DISCUSSION PAPER SERIES

No. 3894

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TRANSITION ECONOMICS

Centre for Economic Policy Research

www.cepr.org

Available online at: www.cepr.org/pubs/dps/DP3894.asp

ISSN 0265-8003

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Discussion Paper No. 3894
May 2003

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CEPR Discussion Paper No. 3894

May 2003

ABSTRACT

Exchange Rate Pass-Through in Candidate Countries

In this Paper we analyse the link between the choice of exchange rate regime and inflationary performance in four EU Accession Countries: the Czech Republic, Hungary, Poland and Slovenia. Estimation of pass-through effect of exchange rate changes to CPI inflation is complemented by I(2) co-integration analysis of stochastic nominal trends. The results allow a clear ranking of countries according to the size of the pass-through effect and the importance of exchange rate shocks to overall inflationary performance. In particular, we find that perfect pass-through effect can be associated with accommodative exchange rate policy, which can moreover become the most important source of inflationary pressures. The analysis suggests that for CEEC-4 the early adoption of the euro can provide the most efficient framework for reducing inflation.

JEL Classification: C32, E42, E52 and E58
Keywords: EMU accession, I(2) co-integration analysis and pass-through effect

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Submitted 18 March 2003

1. Introduction

Following accession to the European Union, candidate countries (CEECs) will eventually have to adopt the euro, as no opt-out clause is allowed for new entrants. Therefore, the main open question about exchange rate policy for new members is the speed of entry into the eurozone. Official positions of the European Commission and the ECB indicate that CEECs should go through the ERM2 mechanism before adoption of the euro. This would imply two years in the ERM2 system with an agreed central parity and a ±15 % band, with a review of Maastricht indicators at the end of the first year. As a result, the minimum time lag for adoption of the euro is two years after joining the EU. One could argue that such a rigid timing might be bypassed. After all, Italy and Finland did not go through ERM2, and Greece entered the eurozone six months earlier the two-year deadline. There is clearly some room for discretion .

Even leaving aside the issue of the interpretation of the specific rules that apply before adoption of the euro, its timing is one of the main macroeconomic issues related to accession. There has been a lively debate on the path towards the euro, focusing on: (i) the pre-conditions for its adoption, within the well-known optimal currency area theory; (ii) the ability of CEECs to fulfill Maastricht criteria, especially inflation criterion, and (iii) the desirability and feasibility of maintaining some flexibility in exchange rates and an independent monetary policy.

In this paper we concentrate on the interplay between the exchange rate regime and the speed of convergence of inflation rates between CEECs and the eurozone. Specifically, we study the phenomenon of pass-through from exchange rate changes to domestic inflation in four CEECs. The exchange rate was the main nominal anchor in most transition economies at the beginning of transition. To curb inflation and maintain macroeconomic stability, the Czech Republic, Hungary and Poland introduced exchange rate-based stabilization programs in the early 90's, while Slovenia followed a combination of targets on M3 and tightly managed exchange rate. Over time several CEECs have moved towards a more flexible regime (*i.e.* the Czech Republic, Poland, Hungary and the Slovak Republic), Slovenia instead continued to maintain managed float, while others, the Baltics and Bulgaria, opted for currency boards. There is a question of whether a move to more flexible exchange rate regimes have helped transition economies to carry on independent monetary policy and to more effectively respond to external nominal shocks. Within flexible exchange rate regimes one has to distinguish between free float and managed float. While in the former exchange rate responds to unpredictable external shocks, in the latter, 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 nominal exchange rate. Such a policy is likely to be incorporated into the pricing decisions of economic agents. For this reason, a strong correlation between exchange rate movements and inflation rates can be observed in managed float regimes. Following this line of reasoning, we estimate the extent of exchange rate pass-through in four candidate countries and find that regimes with a more accommodative stance of exchange rate policy generate higher pass-through than free float. 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.

The empirical importance of exchange rate pass-through has been analyzed in a growing number of papers in recent years. Campa and Goldberg (2001) estimate pass-through to import prices for 25 OECD countries over the period 1975 to 1999. 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. Darvas (2001) provides evidence for the same set of countries as here (the Czech Republic, Hungary, Poland, and Slovenia) for the period 1993 to 2000. However, the previous studies generally suffer from methodological weaknesses, which limit the robustness of the empirical estimates of the pass-through. Using a cointegrated vector autoregressive model we are able to identify the pass-through from exchange rates to prices and to estimate the importance of shocks to the nominal exchange rate in the movements of domestic inflation for the CEEC-4 (the Czech Republic, Hungary, Poland, and Slovenia). With the rather novel econometric approach based on I(2) analysis of nominal variables in cointegration framework, a relationship can be established between changes in the nominal exchange rate and Consumer Price Index (CPI), abstracting from differences in the effects of the exchange rate on import and export prices on one hand, and issues related to Local Currency Pricing (LCP) and Producer Currency Pricing (PCP) on the other hand[1]. As the effect of exchange rate changes on CPI in CEEC-4 is firmly established, policy implications can be drawn for accession countries.

The empirical analysis indicates that, especially for Slovenia and Hungary, there is a very large pass-through from exchange rates to domestic inflation. A smaller impact is found for the Czech Republic and Poland. Similarly, in Slovenia shocks to the exchange rate play a dominant role in determining inflationary pressures. By contrast, in Poland autonomous shocks arising from monopolistic behavior in goods markets and wage pressures dominate the inflation process, with smaller effects from exchange rate shocks. Note that Slovenia and Poland have followed rather different exchange rate policies. Slovenia apparently targeted the real exchange rate throughout the period, trying to maintain external competitiveness. Poland, after the initial use of the exchange rate as a nominal anchor, has progressively moved toward a more flexible exchange rate, culminating in the floating regime that started in April 2000. Therefore, one can conjecture that such different exchange rate regimes have had a fundamental impact on domestic inflation. The real exchange rate rule in Slovenia was probably internalized by price setters, thus becoming a persistent source of inflation. In fact, although Slovenia apparently had the best fundamentals of CEEC-4, it has been unable to reduce inflation below 6-8% in the last five years. By contrast, Poland did not follow an accommodative exchange rate policy. Considering as well that Slovenia is a much more open and smaller economy than Poland, one would expect a smaller pass-through in Poland and a smaller role of exchange rate shocks driving the domestic inflationary process. Hungary and the Czech Republic lie between the two extreme cases, with Hungary more similar to Slovenia and Poland more to the Czech Republic.

The analysis has a number of clear policy implications. The large pass-through from exchange rates to domestic inflation reduces the scope for flexibility in exchange rates. Even abstracting from the issue of propagation of exogenous shocks originating in international financial markets (see Habib (2002) on this issue), flexible exchange rates are not an effective instrument for absorbing asymmetric real shocks (Masten, 2002). Large pass-through is likely to cause policy-makers to attempt ex-post to drive the exchange rate in a way that maintains external competitiveness. As can be seen in the case of Slovenia, such 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. For a small open economy this may imply the adoption of fixed exchange rates. Luckily, candidate countries have the point of arrival, the euro, already set. Their main policy decision is how fast to enter the euro. Results in this paper suggest that there are no significant advantages to delaying entry.

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[1] An extensive survey on different approaches to exchange rate pass-through and their empirical validity is provided by Choudhri, Faruqee, and Hakura (2002).

The paper proceeds as follows. Section 2 presents stylized facts on inflation and exchange rate behavior in CEEC-4. After showing the long-run trend appreciation of the exchange rate and its connections with the Balassa-Samuelson effect, the section emphasizes the relationship between exchange rate regime and inflation dynamics. Section 3 contains the main empirical analysis of the paper, focusing on the pass-through. It is shown that pass-through is highly significant in the four candidate countries examined, although important differences emerge. The effect of pass-through appears to be larger in Slovenia and Hungary than in the Czech Republic and Poland. While Slovenia and Hungary have engaged in relatively tightly managed exchange rates, the Czech Republic and Poland have let their exchange rate float more freely, at least recently. Additionally, the Czech Republic and Poland introduced inflation targets, which helped monetary authorities to maintain inflation at lower levels than in Slovenia and Hungary. However, in Poland inflation declined during a sharp slowdown of the economy. Only the Czech Republic has been able to credibly follow a policy of successfully targeting inflation. It is likely that the much lower external debt of the Czech Republic, compared to Poland, contributed to make a policy of targeting inflation and flexible exchange rates more credible. It would seem that more predictable exchange rate policies, like those followed in Slovenia and Hungary (and Poland until 2000) tend to be associated with larger pass-through coefficients. The size and openness of the countries are also important factors. Section 4 concludes.

