johansen, 2002). k has $n+1$ unknown parameters. $k'\beta = 0$ is therefore a system of r linear equations and $n+1$ unknowns. It has a unique solution when $r = 1+n$. In such a case

μ is uniquely identified. Unless we have some prior statistically non-testable information for parameters of $\widetilde{k}$ this is also the only case when it can be identified. ■

It directly follows from Proposition 1 that identification of pass-through effect implies also that a long-run equilibrium change in the depreciation rate has non-zero equilibrium effect on $n = r - 1$ variables in X_t. This leads to the following corollary to Proposition 1.

Corollary 2 *When pass-through effect is identified in a p-dimensional system, permanent exchange rate changes are associated with a non-zero equilibrium changes in exactly r-1 variables other than inflation.*

Proposition 1 and its corollary have in our case important economic implications. With rank two the pass-through effect can be identified and we also observe that a permanent change in depreciation rate is associated with a zero change in real output. In case of rank three we also achieve identification, but in such a case a permanent change in depreciation rate also permanently affects real output. Note that economic implication of these two cases are different. If the relation turns out to be negative, as it is the case in the empirical application of this paper, this implies that a disinflation process through lower depreciation rate results in a higher level of output. This has much stronger implications for the benefits of the disinflation process.

For a better illustration of identification issue let us consider the following example:

Example 3 *Consider the following system of variables: $X_t = (\pi_t - \pi_t^*, \Delta e_t, i_t - i_t^*, y_t)$. In our four-dimensional system genuine cointegration is found for ranks 1, 2 and 3. With rank 1 a single cointegrating relation can be conveniently written in regression format as*

$$\pi_t - \pi_t^* = \lambda_1 \Delta e_t + \lambda_2 (i_t - i_t^*) + \lambda_3 y_t$$

which corresponds to cointegrating vector $\beta = (1, -\lambda_1, -\lambda_2, -\lambda_3)'$. As discussed above we need to check whether λ_1 can be interpreted as a measure of pass-through effect. Consider a long-run change $k = (\lambda_1, 1, 0, 0)'$, that is, a long-run change in inflation differential by one percentage point accompanied by a long-run change in exchange rate growth by λ_1 percentage points, while leaving the interest rate spread and real output unchanged. Note that $k'\beta = 0$. Hence $k \in sp(\beta_\perp)$ and we could interpret λ_1 as the equilibrium pass-through effect. From economic point of view, however, this long-run change is inadmissible because it leads to a permanent deviation from the uncovered interest rate parity and could be observed only in a country with very strict capital controls. To interpret λ_1 as the pass-through effect we need to explore the feasibility of the long-run change of the form $k = (\lambda_1, 1, \mu, 0)'$ for $\mu > 0$.[8] Clearly such a vector is not orthogonal to β for any $\mu \neq 0$ and hence λ_1 cannot be interpreted as the pass-through effect. To obtain an estimate of pass-through effect we need to allow for long-run changes of the form $k = (\widetilde{\mu}, 1, \mu, 0)'$. The pass-through effect would in this case be

$$\widetilde{\mu} = \lambda_1 + \lambda_2 \mu$$

which without prior knowledge of μ is not identified, because μ cannot be estimated from the parameters of (4.2). Taking $\mu = 1$ arbitrarily is valid only if we can assume stationarity of the interest rate premium. Next consider the case with $r = 2$. Without loss of generality assume that

[8] $\mu \neq 1$ allows for a non-stationary interest rate risk premium.

cointegrating vectors are identified in such a way that the depreciation rate enter only the first relation. In regression format they can be written as

$$\begin{aligned}\pi_t - \pi_t^* &= \lambda_1 \Delta e_t + \lambda_2 (i_t - i_t^*) \\ \pi_t - \pi_t^* &= \lambda_3 (i_t - i_t^*) + \lambda_4 y_t\end{aligned}$$

with corresponding cointegrating vectors: $\beta_1 = (1, -\lambda_1, -\lambda_2, 0)'$ *and* $\beta_2 = (1, 0, -\lambda_3, -\lambda_4)'$. *Clearly a vector of type* $k = (\lambda_1, 1, 0, 0)'$ *is not orthogonal to* β*, which is also true for the vector* $k = (\lambda_1, 1, \mu, 0)'$. *Thus, again we need to consider* $k = (\widetilde{\mu}, 1, \mu, 0)'$, *for which it holds that* $k \in sp(\beta_\perp)$ *for* $\mu = \lambda_1 (\lambda_3 - \lambda_2)^{-1}$ *and it obtains that the equilibrium pass-through effect is*

$$\widetilde{\mu} = \frac{\lambda_1 \lambda_3}{\lambda_3 - \lambda_2}$$

and is in this case identified. However, it should be noted that it is not directly estimated by λ_1. *To find the estimate of equilibirum pass-through we need to solve a simple system of equations. Three cointegrating relations can be written as*

