Exchange Rate Pass-Through in Candidate Countries

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1. Introduction

Following accession to the European Union, candidate transition 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 into 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 month earlier the two-year deadline. There is clearly some discretionality.

Even leaving aside the issue of the interpretation of the specific rules that apply before adoption of the euro, the timing of such adoption is one of the main macroeconomic issues relating to accession. There has been a lively debate on the path towards the euro, focusing on : (i) the pre-conditions for such adoption, within the well-known theory of optimal currency area; (ii) the ability of CEECs to fulfill Maastricht criteria, especially that on inflation, 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 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. This topic has been analyzed in several papers. However, we argue that previous analyses 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 (Czech Republic, Hungary, Poland, and Slovenia). 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, we find that 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 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 had a fundamental impact on domestic inflation. The real exchange rate rule in Slovenia was likely internalized by price setters and thus became a persistent source of inflation. Interestingly, Slovenia that apparently had the best fundamentals of CEEC-4 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 the different degree of openness of the two economies, with Slovenia much more open and much smaller 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 in between the two extreme cases, with Hungary more similar to Slovenia and Poland more to the Czech Republic.

The analysis has also some 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. Large pass-through is likely to induce a response of policy-makers that will attempt ex post to drive the exchange rate in a way that maintains external competitiveness. As in the case of Slovenia such 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 adoption of fixed exchange rates. Luckily, candidate countries have the point of arrival, the euro, already set. Their main policy decision is how fast enter the euro. Results in this paper suggest that there are no significant advantages to delay such an entry. A pre-announced path of moderate depreciation (crawling peg) might be the best option towards the entry in the euro.

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. It is claimed that there is substantial inflationary pressure coming from the non-tradable sector, which cannot be overturn by active exchange rate policy. It is argued in Section 3 that the exchange rate does not play an absorbing role in CEECs. Section 4 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 Czech Republic and Poland. While Slovenia and Hungary engaged in relatively tightly managed exchange rates, Czech Republic and Poland let their exchange rate float more freely, at least recently. Additionally, Czech Republic and Poland introduced inflation targets, which helped monetary authorities to maintain inflation at lower levels than in Slovenia and Hungary. However, the decline in inflation in Poland took place in a period of sharp slowdown of the economy. Only the Czech Republic has been able to credibly follow a policy of successfully targeting inflation. The much lower external debt of the Czech Republic, compared to Poland, likely contributed to make more credible a policy of targeting inflation and flexible exchange rates. We conjecture that a more predictable exchange rate policies, as those followed in Slovenia and Hungary (and

Poland until 2000) tends to be associated with larger pass-through coefficients. Size and openness of the countries are also important factors. Section 5 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 (Table 1).

Table 1: Inflation Rates in CEEC-4

	1990	1991	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001
Czech Republic	10,8	56,7	11,2	20,8	10,0	9,1	8,8	8,5	10,7	2,1	3,9	4,7
Hungary	28,9	35,0	23,0	22,5	18,8	28,2	23,6	18,3	14,3	10,0	9,8	9,2
Poland	585,8	70,3	43,0	35,3	32,2	27,8	19,9	14,9	11,8	7,3	10,1	5,5
Slovenia	549,7	117,7	207,3	32,9	21	13,5	9,9	8,4	8,0	6,1	8,9	8,4

Source: EBRD 2002.