2. Stylized Facts on Inflation and Exchange Rate Dynamics

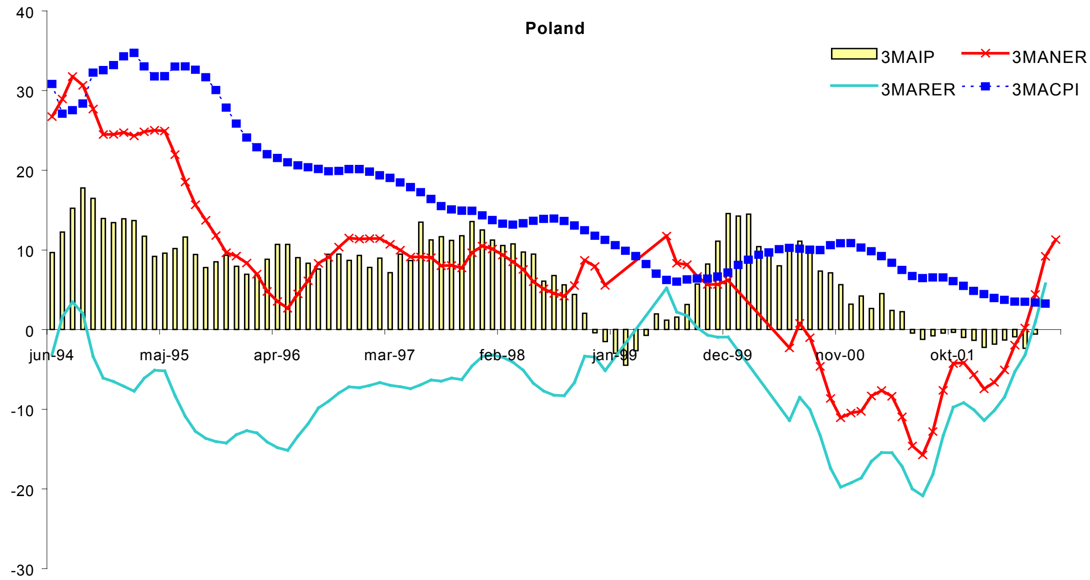
Following the initial jump in price levels associated with price liberalization, inflation has declined gradually in CEEC-4. Reduction to single-digit inflation was much faster in Slovenia and the Czech Republic, countries less affected by large stocks of debt and the attendant need to finance large debt service payments. However, inflation rates seem to be more stubborn in Slovenia and Hungary than in the Czech Republic and Poland. In the last 3-4 years, inflation hovered around 8-9 % in Slovenia and Hungary, with some sign of small decline only in 2002 in a period of economic slowdown. The sharp decline in the Czech Republic and Poland reflects two different realities. The Czech Republic has been successful in reducing inflation through an effective and credible policy of inflation targeting. In Poland the fall in inflation, which declined to around 1% annual rate in 2002, reflects perhaps an overshooting of the desired decline. This was the result of an excessively tight monetary policy that negatively affected the economy during a period of general economic slowdown in Europe. Output performance in Poland during 2002 has been among the worst in candidate countries. A sharp fall in demand and output and persistent unemployment rate at around 18% have contributed to the fall in inflation. This suggests that Polish inflation is rather sensitive to the cyclical position of the economy.

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. A component of this trend appreciation can be considered an equilibrium phenomenon, in line with the Balassa-Samuelson effect, that affects real exchange rates in a catching-up phase. However, there is, in addition, a shorter-term dynamic process connecting exchange rates and inflation. Figure 1 illustrates the dynamics of inflation, and the growth of nominal and real exchange rates in CEEC-4 since 1995. The figure also contains an indicator of the state of the economy, proxied by the growth of industrial production. [2]

Figure 1 indicates the presence of at least two different patterns in CEEC-4. On one hand, there is the case of Slovenia and Hungary. Nominal exchange rate and inflation move broadly together. On the other hand, the Czech Republic and Poland display a high correlation in the movements of nominal and real exchange rates. Finally, Poland seems to show a stronger relationship between inflation and the state of the economy, along the lines of the traditional Phillips curve.

[2] Growth rates of industrial production (IP), real exchange rate (RER), nominal exchange rate (NER), and consumer price index (CPI) are presented by 3-month moving averages.

Figure 1. Growth of Industrial Production, Growth of Nominal and Real Exchange Rates and Inflation in CEEC-4



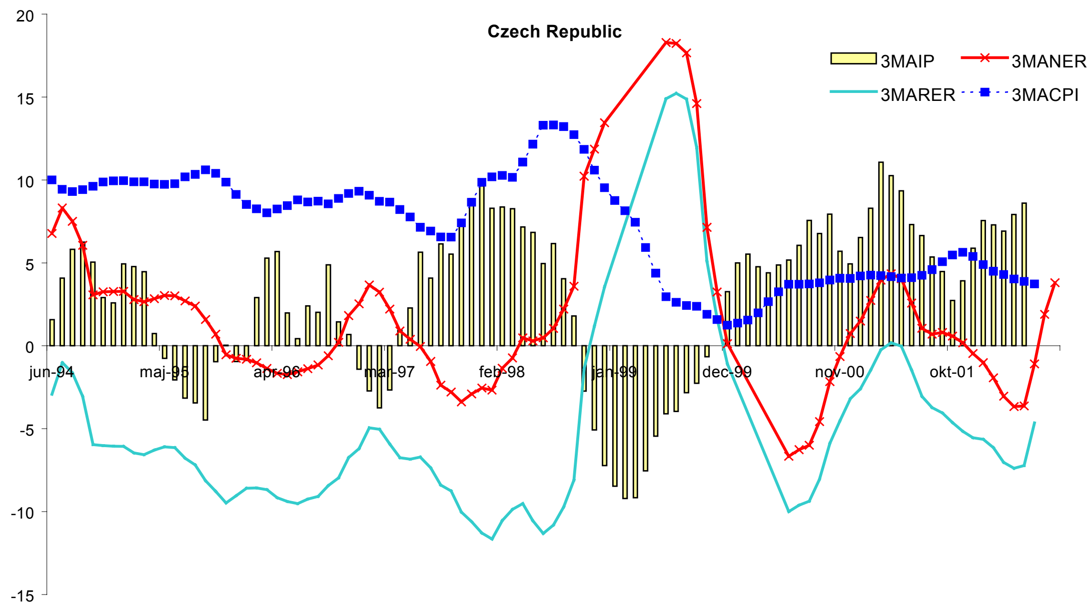
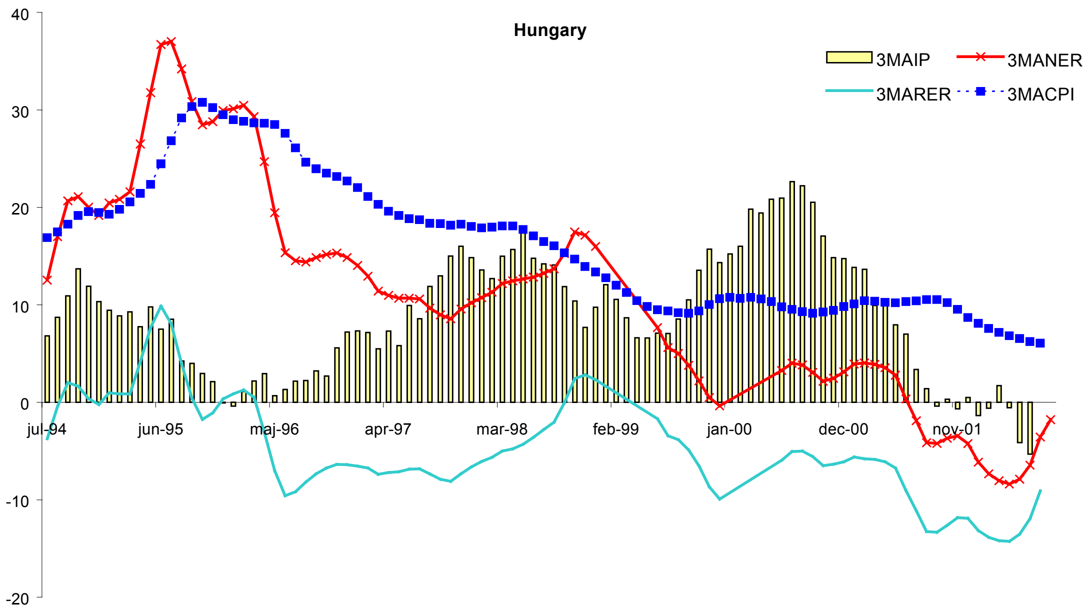
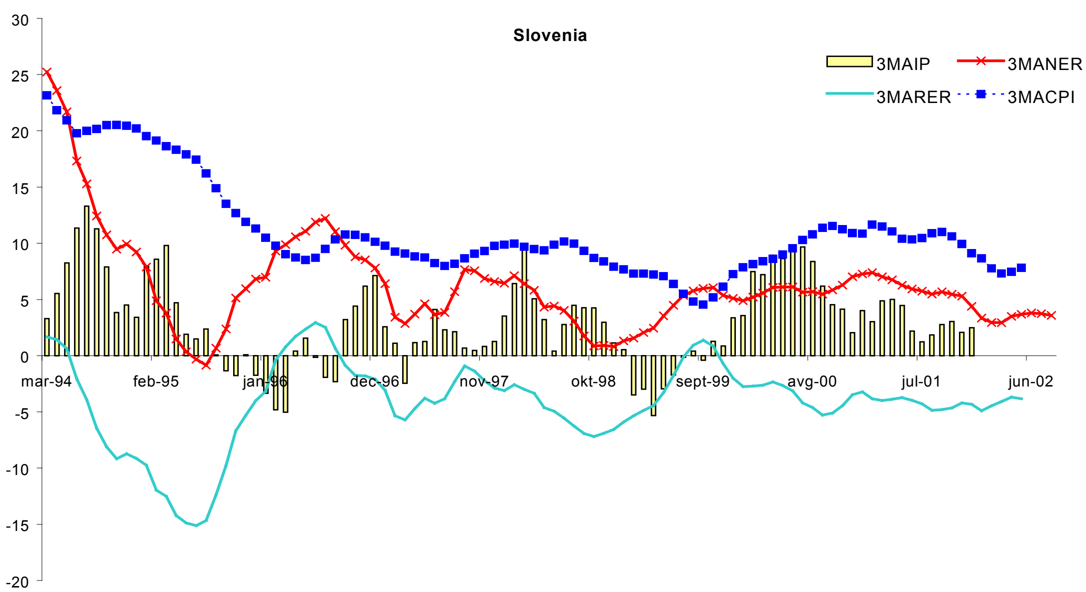