$$\pi_t - \pi_t^* = \lambda_1 \Delta e_t, \ \pi_t - \pi_t^* = \lambda_2 (i_t - i_t^*), \ i_t - i_t^* = \lambda_3 y_t$$

with corresponding cointegrating vectors: $\beta_1 = (1, -\lambda_1, 0, 0)'$, $\beta_2 = (1, 0, -\lambda_2, 0)'$ *and* $\beta_3 = (0, 0, 1, -\lambda_3)$. *The orthogonal complement to these vectors (normalized on the second element) is* $\beta_\perp = \left(\lambda_1, 1, \lambda_1 \lambda_2^{-1}, \lambda_1 \lambda_2^{-1} \lambda_3^{-1}\right)$, *a one-dimensional space. In this case the pass-through effect is identified and directly measured by* λ_1.

Thus, the identification of long-run or equilibrium pass-through effect depends on cointegration rank. In our case, we can identify it for ranks two and three. However, there is a fundamental difference between these two cases. With rank three pass through is measured directly by the cointegration coefficient between inflation and depreciation rate. With rank two a simple system of two equation with two unknowns needs to be solved in order to find a unique long-run adjustment for the interest rate spread. Therefore, identification of pass-through effect requires estimating a unique equilibrium relation between depreciation rate and inflation spread. The last statement leads to the following definition.

Definition 4 *Equilibrium pass-through effect is measured by the coefficient of nominal exchange rate growth on the difference between domestic and foreign CPI inflation in a cointegrating relationship that contains no other variables.*

In line with Proposition 1 identification of long-run or equilibrium pass-through effect depends on cointegrating rank. When identification can be achieved (cases with $r = 2$ or 3), we can also determine the corresponding contemporaneous pass-through effect – a short-run change – that supports a given long-run change. As seen above, a given long-run change $k \in sp(\beta_\perp)$ can be achieved by Γk adding to X_t. We can interpret this change also as the effects of shocks that clearly have permanent effects on variables in X_t. In fact, this is a restriction any type of shock in structural sense (real and nominal) that economic theory can justify to have a permanent change given by $k \in sp(\beta_\perp)$. In addition, short-run changes need not occur only due to exogenous shocks (ε_t in expression 4.2), but also as a consequence of systematic policy reactions and changes in policy targets.

The focus of the paper is on permanent effects of changes in exchange rate growth on inflation, because shocks that are only transitory (three in the present case) cause only transitory fluctuations around long-run equilibrium values and do not necessarily induce significant changes in pricing behavior of firms. Γk contains the contemporaneous effect of a permanent exchange rate change on inflation and other variables. This characterization of "shocks" is not based on any non-testable restriction imposed on reduced form parameters of the (cointegrated) VAR.

3.2 Implications of identification problem for other methodological approaches

The discussion of the identification problem has some implications both for single-equation and SVAR-based studies of pass-through effect. It is clear from the above discussion that a given single-equation estimate of

$$\pi_t - \pi_t^* = \lambda_1 \Delta e_t + \lambda_2 (i_t - i_t^*) + \lambda_3 y_t + \varepsilon_t$$

(also alowing for a dynamic spceification and conditional on being able to find valid instruments) would not enable us to interpret λ_1 as the measure of pass-through effect. A ceteris paribus interpretation of $\lambda_1$as the explicitly assumes that $i_t - i_t^*$ does not change as Δe_t changes. The same reasoning holds also for interpretation of λ_2. Abstracting from the case of complete capital immobility, in real world changes in Δe_t in $i_t - i_t^*$ occur simultaneously. Thus, we cannot use the single-equation estimates of λ_1 and λ_2 to determine the actual size of pass-through effect. The same problem would arise if an inflation rate of an alternative price index that shares common components with π_t would be added to the equation.[9] Also in this case λ_1 would not identify the true pass-through effect. This is also the reason why (4.4) contains only one inflation rate. In general, any specification of the empirical model for estimation of the pass-through effect that contains variables that are according to economic theory in equilibrium dynamically linked to the nominal exchage rate calls for systems estimation of the model, and within that model checking for the identification of the equilibrium effect. This reasoning goes beyond the estimation of only exchange rate pass-through. As follows from Johansen (2002), it is applicable to any type of empirical analysis where correct interpretation of cointegration coefficients is of importance.

The majority of studies using SVAR analysis of pass-through effect do not deal directly with cointegration. It should be noted, however, that reponses of prices and nominal exchange rate to any (structural) permanent shock should be linked by the equilibrium pass-through effect. For this reason is the identification of equilibrium pass-through effect of particular interest and can increase considerably the information content also of short-run pass-through analysis unless the purpose of the impulse responses analysis is limited to the analysis of the dynamic properties of the data (as in Choudri et al., 2002). The issue of identification would be irrelevant only if prices and nominal exchange rates were I(0) variables. Empirically such an assumption has been widely rejected. Thus, in empirically realistic cases identification issue of equilibrium effects becomes relevant. Note that VAR estimation on differenced data to achieve stationarity, without prior testing for cointegration rank implicitly assumes away the existence of equilibrium pass-through effect. If this is really present in the data, the estimates will be inconsistent. Estimation of a VAR on I(1) data in levels yields consistent (but inefficient) estimates; however, neglecting the analysis of cointegration relations neglects also important information about equilibrium

[9] In such a case there are also no valid instruments to obtain consitent single-equation estimates.

pass-through effect. It could result that testing for cointegration rank would signal rank zero of Π matrix in (4.2). However, this would imply the irrelevance of identification problem of the pass-through effect because of disablity to find a stable relation between the nominal exchange rate and prices. Theoretically this is highly unlikely and should be rather seen as a sign of misspecification of the empirical model.