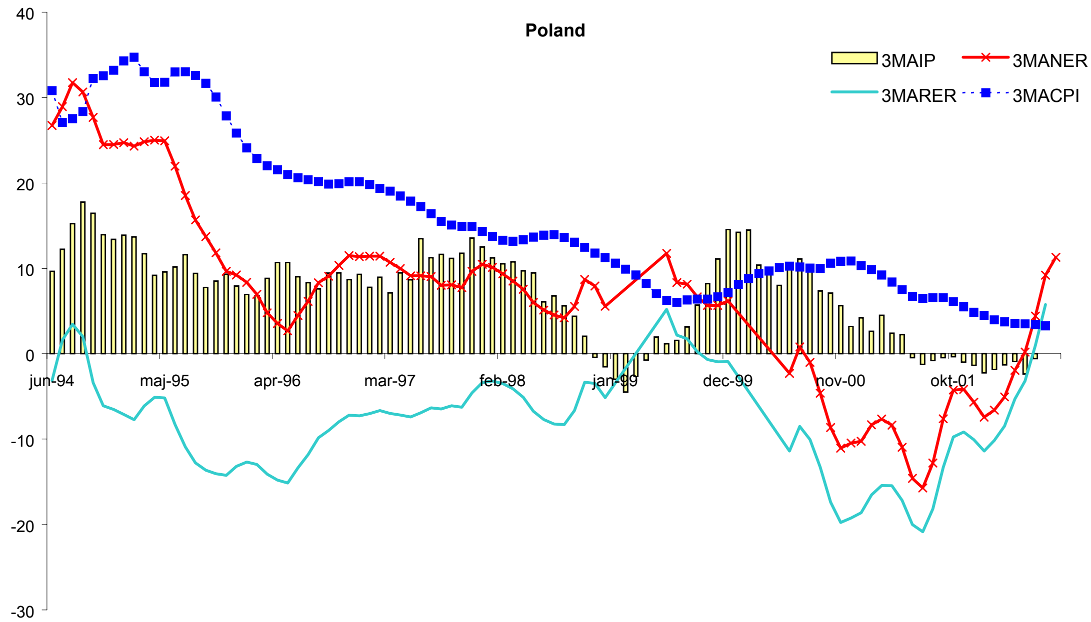
The reduction to single-digit inflation was much faster in Slovenia and the Czech Republic, countries less affected by large stocks of debt and the attendant need to finance large debt service payments. Moreover, inflation rates seem to be more stubborn in Slovenia and Hungary than in the Czech Republic 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 to reduce 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 desired decline. Such overshooting resulted from an excessively tight monetary policy that negatively affected the economy during a period of generalized slowdown in Europe. The output performance in Poland during 2002 has been among the worst in candidate countries. 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. All CEEC-4 experienced such appreciation, which has been common to all transition economies. A component of such trend appreciation can be considered an equilibrium phenomenon, in line with the so-called Balassa-Samuelson effect, that affects real exchange rates in a phase of catching-up. However, in addition to such a trend appreciation there is a shorter term dynamic process connecting exchange rates and inflation. Figure 1 illustrates the movements of inflation nominal and real exchange rates in CEEC-4 since 1995. The figure contains as well an indicator of the state of the economy, proxied by industrial production. [1]

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

[1] Data on industrial production (MAIP), real exchange rate (MARER), nominal exchange rate (MANER), and consumer price index (MACPI) are 3-month moving averages.