Figure 1: continued





Source: Datastream.

In addition to the above considerations, different patterns of inflation dynamics in four candidate countries seem to be associated with different exchange rate regimes. Table 1 presents the exchange rate regimes in selected transition economies with shifts in regimes from less flexible - or fixed - to more flexible as they occurred during transition. Dates presented in bold stand for current exchange rate regime.

Table 1: Exchange Rate Regimes in CEEC-4

	Conventional Peg	**Narrow Band**	**Tightly Managed**	**Broad Band**	**Managed Float**	**Relatively Free Float**
Czech Republic	January 1991			February 1996	**May 1997**	
Hungary	→	March 1995		**October 2001**		
Poland	→	May 1991				**April 2000**
Slovenia			**February 1992**			

Source: Arratibel, Rodriguez-Palenzuela, and Thimann, 2002.

The evolution of exchange rate regimes has been affected by the liberalization of capital controls (Corker et al., 2000). In addition to different exchange rate regimes and the liberalization of capital flows, CEEC-4 employed different monetary policies. While the Czech Republic and Poland set inflation targets, Hungary and Slovenia stuck to the exchange rate and M3 targets, respectively. As shown in Coricelli and Jazbec (2001), the switch of the exchange rate regimes from a less to a more flexible framework with respect to the regimes employed at the beginning of 90's, broadly corresponds to a diminishing effect of structural reforms on the real exchange rate determination in transition economies. In the mid-90's, CEEC-4 were on average in the fifth or sixth year of the transition process when productivity and demand factors began to affect the real exchange rate more than structural reforms. Additionally, the shift in exchange rate regimes in the Czech Republic and Hungary broadly corresponds to the liberalization of trade and current account convertibility, while in Poland a shift toward free floating happened only in 2001. Although Slovenia officially targeted M3 throughout the last decade, the tightly managed exchange rate regime was substantially supported by capital controls on short-term capital flows together with extensive sterilization policy. Despite the variety of approaches to the exchange rate policy, the CEEC-4 have all made substantial progress in reducing inflation, which has, on average, been below 10 percent since 1998. Although the anti-inflationary programs in CEEC-4 have been successful in bringing down inflation from almost hyperinflationary levels at the beginning of transition, the inflation rates are still above the rates required for entry to EMU. As already mentioned, part of the reasons for higher inflation rates could be found in the working of Balassa-Samuelson effect and the remaining convergence of relative prices (on the latter see Čihak and Holub (2002)). However, it is suspected that the combination of an exchange rate regime and monetary policy could substantially contribute to the differences in inflation rates in CEEC-4 as the Czech Republic and Poland have on average produced lower inflation rates than Hungary and Slovenia in the last three years. As the Czech Republic and Poland maintain relatively less managed exchange rate regimes than Hungary and Slovenia, and additionally employ inflation targets, it is believed that the combination of a relatively more flexible exchange rate regime and inflation target produces lower inflation. Abstracting from the relationship between the exchange rate regime and inflation, it is worth pointing out that although CEEC-4 liberalized their trade and foreign exchange system to the standards and performance norms of advanced industrial economies, they still lag behind reforms in price liberalization, especially in the public sector and non-market services.

3. Exchange Rate Pass-Through

With respect to exchange rate regimes, all CEEC-4 moved from fixed to more flexible exchange rate regimes during transition, perhaps also to be able to curb inflation rates toward required levels, although the main reason for the move was believed to be the pressure caused by a surge in foreign capital inflows. In so doing, CEEC-4 added a potential new source to higher inflation rates and its stubbornness in addition to the working of the Balassa-Samuelson effect and relative price convergence.

As real appreciation in transition economies has resulted in higher inflationary pressure rather than nominal appreciation, part of the inflationary pressure could derive from goods and labor market rigidities. It is thus not surprising that countries with higher relative non-tradable wage growth – either growth of wages in market or public services – on average face higher inflation rates. This brings up the issue of the relationship between exchange rate policy and disinflation in an economy with price-wage and inflation inertia in the non-tradable sectors. A useful reference framework for discussing the costs and benefits of different speeds of disinflation is a two-sector model with monopolistic power in the non-tradable sector. In the context of perfect capital mobility, interest rates in candidate countries would be determined by foreign interest rates and expected depreciation of the exchange rate. In the staggered price model of Calvo (1983) with price level inertia in the non-tradable sector, it is easy to show that by reducing the rate of depreciation of the exchange rate, a country can reduce the overall rate of inflation with little if any fall in output in the non-tradable sector. A more interesting model is a recent extension of staggered price models by Calvo, Celasun and Kumhof (2002) that takes into account the average rate of inflation for the price setting of firms in a monopolistically competitive market. The intuition of the model is that firms choose a price rule that includes a revision of price schedule depending on the rate of inflation in the economy. This implies that firms internalise the effects of policies such as that of a persistent rate of depreciation of a central bank that wants to target the real exchange rate. In this version of the model, there is inflation inertia in addition to price-level inertia. The implication is that a disinflation policy implemented through a reduction of the rate of depreciation of the exchange rate induces a temporary decline in output in the non-tradable sector. However, in this model, disinflation brings welfare gains as it reduces the welfare losses associated with monopolistic power in the non-tradable sector. A disinflation policy can thus be seen as a way of reducing the welfare losses of monopolistic price setting. This line of reasoning seems very relevant for an exchange rate policy in candidate countries.

3.1. Methodological Issues and Limitations of Existing Empirical Studies

The empirical analysis presented below addresses two issues. The first is the identification of pass-through effect, and the second is reconciliation of results with theoretical foundations of New Keynesian models that enables us to address policy implications for the process of accession to EMU.[3] The measure of pass-through we are interested in is the effect of changes in nominal exchange rate growth on CPI inflation, a common final target variable of monetary authorities. In particular, we are interested in equilibrium effects of exchange rate changes on inflation, real interest rates and output, and short-run adjustment to deviations from equilibrium.

The analysis in this paper improves existing studies of pass-through in three ways. First, the analysis is conducted within the framework of cointegrated vector autoregression model (CVAR), while use of structural VAR methods is more common in the literature. See, for example, McCarthy (2000),

[3] This means that we are not interested in gauging the prevalence of local currency or producer currency pricing in candidate countries.

Campa and Goldberg (2002) and Choudhri, Faruqee, and Hakura (2002).[4] 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 very surprising since theoretically long-run co-movement of prices and exchange rate seems very plausible. Not properly testing for cointegration causes two problems: first, the inconsistency of estimated parameters; second, neglecting cointegration means neglecting the intrinsic meaning of equilibrium long-run relationship between the nominal exchange rate and prices. Identifying long-run equilibrium relations and analyzing the adjustment to disequilibria allows us to evaluate some important theoretical aspects of New Keynesian models.