3.3 Results of I(1) Analysis

Lag length of the systems has been chosen by complementary use of standard information criteria and the usual Wald-type tests for a valid lag reduction. It proved sufficient to include two endogenous lags for Slovenia, three for Hungary and the Czech Republic and four for Poland.[10] Tests for model misspecification are presented in Tables 3 and 4 in the appendix. We can conclude that final models do not suffer from misspecification. There are only some signs of non-normality of the residuals for Hungary and Poland, but it should be emphasized that the key assumption for the validity and robustness of cointegration analysis is that the residuals be stochastically independent and this is foremost confirmed by the absence of residual autocorrelation.[11]

The choice of cointegration rank is 3, uniformly across all four countries. Both the asymptotic and bootstrap versions of the trace test indicate this very clearly for Slovenia, Hungary and Poland, while rank 2 is also possible for the Czech Republic.[12] We have nevertheless chosen rank 3 also for the latter country as the systems show significant and strong equilibrium correction to the third cointegrating relation, and because we wanted to maintain direct comparability of results between all four countries. Visual inspection of the estimated third cointegrating vector presented in Figures 3 - 6 also show no obvious signs of non-stationarity for all countries.

The left panel of Table 1 presents the estimates of just-identified cointegrating vectors. The right panel reports the corresponding adjustment coefficients. Signs of all coefficients are consistent with economic theory. The most informative for the analysis of pass-through effects is the second cointegrating vector β_2 and, in particular, its first coefficient. As explained in previous section, it can be interpreted as long run or equilibrium pass-through effect. We can observe that it is the largest in Slovenia and practically identical to 1. For Hungary it is only marginally different, and for both countries the restriction that it is actually equal to 1 cannot be rejected, (see the corresponding likelihood ratio tests reported under Table 1 with corresponding p-values above 0.90). For Poland the point estimate of this coefficient is 0.80; however, we still cannot say that it is statistically significantly different from 1. The corresponding χ^2 test has a p-value of 0.55. The smallest is the point estimate of the coefficient for the Czech Republic, around 0.6. However, if we impose weak exogeneity of the industrial production index (statistically supported) and then test jointly the hypothesis that the pass-through coefficient is equal to 0.5, we cannot reject the restrictions (corresponding p-value is 0.09).

[10] As a robustness check also systems with 4, 6 and 8 lags have been estimated. It was encouraging to find that the choice of rank does not change with lag length. Moreover, lag length leaves the estimates of cointegrating space virtually unchanged. The variation in parameter estimates is so small that it does not change the conclusions presented in the paper.

[11] In each system outliers have been accounted for by including simple impulse dummies. The parameter stability of the VARs has been tested with recursive 1-Step and Break-Point Chow tests. The tests reveal no signs of parameter instability for all countries. The same turns out to be the case for the Hansen-Johansen test of constancy of cointegration coefficients. These test are available upon request.

[12] Bootstrap of the trace test has been performed in SVAR program written by Anders Warne.

To complete the exposition of I(1) analysis it is also important to look at the corresponding coefficients in Table 1, measuring the adjustment to the long-run relations. The most important finding is that output does not respond to deviations from the second and third cointegrating relations that represent open-economy parity relations. This strengthens the interpretation of coefficient β_{21} as the measure of pass-through effect because it implies that only nominal variables adjust to restore the equilibrium. In the interpretation of coefficient β_{21} this allows us to abstract from real shocks that can cause trend movements in the real exchange rate.

Table 1: Estimated Cointegration Relations and Loading Coefficients

	β_1	β_2	β_3	α_1	α_2	α_3
			Czech Republic[a] (1993:12 - 2002:7)			
y	1.00	-	-	0.00	0.00	0.00
Δe	-	**-0.64** (0.06)	-	**-62.06**	**-1.03**	**-1.28**
$\pi - \pi^*$	-	1.00	**-1.28** (0.19)	**-20.83**	0.02	**-0.79**
$i - i^*$	**0.03** (0.001)	-	1.00	-0.68	-0.00	0.02
			Hungary[b] (1993:2 - 2002:7)			
y	1.00	-	-	**-0.09**	0.00	0.00
Δe	-	**-0.97** (0.10)	-	-4.87	**-0.88**	**-0.98**
$\pi - \pi^*$	-	1.00	**-1.49** (0.11)	2.62	0.01	**-0.73**
$i - i^*$	**0.03** (0.004)	-	1.00	**-1.10**	0.00	0.01
			Poland[c] (1993:1 - 2002:4)			
y	1.00	-	-	-0.02	0.00	0.001
Δe	-	**-0.80** (0.06)	-	**-36.95**	**-0.96**	-0.72
$\pi - \pi^*$	-	1.00	**-0.84** (0.08)	**-27.35**	-0.00	**-0.80**
$i - i^*$	**0.03** (0.006)	-	1.00	-0.84	-0.00	0.01
			Slovenia[d] (1993:3 - 2002:3)			
y	1.00	-	-	**0.03**	-0.00	-0.00
Δe	-	**-1.01** (0.10)	-	**-26.49**	**-0.62**	**-0.78**
$\pi - \pi^*$	-	1.00	**-2.32** (0.20)	-2.66	0.02	**-0.75**
$i - i^*$	**0.01** (0.001)	-	1.00	**-18.51**	0.00	**0.38**

Notes: **Bold** indicates significance. Standard errors in parentheses.
[a] $H_0 : \beta_{21} = -1,\ \chi^2(1) = 2.52\ (0.11)$. Weak exogeneity of y: $\chi^2(3) = 3.03\ (0.39)$. Weak exogeneity of $y + H_0 : \beta_{21} = -1 : \chi^2(4) = 8.07\ (0.09)$
[b] $H_0 : \beta_{21} = -1,\ \chi^2(1) = 0.01\ (0.92)$.
[c] $H_0 : \beta_{21} = -1,\ \chi^2(1) = 0.36\ (0.55),\ H_0 : \beta_{32} = -1,\ \chi^2(1) = 0.76\ (0.38)$.
[d] $H_0 : \beta_{21} = -1,\ \chi^2(1) = 0.00\ (0.99)$.

It is surprising at first sight that only the exchange rate adjusts strongly and significantly to the second cointegrating relation. One would expect this also for the inflation rate differential. However, corollary to Proposition 1 (see also the left panel of Table 2 below) implies that any permanent increase in exchange rate depreciation that also increases inflation must be matched with a corresponding increase in the interest rate spread. We can observe from the vector of adjustment coefficients to the third cointegrating relation α_3 a very strong equilibrium adjustment of inflation (positively in response to a positive exchange rate shock that increases the interest rate spread).

3.4 Short-run and long-run changes

The orthogonal complement to β, denoted by $\beta_\perp$, is a 4x1 vector for each country. Estimates are reported in the left panel of Table 2. Admissible long-run co-movements of the variables analyzed are thus summarized by a one-dimensional space, which has qualitatively the same structure across countries. In the explanation lets consider the case of Slovenia: $\beta_\perp = (-0.025, 1, 1, 2.5)'$. This means that an equilibrium or permanent change in exchange rate growth accompanied by an equivalent increase in inflation differential, a disproportionate increase in interest rate spread, and consequently a lower level of output. The effect on the level of output is significant for all countries. Because the change in interest rate spread is more than proportional, ex-post real interest rate spread also increases, which is most likely the cause of lower output.[13] A disproportionate effect on the interest rate differential is a consequence of non-stationarity of the risk premium and the real interest rate spread. This non-stationarity is observed for all countries and is not surprising as these countries had gone through the process of transition to market economies. It is important to note that any other vector linearly independent of $(-0.025, 1, 1, 2.5)'$ will violate the orthogonality with respect to β and hence cannot span the equilibrium long-run changes in variables of interest. In particular, monetary authorities cannot permanently depreciate the currency without causing an equal change in the difference in CPI inflation rates. In other words, we must note that $(-0.025, 1, \lambda, 2.5)' \notin sp(\beta_\perp)$ for any $\lambda \neq 1$. This means that we can in line with Proposition 1 indeed interpret the coefficient as a measure of equilibrium pass-through into CPI inflation. Moreover, as noted above, any permanent rise in the rate of depreciation results in a rise in real interest rate differential and a negative effect on output.

Table 2: Orthogonal complement to cointegrating space initial changes

	$\beta'_\perp$				Γk			
	y	Δe	$\pi - \pi^*$	$i - i^*$	y	Δe	$\pi - \pi^*$	$i - i^*$
Czech Republic	-0.018	1	0.6	0.625	-0.038	7.87	-0.59	0.39
Hungary	-0.045	1	1	1.5	-0.092	1.00	1.05	1.01
Poland	-0.020	1	0.8	0.67	-0.047	0.33	0.63	0.49
Slovenia	-0.025	1	1	2.5	-0.019	0.84	0.84	1.58

Hungary again shows very similar properties to Slovenia, the only difference being that a policy of further exchange rate stabilization while yielding the same gain in inflation reduction yields a somewhat smaller, but still more than proportional, reduction in interest rate spread, and a larger positive effect on output. Poland and the Czech Republic share many similarities

[13] To see this, note that we can rewrite the third cointegrating relation generically denoted as $i - i^* = \beta_{31}(\pi - \pi^*)$ as $r - r* = (1 - \beta_{31})(i - i^*)$, where r denotes the ex-post real interest rate.

and appear distinct from Hungary and Slovenia. Their point estimate of pass-through is smaller than one, and consequently, there is also a smaller negative effect on output and a smaller required increase in interest rate spread to support a potential policy of accelerated exchange rate depreciation.[14] Again this fits with our priors about the nature of exchange rate policy in these two countries.