Second, the pass-through effect is estimated without relying on the identification of structural shocks. These can be identified using non-testable restrictions, which are very often imposed arbitrarily and in high-dimensional systems even with weak theoretical justification. In this respect our conclusions are more robustly supported by the data.

The third potential deficiency of existing studies is that in general they do not address the possibility of prices, the nominal exchange rate and nominal wages being integrated of order 2, which is an increasingly common finding in the literature (see Banerjee, Cockerell, and Russel (2001), Juselius (1999, 2001), Coenen and Vega (2001) and Ericsson, Hendry, and Prestwich (1998)). Also Kongsted (1998), for example, analyzes pricing-to-market behavior explicitly within an I(2) cointegration framework. I(2)-ness of prices effectively means that inflation rate is not stationary, *i.e.* it is driven by a stochastic trend. Shocks to inflation in this respect have a fully persistent effect on the level of inflation. Treating inflation as stationary results in invalid statistical inference.[5] Thus, all results obtained without testing for I(2)-ness in the price level before treating inflation as stationary can be seriously questioned.

Without controlling for the state of the economy, empirical investigation of the pass-through effect on aggregated data suffers from an identification problem of pure effects of exchange rate changes on prices. Without the controls the estimate of pass-through in an economy facing real appreciation will

[4] Billmeier and Bonato (2002) use also a CVAR approach, but there the pass-through effect into two different price indices is incorrectly identified.

[5] Econometric investigation of the pass-through effect on quarterly data in accession countries is seriously hindered by short time series. Namely, the inclusion of a number of control variables that are important for exchange rate determination quickly leads to a dimension of the system that does not allow for a fully-fledged cointegration analysis. This data intensive technique employed here led us to use monthly data.

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necessarily be larger than one, which is what Campa and Goldberg (2002) report for certain countries and also what we found in our preliminary I(2) analysis. But such a coefficient clearly does not take into account only the effects of the nominal exchange rate on the price level, but also the effects of other economic variables, and hence cannot be interpreted as a pass-through effect in the usual sense. For this reason the analysis also includes a measure of real output to proxy the state of the business cycle and productivity developments. Furthermore, foreign inflationary developments are controlled for, and interest rate spread is included to account for international parity forces working through uncovered interest rate parity.

3.2 A Brief Description of I(2) Model

In this section we provide a minimal theoretical exposition of I(2) model that should be sufficient for understanding the results presented below. For a detailed analysis of moving average representation of I(2) models the reader should consult, for example, Johansen (1995a), Paruolo (1996), Rahbek et al. (1999) and references therein. Consider a p-dimensional VAR with deterministic term D_t (in the present case it contains a constant, centered seasonal dummies and selected impulse dummies, which account for outliers). Under the assumption of data being I(1) we would consider the following model that combines first differences and level terms as

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

with a corresponding reduced rank condition $\Pi = \alpha\beta'$. For I(2) analysis it is convenient to rewrite the model in terms of acceleration rates.

$$\Delta^2 X_t = \Pi X_{t-1} - \Gamma \Delta X_{t-1} + \sum_{i=1}^{k-2} \Psi_i \Delta^2 X_{t-i} + \Phi D_t + \varepsilon_t \ ,$$

where $\Gamma = I - \sum_{i=1}^{k-1} \Gamma_i$ and $\Psi_i = -\sum_{j=i+1}^{k-1} \Gamma_i, i = 1,...,k-2$. Then the I(2) model is defined by two reduced rank conditions

$$\Pi = \alpha\beta' ,$$

where α and β are $p \times r$ matrices of full rank $r < p$, and

$$\alpha_{\perp}^{'} \Gamma \beta_{\perp} = \xi \eta',$$

where ξ and η are $(p-r) \times s$ matrices with $s \leq p-r$. (See Johansen (1995a) for a proof.) Furthermore, let $\alpha_{\perp}$ and $\beta_{\perp}$ be decomposed into I(1) and I(2) directions:

$$\alpha_{\perp} = \{\alpha_{\perp 1}, \alpha_{\perp 2}\} \text{ and } \beta_{\perp} = \{\beta_{\perp 1}, \beta_{\perp 2}\},$$

where $\alpha_{\perp 1} = \alpha_{\perp}(\alpha_{\perp}'\alpha_{\perp})^{-1}\xi$ is a $p \times s$ matrix, $\alpha_{\perp 2} = \alpha_{\perp}\xi_{\perp}$ a $p \times (p-r-s)$ matrix. $\beta_{\perp 1} = \beta_{\perp}(\beta_{\perp}'\beta_{\perp})^{-1}\eta$ and $\beta_{\perp 2} = \beta_{\perp}\eta_{\perp}$ have analogous dimensions, and $\xi_{\perp}$ and $\eta_{\perp}$ are orthogonal complements to ξ and η respectively. From this decomposition it is evident that the number of I(1) stochastic trends in the model is *s*, and the number of I(2) trends is *p-r-s*.

Under the two reduced rank conditions the moving average (MA) representation of the VAR model is

$$X_t = C_2 \sum_{s=1}^{t} \sum_{i=1}^{s} (\varepsilon_i + \Phi D_i) + C_1 \sum_{i=1}^{t} (\varepsilon_i + \Phi D_i) + C_2(L)(\varepsilon_i + \Phi D_i) + A + Bt,$$

where coefficients *A* and *B* depend on initial conditions, and the main matrix that we are interested in here, C_2, can be expressed as

$$C_2 = \beta_{\perp 2}\left(\alpha_{\perp 2}^{'} \theta \beta_{\perp 2}\right)^{-1} \alpha_{\perp 2}^{'}.$$

The matrix θ is defined as $\theta = \Gamma \overline{\beta} \overline{\alpha}' \Gamma + \sum_{i=1}^{k-1} i\Gamma_i$, $\overline{\beta} = \beta(\beta'\beta)^{-1}$ and analogously for $\overline{\alpha}$.

From the MA representation we can observe that the process is dominated by the second order stochastic trend that can be denoted by $\alpha_{\perp 2}^{'} \sum_{s=1}^{t} \sum_{i=1}^{s} \varepsilon_i$. These second order stochastic trends affect the variables in X_t with weights determined by $\beta_{\perp 2}$.

In this part of the analysis we wish to focus on these two matrices. For each country under analysis we identify one stochastic I(2) trend and we are interested in the relative importance of innovations to different variables in this trend. It is of particular interest to see (from $\alpha_{\perp 2}$ vector) whether shocks to the nominal exchange rate most strongly contribute to the trend that dominates the long-run behavior of nominal quantities in the economy, or whether this trend is equally or more strongly affected by shocks to the CPI or PPI, or by shocks to nominal wages. In the first case this would imply that the majority of inflationary pressures come via the exchange rate, roughly speaking, due to the pass-through effect.[6] The second case would correspond to the majority of inflationary pressures coming from the pricing behavior of firms. In particular, a higher weight of CPI than PPI points in the direction of monopoly pricing behavior in the non-tradable and service sector.[7] And finally, a high share of nominal wages would imply that important inflationary pressures come from aggressive trade unions.

Secondly, of particular interest is also the $\beta_{\perp 2}$ vector. It gives the proportions through which the I(2) trend feeds into individual variables, thus indicating which variable is most adequately described as an I(2) variable or an I(1) variable. Here we explore whether the index of industrial production is also an I(2) variable, or whether the I(2) stochastic trend is predominately a nominal one. [8]

It is worth repeating that the aim of the I(2) analysis presented in this section is not to directly estimate the pass-through effect,[9] but to highlight the relative importance of shocks to different nominal variables in the identified nominal I(2) trends.[10] For this reason we consider a system of domestic nominal variables, to which the index of industrial production has been added in order to explore a potential presence of I(2) trends also in this variable. In particular, for every country we consider the following systems of variables:

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[6] The pass-through effect operates broadly through three basic channels: (1) direct effect through prices of imported goods in the CPI; (2) effect through prices of imported intermediate goods; and (3) the effect through expectations that includes also the expected responses of monetary policy (Garcia and Restrepo, 2001).

[7] A formal analysis of this notion would need to explicitly distinguish between tradable and non-tradable prices. Unfortunately these indexes are available only for Slovenia.

[8] Other results of I(2) analysis are not reported as they are not of central importance. For example, for three out of four countries there is only one cointegrating vector, which necessarily implies that this relation is polynomially (or dynamically) cointegrating. For the case of two cointegrating vectors (Slovenia), we find a directly stationary combination of levels of variables; however, an identification issue arises whenever we have in one relation more variables that are endogenously affected by the nominal exchange rate (see above). The same identifications issue applies also for the stationary relations among first differences of variables (β_1 vectors).