Right panel of Table 2 reports required initial (short-run) changes (see Johansen, 2002) in current values of endogenous variables Γk that yield a long-run response of the system spanned by $\beta_{\perp}$.[15] These results show that results on pass-through effect cannot be discussed independently of the specific type of the exchange rate regime in a certain country. It can be observed that in Hungary and Slovenia, two countries with most accommodative exchange rate regimes in the past, a "shock" that permanently increases the growth of the exchange rate is accompanied with practically equal change in inflation differential already on impact. The hypothesis that these contemporaneous changes can be induced by economic policy cannot be accepted based only on these results. However, it is also true that the hypothesis that this reflects a high degree of policy accommodation cannot be entirely rejected. It can be at least considered as an indicator of accommodative exchange rate policy that aims to stabilize short-run fluctuations in the real exchange rate. Very high contemporaneous correlation between changes in exchange rate growth and inflation can thus be seen as a sign of the interplay between policy accommodation and forward-looking pricing behavior that induces an important systematic component into domestic inflation. In such circumstances, if the central bank operating a managed float regime does not internalize the inflationary consequences of its past actions, might find itself in a vicious circle where trying to respond to inflationary pressures leads instead to a full accommodation of these. Inflation thus becomes more persistent not because of various forms of nominal rigidities but because of accommodative monetary policy.[16]

Exactly the opposite is the case of the Czech Republic, where instantaneous correlation is practically zero (even with a negative point estimate). This is consistent with observed path of Czech exchange rate growth, which, even though non-stationary, does not show any tendency to drift away from zero even though it exhibits large and persistent swings, to a large extent caused by speculative capital flows. For Poland the results show that consistently with imperfect, but still large, pass-through effect, exchange rate policy seems to act as to gradually and imperfectly accommodate inflationary pressures, hence not acting as a strong source of inflationary pressures itself.

The reasoning based on the results in right panel of Table 2 is confirmed also with a structural VEC analysis (see Warne, 1993). For all four countries we have identified the structural parameters by assuming a recursive ordering of three transitory shocks. With this we obtain also the estimates of structural contemporaneous effects of the permanent shock (only one in

[14] Note that for Poland even though the point estimate of β_{31} in Table 1 is smaller than 1, it is not significantly different from 1. This means that Polish risk premium and real interest rate differential have been stationary in the period under analysis. A change in the rate of depreciation of zloty thus does not yield lower equilibrium real interest rates.

[15] The asymptotic distribution of parameters in right panel of Table 2 is mixed Gaussian. For this reason, the corresponding standard errors are not reported. The figures in right panel of Table 2 were calculated using significant parameters in matrix only.

[16] Disinflation can be achieved with the cooperation of the government, for example by intervening administratively in the wage bargaining process. Long-run sustainability of administrative measures can be seriously questioned, however.

our systems), which is of our primary concern. The impulse responses of nominal exchange rate growth and inflation differential to the permanent shock – single stochastic driving trend - are plotted in Figure 7. As could be anticipated from previous discussion for Hungary and Slovenia we can clearly observe considerably smaller short-run fluctuation and much smoother transitional dynamics to the new steady state. Moreover, Hungary and Slovenia have a much faster pass-through effect. Most of the permanent change in exchange rate growth is transferred into inflation differential within one year. From the theoretical model in Coricelli et al. (2004) it also follows that if we restrict ourselves to analyzing only the responses to transitory shocks, as is most commonly done in SVAR-based studies of pass-through, the observed speed and magnitude of pass-through would be smaller. However, effects of transitory shock are of relatively small importance when we want to address the disinflation policies necessary to obtain a given policy objective i.e. entry to the Euro area. For this, the effects of permanent changes are of central importance.

4 Concluding Discussion

The purpose of the paper is to contribute to the literature on exchange rate pass-through estimation from methodological and empirical point of view. The empirical analysis is performed within the cointegrated VAR framework that enables the estimation of the equilibrium pass-through effect of nominal depreciation to domestic inflation. As shown by Johansen (2002), this requires a proper interpretation of cointegration coefficients, which, combined with economic theory, in the present framework translates into the problem of identification of the equilibrium pass-through effect. Systems of variables for pass-through estimation must contain all cruicial variables that account for complex macroeconomic interdependence in open economies. In such a case, we cannot discuss equilibrium effects of a one percentage point change in nominal depreciation on inflation without taking into account also the endogenous effect on at least the interest rates and output. As it turns out, this implies that identification cannot be automatically achieved for all orders of cointegration rank. Moreover, each possible value of cointegration rank leads to different conclusions about the equilibrium effects of exchange rate changes. In the present case, identification of equilibrium pass-through effect can be achieved. Cointegration rank three in a four-dimensional system implies that a permanent increase in depreciation rate is accompanied not only by the increase in inflation and nominal interest rates but also by a decrease in real output.[17] Compared to a hypothetical situation of rank two, where the equilibrium effect on real output would be zero, this leads to stronger conclusion about limited welfare effects of exchange rate flexibility when this is accompanied by policy accommodation. Combining this qualitative result with the size of estimated pass-through effect leads to some meaningful conclusions.

The paper finds a strong pass-through from nominal exchange rates to domestic inflation. In such a context, the dichotomy between inflation targeting and a regime with limited exchange rate variability (like ERM II) is more apparent than real. Moreover, in many instances, flexibility of exchange rates turns out to be a policy of accommodation of inefficiencies and monopoly power in non-tradable sectors. It is argued in this paper that a higher degree of policy accommodation leads to higher pass-through, implying that differences in exchange rate regimes contribute

[17] The effect on output is negative due to the increase in real interest rates.

significantly to differences in inflation rates we observe among CEEC-4. The presence of high and fast pass-through effect, which has been found for two countries with most accommodative exchange rate policies: Hungary and Slovenia, implies that disinflation can be achieved by a different choice of exchange rate policy. It follows from the discussion above that this need not have a negative effect on output. Thus, we can also conclude that disinflation would have low (if any) costs in terms of output decline and with potential gains in welfare.

What are the policy implications that can be drawn from presented empirical results? The most important conclusion is that in any policy design the important effect of the nominal exchange rate on prices should not be underestimated. It has been show theoretically that accommodative exchange rate policy is associated with high pass-through effect. This has been confirmed also in the empirical analysis. The path of nominal exchange rate within a more general exchange rate regime arrangement during disinflation should thus be given an important role. Moreover, the benefits of exchange rate flexibility seem rather limited. The reason for this are not structural characteristics of the coutnries analyzed per se, but accommodative monetary policy. If policy accommodation leads to a high pass-through into the CPI inflation rate, it reduces the expenditure switching effect and consequently the welfare gains from exchange rate flexibility.

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Appendix

Table 3: Multivariate misspecification tests and characteristic roots and trace tests for the I(1) systems

Czech Republic (1993:12 - 2002:7)						
Residual autocorrelation 1-6	$F(96,184) = 1.17$				p-val $= 0.18$	
Normality	$\chi^2(8) = 5.71$				p-val $= 0.68$	
Heteroscedasticity	$F(260,371) = 0.59$				p-val $= 1.00$	
	Modulus of 6 largest characteristic roots					
Unrestricted VAR	1.02	0.91	0.74	0.74	0.64	0.64
Under $r = 1$	1.00	0.91	0.74	0.74	0.63	0.63
Trace test	1.29	12.78	44.90	101.32		
p-value (asymptotic)	0.26	0.14	0.00	0.00		
p-value (bootstrap)	0.45	0.18	0.00	0.00		
r	3	2	1	0		
Hungary (1993:8 - 2002:7)						
Residual autocorrelation 1-7	$F(112,177) = 1.11$				p-val $= 0.27$	
Normality	$\chi^2(8) = 18.53$				p-val $= 0.03$	
Heteroscedasticity	$F(260,390) = 0.55$				p-val $= 1.00$	
	Modulus of 6 largest characteristic roots					
Unrestricted VAR	1.01	0.93	0.75	0.75	0.50	0.50
Under $r = 1$	1.00	0.91	0.75	0.75	0.49	0.49
Trace test	1.25	16.96	56.43	130.61		
p-value (asymptotic)	0.26	0.03	0.00	0.00		
p-value (bootstrap)	0.36	0.04	0.00	0.00		
r	3	2	1	0		

Note: The Paruolo test for the presence of I(2) trends rejects the presence of such trends for all countries and all choices of rank with a zero p-value. These test results are available upon request.

Table 4: Multivariate misspecification tests and characteristic roots and trace tests for the I(1) systems