[9] These issues are more explicitly addressed in cointegration analysis in the I(1) framework in the next section.

[10] We have experimented also with trivariate systems containing only a selected domestic price index, the nominal exchange rate and the German price index. In all cases the coefficient of the exchange rate relative to domestic prices (in a relation that cointegrates directly from I(2) to I(0)) is consistently larger than one. Results differ when PPI is used instead of CPI, for example, and across countries, but the nevertheless, the basic finding remains unaltered. These results are available from the authors upon request.

$$X_t = \left(y_t, e_t, cpi_t, ppi_t, w_t\right)'$$

$t = 1993:1,...,2002:5$,

y_t - index of total industrial production,[11]

e_t - nominal exchange rate (units of domestic currency per Euro),

cpi_t - consumer price index (food and energy excluded),

ppi_t - producer price index, and

w_t - average nominal wages.[12]

Based on the results of I(2) analysis[13] we continue with the analysis of an I(1) system of industrial production index, growth of nominal exchange rate, difference between domestic and foreign inflation, and interest rate spread. In this system we are able to identify an estimate of pass-through effect.

3.3 I(2) Analysis of Nominal Stochastic Trends

Testing for cointegrating rank has been conducted in unrestricted VAR models. To each system a linear trend term has been added, while impulse dummies have accounted for outliers.[14] It is worth noting that testing for rank remains asymptotically valid even in the presence of such dummies. 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. Three lags proved to be sufficient for Hungary, the Czech Republic and Poland, while 4 lags were included in the final model for Slovenia. All models are statistically well specified as seen from Table A1 in the Statistical Appendix.[15]

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[11] This is the only measure of real output available on monthly frequency.
[12] Manufacturing only for the Czech Republic.
[13] We do not use a direct reduction from a I(2) to I(1) model, because formal testing revealed that it would not yield an economically meaningful nominal to real reduction. Instead, I(2) analysis determined our choice of variables in I(1) models and, more importantly, contemplated the analysis of the size of pass-through with most important sources of nominal shocks, which is a crucial element in policy analysis.
[14] Deterministic trend enters as orthogonal to cointegrating space, both in I(2) and I(1) cointegration analysis, respectively, because this yields economically more meaningful results.
[15] The only sign of misspecification is the test of normality of residuals for Poland; however, it has been checked that this is due to excess kurtosis, and in this case testing for rank is moderately robust (Hansen and Rahbeck, 1999).

Table A1 also reports the roots of the systems with and without rank restrictions being imposed. A high number of large roots in each system is already a sign of I(2) trends. This is confirmed in Table 2, where formal tests for the presence of I(2) are presented. This is based on the methodology developed by Paruolo (1996). With the tests statistic $S_{r,s}$, we test the null hypothesis that $r \leq p$ and the number of I(2) components is equal to $p-r-s$, against the alternative $r \leq p$.[16] Based on the results, one I(2) trend is the preferred choice for all countries.[17] The ranks chosen are 1 for Hungary, the Czech Republic and for Poland, and 2 for Slovenia. From the last line of each panel of Table A1 we can in addition observe that under these rank restrictions our choice of one I(2) trend was correct, as in all cases there is only one very large root left in the system.

Table 2: Tests of I(1) and I(2) Cointegrating Ranks

r	***$S_{r,s}$***					***Q_r***
	The Czech Republic					
0	489.6 (0.00)	289.3 (0.00)	205.5 (0.00)	148.7 (0.00)	98.7 (0.00)	84.1 (0.00)
1		272.0 (0.00)	138.6 (0.00)	70.7 (0.09)	32.0 (0.95)	29.1 (0.76)
2			204.6 (0.00)	79.3 (0.00)	14.0 (0.99)	11.8 (0.93)
3				100.1 (0.00)	8.2 (0.93)	5.89 (0.71)
4					44.4 (0.00)	0.27 (0.60)
	Hungary					
0	333.9 (0.00)	247.1 (0.00)	166.5 (0.00)	111.2 (0.01)	81.4 (0.06)	89.2 (0.00)
1		245.8 (0.00)	159.8 (0.00)	89.2 (0.00)	38.4 (0.75)	43.2 (0.13)
2			123.4 (0.00)	50.4 (0.08)	15.7 (0.98)	17.2 (0.63)
3				76.2 (0.00)	27.8 (0.01)	3.82 (0.91)
4					14.7 (0.04)	0.34 (0.56)
	Poland					
0	657.6 (0.00)	419.2 (0.00)	259.5 (0.00)	163.3 (0.00)	115.8 (0.00)	94.2 (0.00)
1		411.6 (0.00)	237.0 (0.00)	113.7 (0.00)	54.1 (0.12)	49.1 (0.04)
2			196.1 (0.00)	73.6 (0.00)	25.0 (0.61)	20.7 (0.39)
3				62.6 (0.00)	12.6 (0.61)	6.61 (0.63)
4					21.7 (0.00)	0.00 (0.96)
	Slovenia					
0	383.5 (0.00)	290.9 (0.00)	215.4 (0.00)	164.4 (0.00)	127.9 (0.00)	103.2 (0.00)
1		234.3 (0.00)	155.8 (0.00)	102.6 (0.00)	67.9 (0.01)	49.3 (0.03)
2			142.9 (0.00)	65.6 (0.00)	37.3 (0.07)	16.7 (0.67)
3				94.3 (0.00)	28.1 (0.01)	5.55 (0.75)
4					17.3 (0.02)	0.48 (0.49)
p-r-s	5	4	3	2	1	0

* *Corresponding p-values in brackets.*

[16] s is the number of I(1) components.

[17] There are also some signs of two I(2) components for the Czech Republic; however, we choose one component as we believe that the system of variables considered in this section is best described by a single I(2) stochastic trend.

Table 3 reports the estimates of $\alpha_{\perp 2}$ and $\beta_{\perp 2}$ vectors.[18] Estimation of an I(2) model is based on the 2-step procedure proposed by Johansen (1995b). Whenever statistically supported (see test statistics below Table 3), α and β matrices obtained in the first step enter the second step restricted.[19]

One common feature of the results is that output and nominal wages do not contribute to the I(2) stochastic trend, which can thus be assumed to be determined only by the shock to the three remaining variables: the nominal exchange rate, CPI and PPI. The share of nominal exchange rate in the nominal stochastic trend is the highest in Slovenia, approximately twice as large as the corresponding shares of the CPI and PPI, which are roughly equal. This implies that the main inflationary pressures in Slovenia come from shocks to the nominal exchange rate, and much less from the autonomous pricing behavior of imperfectly competitive firms. Exactly opposite is the case of Poland, where the share of the nominal exchange rate is almost negligible, whereas the I(2) trend can be attributed to shocks to the CPI and PPI in roughly equal proportions.

For the Czech Republic all three variables seem to contribute to inflationary movements in the economy; however, the share of the nominal exchange rate is considerably smaller than the shares of two price indexes. From the two, the share of PPI is higher. For Hungary the situation is different in the sense that no inflationary pressures come from shocks to the PPI. They originate in shocks to the exchange rate, and more importantly, from shocks to the CPI. As one of the major differences from the PPI and the CPI is that the latter also reflect prices of non-tradable goods, we could infer that in Hungary an important share of inflationary pressures comes from the non-tradable sector. This could arise from a combination of monopolistic pricing, wage pressure and administrative price changes in non-tradable sectors.

[18] The estimates have been obtained with computer code written by Clara Jorgensen for CATS in RATS.
[19] Note that a row of zeros in the α matrix in I(2) context does not necessarily imply that the corresponding variable is weakly exogenous an thus does not necessarily represent one of the common trends in the model. Two additional condition for weak exogeneity in I(2) systems have to be tested (Paruolo and Rahbek (1999)), which was left for future extensions of the paper.

Table 3: Second Order Stochastic Trends and Corresponding Loadings Coefficients

	y	*e*	*cpi*	*ppi*	*w*
			The Czech Republic[a]		
$\alpha'_{\perp 2}$	-0.003	-0.013	0.019	-0.025	-0.002
$\beta'_{\perp 2}$	-0.675	1.663	1.929	1.357	2.667
			Hungary[b]		
$\alpha'_{\perp 2}$	0.000	0.017	-0.031	0.002	0.003
$\beta'_{\perp 2}$	0.222	-1.966	-1.374	-1.557	-0.816
			Poland		
$\alpha'_{\perp 2}$	0.0006	0.0003	-0.010	0.010	0.004
$\beta'_{\perp 2}$	0.954	1.826	1.206	1.210	1.575
			Slovenia[d]		
$\alpha'_{\perp 2}$	-0.010	-0.041	-0.024	-0.021	-0.010
$\beta'_{\perp 2}$	-0.474	-1.387	-1.592	-1.465	-2.163

[a] First two elements of α vector restricted to 0, $\chi^2(2) = 0.28$ p-val.=0.87.