Poland (1994:1 - 2002:4)						
Residual autocorrelation 1-7		$F(112, 132) = 1.28$			p-val $= 0.09$	
Normality		$\chi^2(8) = 19.51$			p-val $= 0.01$	
Heteroscedasticity		$F(320, 247) = 0.37$			p-val $= 1.00$	
	Modulus of 6 largest characteristic roots					
Unrestricted VAR	0.98	0.91	0.91	0.76	0.76	0.62
Under $r = 1$	1.00	0.89	0.67	0.67	0.59	0.59
Trace test	5.42	20.91	57.02	221.56		
p-value (asymptotic)	0.02	0.01	0.00	0.00		
p-value (bootstrap)	0.10	0.03	0.00	0.00		
r	3	2	1	0		
Slovenia (1993:3 - 2002:3)						
Residual autocorrelation 1-7		$F(112, 197) = 1.13$			p-val $= 0.22$	
Normality		$\chi^2(8) = 10.43$			p-val $= 0.24$	
Heteroscedasticity		$F(180, 478) = 0.87$			p-val $= 0.38$	
	Modulus of 6 largest characteristic roots					
Unrestricted VAR	0.97	0.89	0.41	0.41	0.34	0.34
Under $r = 1$	1.00	0.88	0.41	0.41	0.35	0.35
Trace test	0.28	26.97	204.486	406.34		
p-value (asymptotic)	0.60	0.00	0.00	0.00		
p-value (bootstrap)	0.41	0.06	0.00	0.00		
r	3	2	1	0		

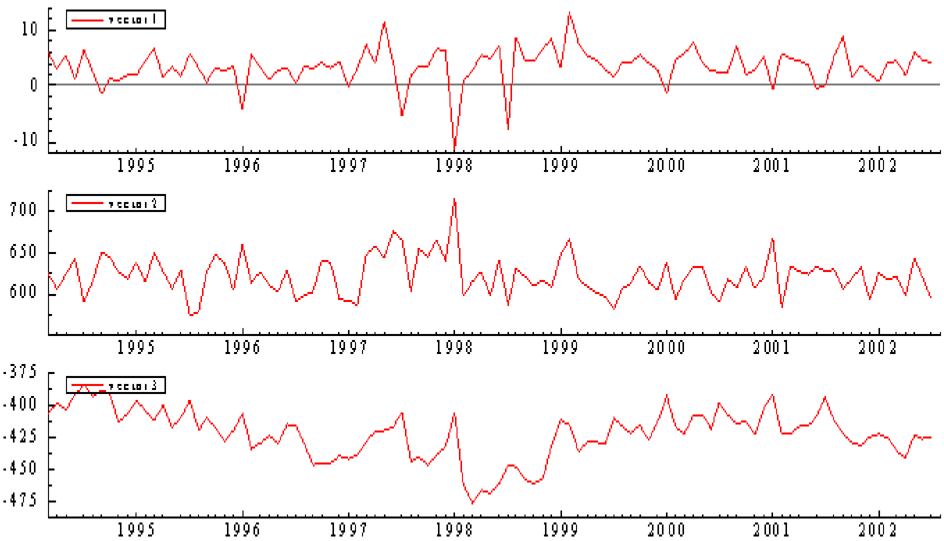

Figure 3: Cointegrating Vectors (unconcentrated) for the Czech Republic

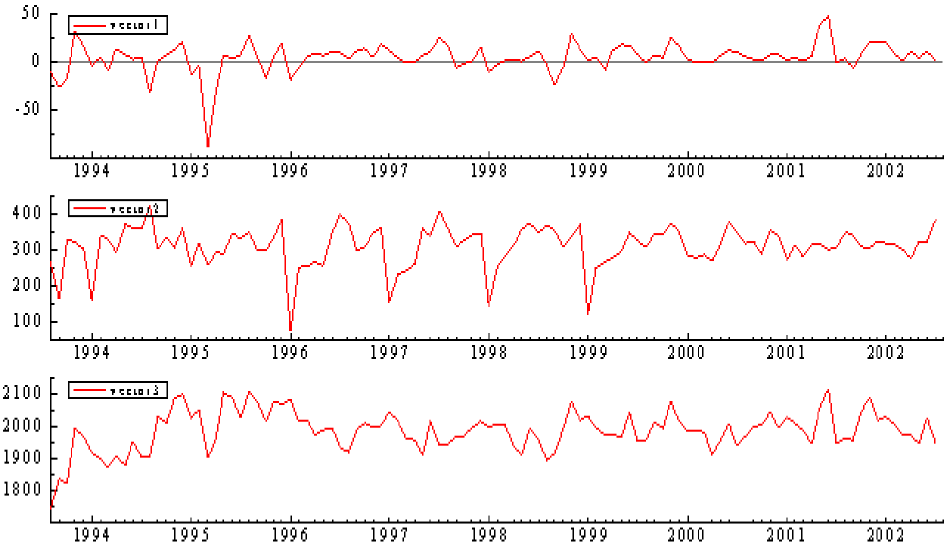

Figure 4: Cointegrating Vectors (unconcentrated) for Hungary

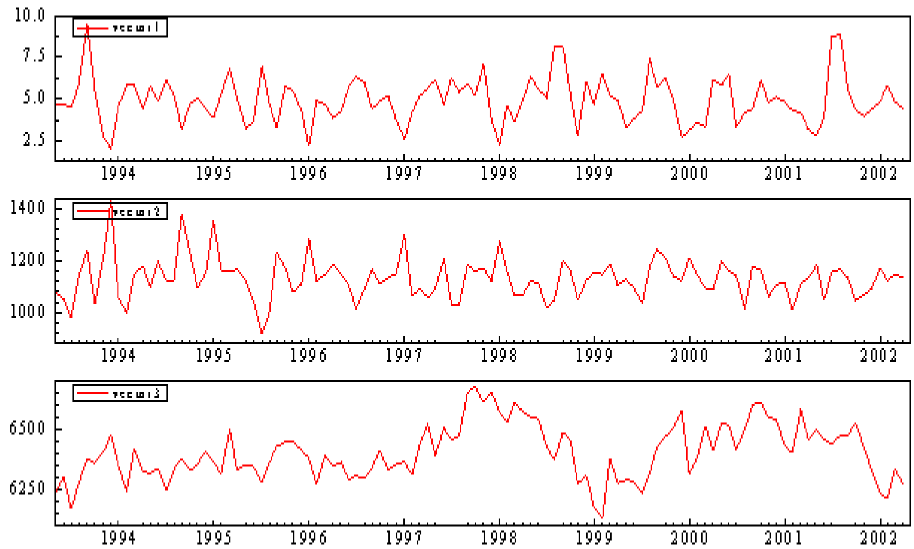

Figure 5: Cointegrating Vectors (unconcentrated) for Poland

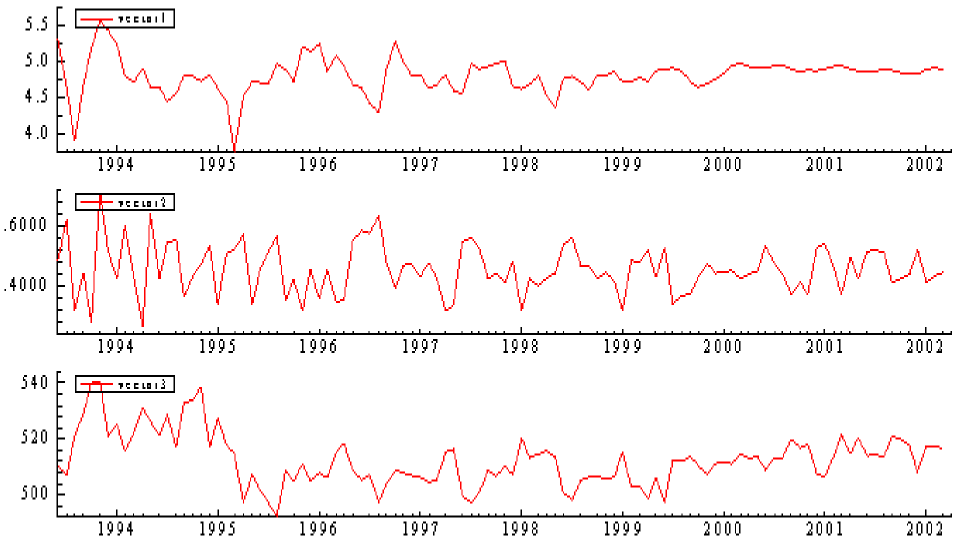

Figure 6: Cointegrating Vectors (unconcentrated) for Slovenia

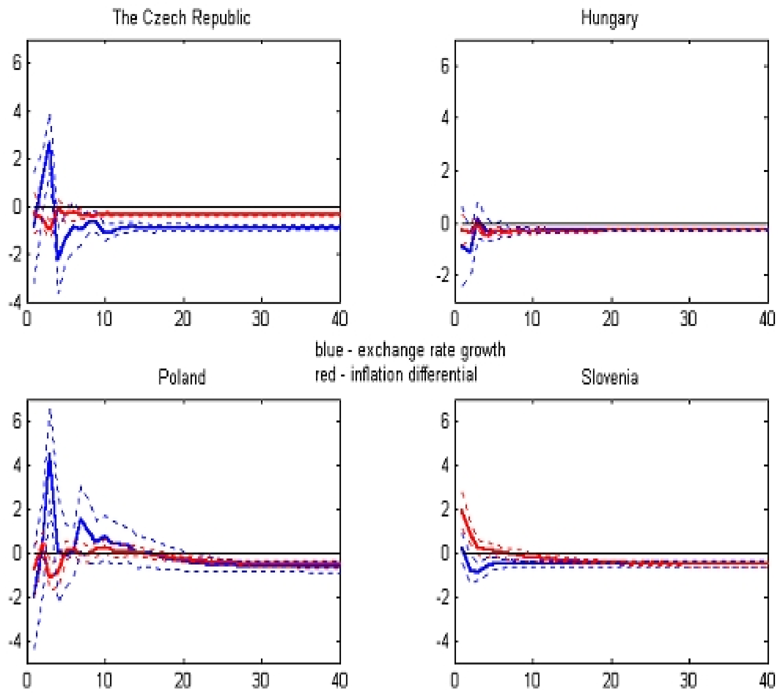

Figure 7: Impulse responses of nominal exchange rate growth and inflation differential to the permanent shock (with 95 % asymptotic confidence intervals)