[b] First and the last element of α vector restricted to 0, $\chi^2(2) = 1.48$ p-val.=0.48.

[c] First and the last element of α matrix, and third element of β restricted to 0, $\chi^2(3) = 0.44$ p- val.=0.93.

[d] Second row of α matrix and fourth row of β matrix restricted to zero, and linear homogeneity of the CPI index and nominal wages imposed, $\chi^2(5) = 10.08$ p-val.=0.07.

Examination of the $\beta_{\perp 2}$ vectors, effectively measuring the loadings to second order stochastic trends, yields a final and sufficient condition for identification of candidate I(2) variables in our system. A common feature in Table 3 is that the loading coefficient of the I(2) trend into output is considerably smaller than other coefficients. Formal tests (not presented here) cannot reject the null of these coefficients being zero. This definitely qualifies the index of industrial production as being integrated of order one in all countries. Based on this finding the IIP enters in levels to models in the next section, where we consider I(1) systems.

The second common feature is that nominal wages (with the exception of Hungary) respond very strongly to driving I(2) trend. This implies that nominal wages adjust strongly to price developments in order to achieve dynamic adjustment of equilibrium real wages. However, the loading to nominal wages cannot be treated as equal to the loadings of any of the two price indexes. Thus, as nominal stochastic trends load disproportionately to wages and prices, real wages are also potentially I(2). The usefulness of real wages as a variable in I(1) analysis is, from this point of view, seriously questioned.

There is evidence that the ratio of CPI to PPI cointegrates down to I(1). Under the assumption that all products in the PPI are tradable this measure roughly corresponds to the ratio of non-tradable versus tradable prices. However, such an assumption is indeed very rough and the proposed "nominal to real" reduction would be valuable if we really disposed with tradable and non-tradable price indexes for all countries. Furthermore, using the ratio of prices would not enable us to identify the pass-through effect. For this reason in I(1) analysis we use only one price index, namely the CPI, in differenced form, *i.e.* inflation rate.

It is true that simple first-differencing of variables to reduce the system from I(2) to I(1) does not preserve original cointegrating relations from the I(2) system, and thus cannot be used as a valid reduction from I(2) to I(1). However, the reductions that are supported by the I(2) cointegration analysis are not always economically meaningful (see Kongsted (2002) for exposition of tests of nominal-to-real reductions), which is also the case for our analysis that is oriented towards identification of the pass-through effect. As a result, we must consider an I(1) system that does not fully reflect the cointegrating properties of the data when modeled as I(2), but it nevertheless yields economically meaningful results, while still taking into account the fact that prices and nominal exchange rate are integrated of order two.

To summarize, in this section we have established that the only appropriate approach to the analysis of price movements is to treat them as variables integrated of order 2. In other words, inflation rate results to be nonstationary in all four countries. I(1) analysis should therefore operate directly with inflation rates as there is, in addition, no other economically meaningful transformation supported by I(2) analysis. The same line of reasoning also holds for the nominal exchange rate, which again resulted as an I(2) variable.

The second important finding of I(2) analysis is the identification of the relative importance of shocks to different variables in I(2) trends. A central question of this paper is how different exchange rate regimes influence the overall inflationary performance of an economy. Our priors were that a regime that systematically depreciates the domestic currency leads to firms strongly incorporating expected depreciations into their pricing behavior. As a result, exchange rate policy becomes an important source of inflationary pressures and leads to an average inflation rate considerably above the one corresponding to structural dynamics of the economy. In this respect the share of exchange rate shocks in the nominal stochastic trend is the highest (and moreover dominant) in Slovenia, followed by Hungary, the Czech Republic and Poland.[20]

3.4 I(1) Analysis of Inflation Rate Differential With Respect to Germany

In this section we consider an I(1) system that enables us to identify the long-run equilibrium relation between growth of the nominal exchange rate growth and inflation differential with respect to Germany. In particular we analyze the following system:

$$X_t = \left(y_t, \Delta e_t, \pi_t - \pi_t^*, i_t - i_t^*\right)$$

y_t - index of total industrial production,

Δe_t - growth of nominal exchange rate,

$\pi_t - \pi_t^*$ - inflation differential with respect to Germany, and

$i_t - i_t^*$ - nominal interest rate differential with respect to Fibor/Euribor (3 month).

We have chosen to use the level of industrial production index as it does not exhibit signs of I(2)-ness. The nominal exchange rate has to be differenced, however, in order to rule out I(2)-ness with certainty. Domestic and foreign inflation rate enter as a homogeneous relation, because, firstly, the relation between nominal exchange rate growth and inflation differential is what we are primarily interested in, and, secondly, only in this way it is possible to circumvent the identification problems addressed in the previous section. By analogy the nominal interest rates enter the system also as a spread.

The lag length of each system has been chosen in the same manner as for the systems in the previous section. It proved sufficient to include two endogenous lags for Slovenia, and three for Hungary, the Czech Republic and Poland.[21] Tests for model misspecification are presented in Table A2 in the

[20] The results for the latter country should be taken with some reservation, however, as its VAR model was not statistically completely satisfactory and the consequences of excess kurtosis in I(2) cointegration analysis are theoretically not yet explored.

[21] As a robustness check we have estimated the systems also with 4, 6 and 8 lags. 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

Appendix. Again 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 again 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.[22]

Our choice of cointegration rank is 3, uniformly across all four countries. The trace test indicates this very clearly for Slovenia and Hungary, while rank 2 is also possible for the Czech Republic and Poland. We have nevertheless chosen rank 3 also for these two countries 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 A1 – A4 also show no obvious signs of non-stationarity for the Czech Republic and Poland. Moreover, for all countries we find strong and significant adjustment to the third cointegrating relation.

The left panel of Table 4 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 second coefficient. The inverse of this coefficient 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 in Table 4) with corresponding p-values above 0.90. For Poland the point estimate of this coefficient is smaller than 1, 0.8 to be precise; 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, below 0.5. 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).

virtually unchanged. The variation in parameter estimates is so small that it does not change the conclusions presented in the paper.

22 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.

Table 4: Estimated Cointegrating Relations and Loading Coefficients

	β_1	β_2	β_3	α_1	α_2	α_3
			The Czech Republic[a] (1993:12 – 2002:7)			
y	**1.00**	0.00	0.00	0.00	0.00	0.00
Δe	0.00	**1.00**	0.00	**-62.06**	**-1.03**	**-1.28**
$\pi - \pi^*$	0.00	**-2.17**	**1.00**	**-20.83**	0.02	**-0.79**
i-i^*	**0.03**	0.00	**-0.78**	-0.68	-0.00	0.02
			Hungary[b] (1993:2 – 2002:7)			
y	**1.00**	-	-	**-0.09**	0.00	0.00
Δe	-	**1.00**	-	-4.87	**-0.88**	**-0.98**
$\pi - \pi^*$	-	**-1.03**	**1.00**	2.62	0.01	**-0.73**
i-i^*	**0.03**	-	**-0.67**	**-1.10**	0.00	0.01
			Poland[c] (1993:1 – 2002:4)			
y	**1.00**	0.00	0.00	-0.022	0.00	**0.001**
Δe	0.00	**1.00**	0.00	**-59.79**	**-1.23**	**-1.02**
$\pi - \pi^*$	0.00	**-1.16**	**1.00**	**-25.37**	-0.01	**-0.78**
i-i^*	**0.03**	0.00	**-1.19**	-0.64	-0.00	0.01
			Slovenia[d] (1993:3 – 2002:3)			
y	**1.00**	0.00	0.00	**0.03**	-0.00	-0.00
Δe	0.00	**1.00**	0.00	**-26.49**	**-0.62**	**-0.78**
$\pi - \pi^*$	0.00	**-0.99**	**1.00**	-2.66	0.02	**-0.75**
i-i^*	**0.01**	0.00	**-0.43**	**-18.51**	0.00	**0.38**

Bold face indicates significance.

[a] Weak exogeneity of output: $\chi^2(3) = 3.03$, p-val.=0.39, Weak exogeneity of interest rate differential: $\chi^2(3) = 9.42$, p-val.=0.02. Weak exogeneity of output + H_0: $\beta_{22} = -2$; $\chi^2(4) = 8.07$, p-val.=0.09.

[b] H_0: $\beta_{22} = -1$; $\chi^2(1) = 0.01$, p-val.=0.92

[c] H_0: $\beta_{22} = -1$; $\chi^2(1) = 0.24$, p-val.=0.61, H_0: $\beta_{33} = -1$; $\chi^2(1) = 0.76$, p-val.=0.38

[d] H_0: $\beta_{22} = -1$; $\chi^2(1) = 0.00$, p-val.=0.99

The interpretation of cointegration coefficients is most informative if we look at the orthogonal complement to $\beta = (\beta_1, \beta_2, \beta_3)$. For each country this is a *4x1* vector, which we denote by $\beta_\perp$ and report below in Table 5. This vector shows admissible long-run co-movements of the variables analyzed. Since the estimates for all countries are quite similar, it also holds that orthogonal complements to cointegrating space show qualitatively the same structure. Inspection of the orthogonal complement to cointegrating space here is necessary if results of our analysis are to be used for policy analysis. Technically this follows from the fact that interpretation of coefficients in a cointegrating relation does not obey a simple ceteris paribus logic, but it needs to account also for equilibrium adjustments of other variables in the system (see Johansen (2002) for a detailed elaboration).

Table 5: Orthogonal complements to cointegrating space - $\beta_{\perp}^{'}$

	y	Δe	$\pi - \pi *$	*i-i**
Czech Republic	-0.018	1	0.5	0.625
Hungary	-0.045	1	1	1.5
Poland	-0.020	1	0.8	0.67
Slovenia	-0.025	1	1	2.5

For the case of Slovenia $\beta_{\perp}$ is *(-0.025, 1,1,2.5)'*. This means that an equilibrium or permanent change in exchange rate growth will necessarily lead to an equivalent increase in inflation differential, a disproportionate increase in interest rate spread, and consequently a lower level of output. Because the rise 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.[23] 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 indices *i.e.* a vector of type *(-0.025, 1,* $\lambda \neq 1$*,2.5)* is not orthogonal to β. This means that we can indeed interpret the inverse of second coefficient in the second cointegrating relation as a measure of pass-through into CPI inflation. Moreover, any permanent rise in the rate of depreciation results in a rise in real interest rate differential and a negative effect on output.

Hungary is in this respect again very similar 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

[23] To see this, note that we can rewrite the third cointegrating relation generically denoted as $\pi - \pi^{*} = \beta_{33}(i - i^{*})$ as $r - r^{*} = (1 - \beta_{33})(i - i^{*})$, where r denotes the real interest rate.

output. Poland and the Czech Republic share many similarities and appear slightly 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. Again this fits with our priors about the nature of exchange rate policy in these two countries.

To complete the exposition of I(1) analysis it is important also to look at the corresponding α coefficients, 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, which represent open-economy parity relations. This strengthens the interpretation of the inverse of β_{22} coefficient as the measure of pass-through effect. It implies that after an exchange rate shock to this relation, output almost does not adjust in equilibrium, and in the interpretation of β_{22} coefficient this allows us to abstract from output movements that cause trend movements in the real exchange rate, namely appreciation.

It is perhaps a bit confusing 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, the coefficients in Table 5 imply 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), which again supports our uniform choice of rank 3. Therefore, shocks to the second cointegrating relation alone cannot be interpreted as exchange rate shocks induced by monetary policy because they are not matched by a contemporaneous increase in interest rate spread and hence cannot be expected to have an effect on inflation *per se*. The fact that inflation does not adjust to the shocks to the second equilibrium relation is consistent with a monetary authority that aims at stabilizing the short-run variations in the real exchange rate. Private agents are likely to anticipate such a policy and incorporate it into their pricing decisions. Given that the policy rule can only gradually eliminate deviations of the real exchange rate from its equilibrium level, such a policy tends to slow down the overall adjustment in the economy.[24]

It clearly emerges from Table 4 that inflation differentials are not weakly exogenous for the long-run parameters of the systems. This is a common observation in the literature, so the result is expected. It nevertheless has an important economic implication. It means that domestic inflation can be controlled, i.e. made stationary around a selected target value, by selecting an appropriate path of policy variables (exchange rate and interest rate) (Johansen and Juselius, 2001). We have demonstrated that in the economies the exchange rate channel in the transmission mechanism is very important. This implies that an appropriate choice of exchange rate regime can contribute decisively to overall inflationary performance.

Empirical results can be summarized as follows. A higher growth rate of nominal exchange rate results in equally higher difference between domestic and foreign inflation in Slovenia and Hungary. Moreover, from I(2) analysis for Slovenia it follows that innovations to the exchange rate are the most important source of inflationary pressures. In Hungary, on the other hand, exchange rate innovations are comparatively less important. The point estimate for Poland shows a coefficient between exchange rate growth and inflation differential that is smaller than one, but not significantly different. Nevertheless, we tentatively conclude that the effect of the exchange rate growth on inflation is smaller than in Slovenia and Hungary. Again, the analysis of I(2) nominal trend corroborates this conclusion. The country with the lowest effect of exchange rate on prices is the Czech Republic. This is also in line with the I(2) analysis, where we see that innovations of both price indexes are more important components of the I(2) nominal stochastic trend.

4. Conclusions

Despite the variety of approaches to the exchange rate policy, CEEC-4 have all made substantial progress in reducing inflation, which has been on average below 10 percent since 1998. Part of the explanation for inflation rates that are still higher than in the EU could be found in the working of the

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[24]See Uribe (2003) for a discussion of the relationship between the real exchange rate target and dynamic adjustment of an economy with forward-looking agents.

Balassa-Samuelson effect and the process of overall relative price convergence in CEEC-4 on average. However, we argue that the combination of exchange rate regime and monetary policy contribute to the differences in inflation rates among CEEC-4. The paper finds a strong pass-through from nominal exchange rates to domestic inflation. In such a context, the dichotomy between inflation targeting and exchange rate targeting 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.

In the last three years, inflation rates have been lower in the Czech Republic and Poland than in Hungary and Slovenia. As the Czech Republic and Poland maintain relatively less managed exchange rate regimes than Hungary and Slovenia, and additionally employ inflation targets, it is believed that such a combination produces lower inflation. However, in the case of Poland the costs in terms of output and unemployment appear very large. In the case of the Czech Republic it appears that the exchange rate features as a main intermediate target to achieve the final target on inflation, as is natural in a small open economy. Before adopting the euro, all candidate countries will have to enter the ERM2 system with an agreed central parity and a ±15 % band. It is argued in the paper that the pre-adoption period may generate persistent inflationary pressure, as candidate countries will probably try to maintain external competitiveness and use exchange rate as a shock abosorber. One can thus expect rising interest rates and output volatility in ERM2 prior to actual adoption of the euro. Such volatility will be affected by the regime of full capital mobility that the countries have to adopt upon entry in the European Union. Results in the paper suggest that the best policy for CEEC-4 would be the adoption of the euro as early as possible. Before actual adoption, a pre-announced path of moderate depreciation (crawling peg) might be the second-best option for exchange rate policy.

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STATISTICAL APPENDIX

Table A1: Misspecification tests and characteristic roots (multivariate tests) for the system $X_t = (y_t, e_t, cpi_t, ppi_t, w_t)'$

The Czech Republic (1993:12 – 2002:7)						
Res. autocorr. 1-7	F(175,188) = 0.93			p-val.=0.69		
Normality	$\chi^2(10) = 13.42$			p-val.=0.20		
Heteroscedasticity	F(480,455) = 0.93			p-val.=1.00		
	Modulus of 6 largest characteristic roots					
Unrestricted VAR	1.02	0.99	0.91	0.91	0.86	0.86
r=1	1.00	1.00	1.00	1.00	0.96	0.86
Hungary (1993:2 – 2002:7)						
Res. autocorr. 1-7	F(150,104) = 1.02			p-val.=0.45		
Normality	$\chi^2(10) = 13.17$			p-val.=0.21		
Heteroscedasticity	$\chi^2(780) = 809.24$			p-val.=0.22		
	Modulus of 6 largest characteristic roots					
Unrestricted VAR	0.99	0.98	0.93	0.86	0.85	0.85
r=1	1.00	1.00	1.00	1.00	0.98	0.86
Poland (1993:1 – 2002:4)						
Res. autocorr. 1-7	F(175,124) = 1.11			p-val.=0.27		
Normality	$\chi^2(10) = 32.62$			p-val.=0.00		
Heteroscedasticity	F(480,278) = 0.36			p-val.=1.00		
	Modulus of 6 largest characteristic roots					
Unrestricted VAR	0.98	0.96	0.96	0.80	0.79	0.79
r=1	1.00	1.00	1.00	1.00	0.95	0.77
Slovenia (1993:3 – 2002:3)						
Res. autocorr. 1-7	F(175,163) = 1.14			p-val.=0.20		
Normality	$\chi^2(10) = 9.13$			p-val.=0.52		
Heteroscedasticity	F(630,265) = 0.33			p-val.=1.00		
	Modulus of 6 largest characteristic roots					
Unrestricted VAR	0.98	0.96	0.96	0.82	0.82	0.75
r=2	1.00	1.00	1.00	0.97	0.86	0.86

Table A2: Misspecification tests (multivariate), characteristic roots and trace tests for the system $X_t = \left(y_t, \Delta e_t, \pi_t - \pi_t^*, i_t - i_t^*\right)$

The Czech Republic (1993:12 – 2002:7)						
Res. autocorr. 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
r=3	1.00	0.91	0.74	0.74	0.63	0.63
Trace test	1.29	12.78	44.90	101.32		
p-value	0.26	0.14	0.00	0.00		
R	3	2	1	0		
Hungary (1993:2 – 2002:7)						
Res. autocorr. 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		
Trace test	Modulus of 6 largest characteristic roots					
Unrestricted VAR	1.01	0.93	0.75	0.75	0.50	0.50
r=3	1.00	0.91	0.75	0.75	0.49	0.49
Trace test	1.25	16.96	56.43	130.61		
p-value	0.26	0.03	0.00	0.00		
R	3	2	1	0		
Poland (1993:1 – 2002:4)						
Res. autocorr. 1-7	F(112,197) = 1.15			p-val.=0.20		
Normality	$\chi^2(8) = 18.86$			p-val.=0.02		
Heteroscedasticity	F(260,428) = 0.79			p-val.=0.98		
	Modulus of 6 largest characteristic roots					
Unrestricted VAR	0.99	0.87	0.67	0.67	0.59	0.59
r=3	1.00	0.89	0.67	0.67	0.59	0.59
Trace test	1.37	7.19	53.44	107.85		
p-value	0.24	0.56	0.00	0.00		
R	3	2	1	0		
Slovenia (1993:3 – 2002:3)						
Res. autocorr. 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
r=3	1.00	0.88	0.41	0.41	0.35	0.35
Trace statistic	0.28	26.97	204.86	406.34		
p-value	0.60	0.00	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 from the authors upon request.

Figure A1: Cointegrating Vectors (unconcentrated) for the Czech Republic

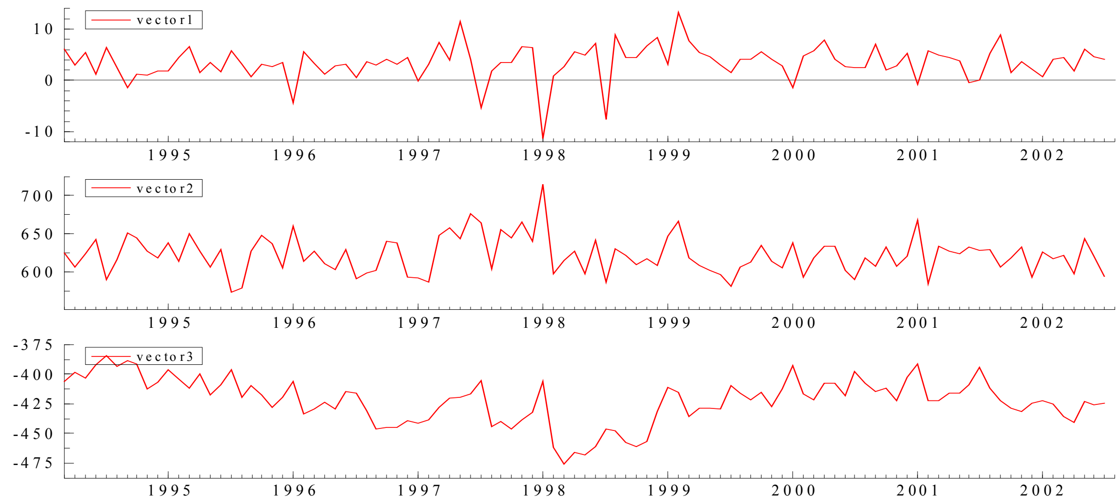

Figure A2: Cointegrating Vectors (unconcentrated) for Hungary

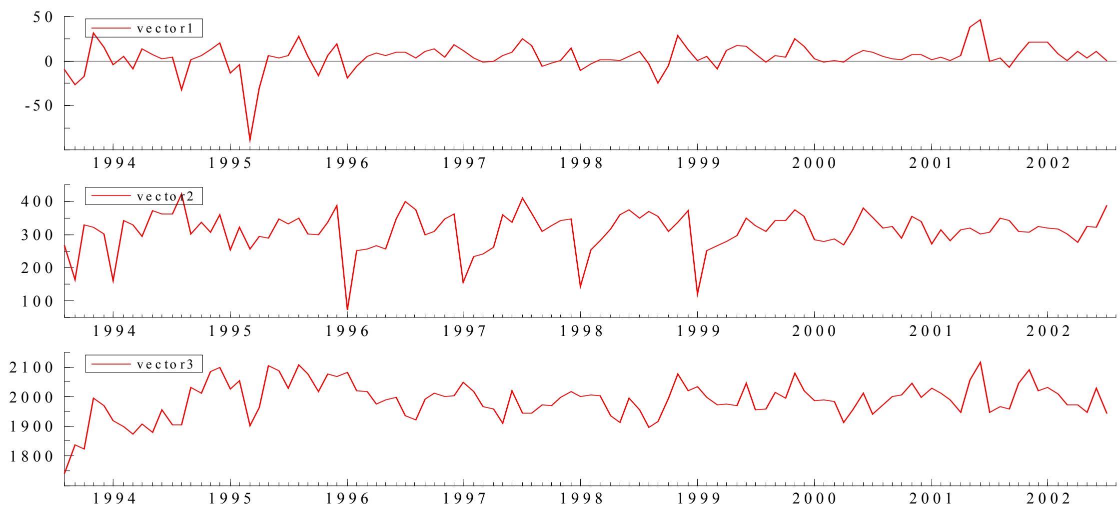

Figure A3: Cointegrating Vectors (unconcentrated) for Poland

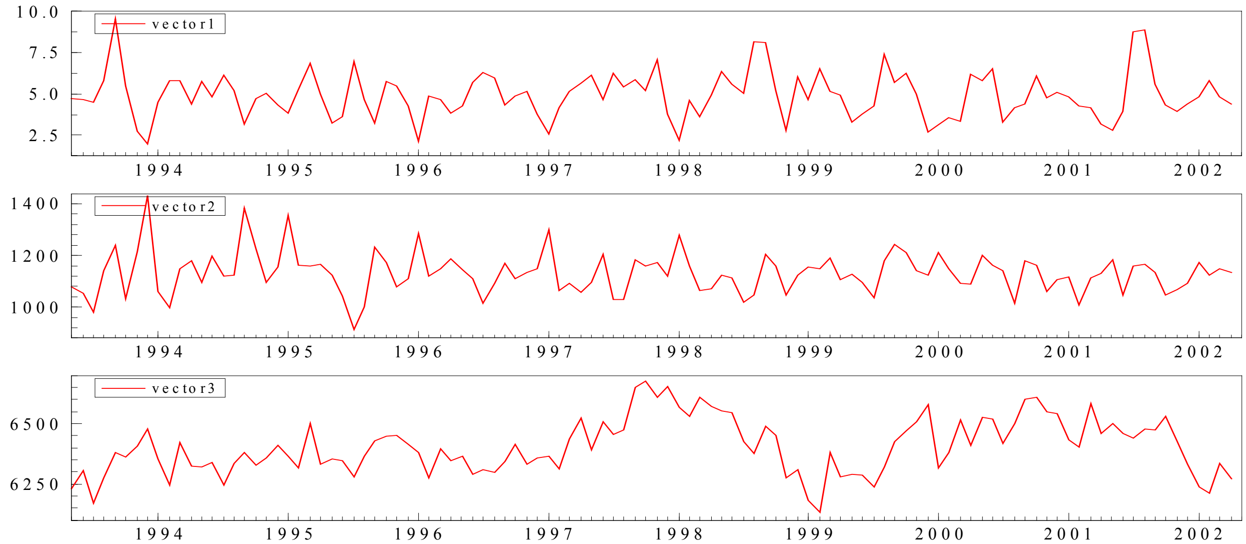

Figure A4: Cointegrating Vectors (unconcentrated) for Slovenia