Figure 1. Nominal and Real Exchange Rates in CEEC-4



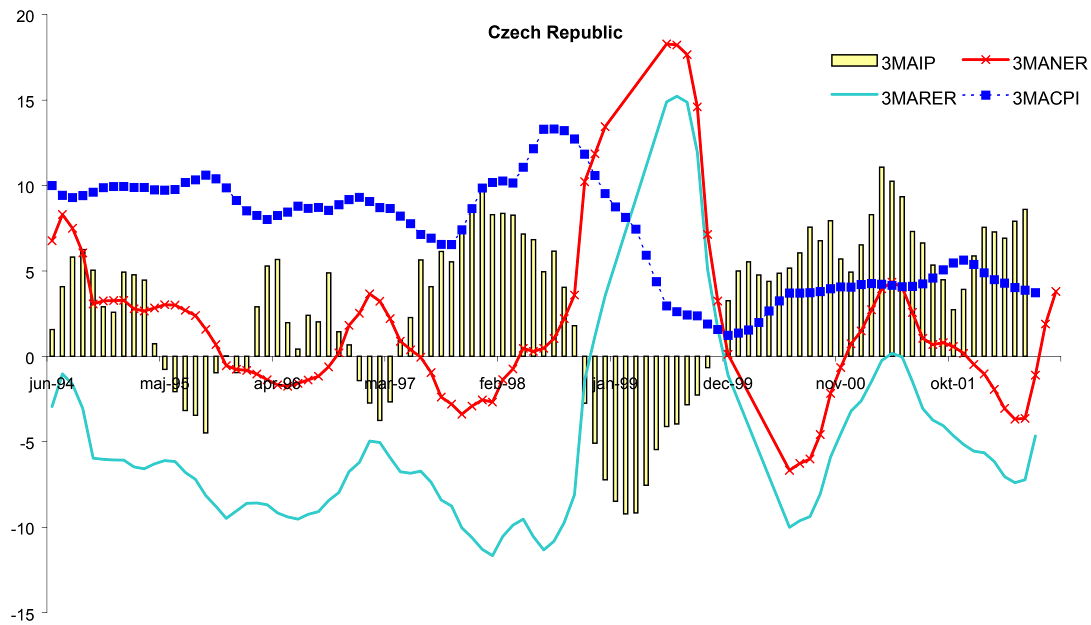
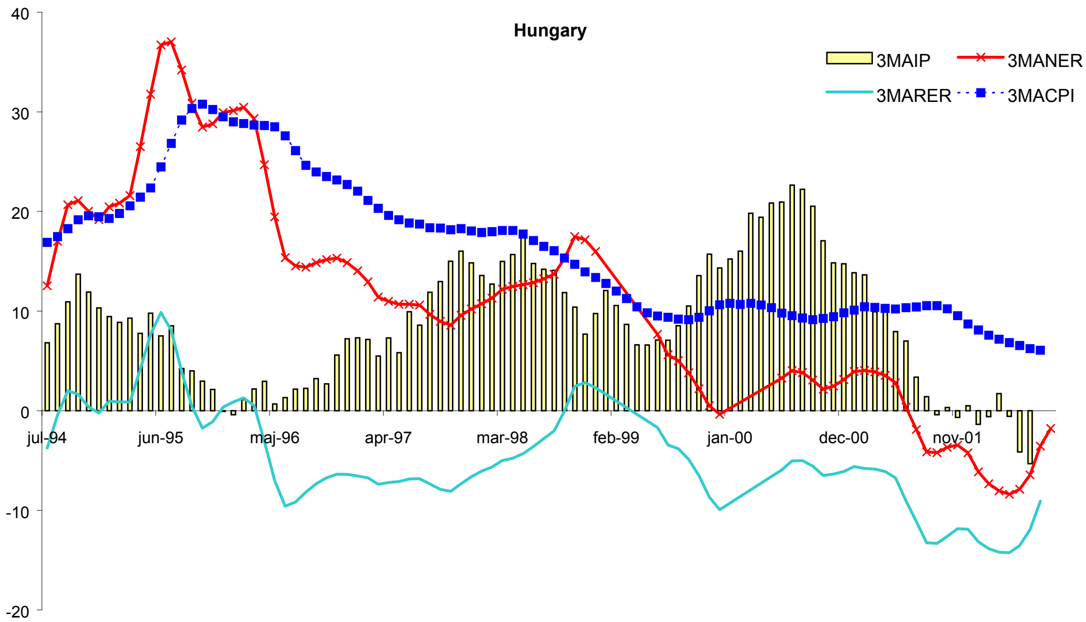
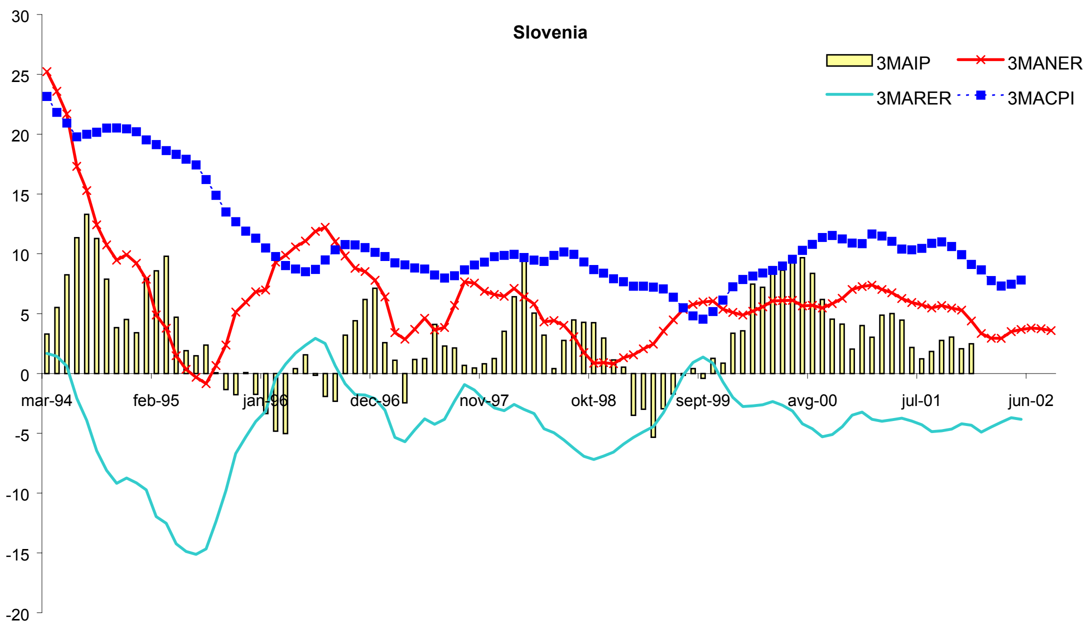


Figure 1: continued





Source: Datastream.

In addition to the above considerations, the different patterns of inflation dynamics in the four candidate countries seem to be associated to different exchange rate regimes. Table 2 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 2: 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). Most CEEC-4 liberalized long-term capital flows, while some controls on short-term flows are still in place in Hungary, Poland, and Slovenia. In addition to different exchange rate regimes and liberalization of capital flows, CEEC-4 employed different monetary policies. While Czech Republic and Poland set inflation targets, Hungary and Slovenia stuck to the exchange rate and M3 targets, respectively. According to EBRD transition indicators, CEEC-4 represents a homogenous group of the most advanced transition economies. Table 3 presents the evolution of selected EBRD transition indicators on price and trade liberalization. Both indexes of price and foreign exchange and trade liberalization, respectively, improved around 1996 and 1997. As shown in Coricelli and Jazbec (2001), this improvement corresponds to a diminishing effect of structural reforms on the real exchange rate determination in transition economies. In the middle of the 90's, CEEC-4 were on average in the fifth or sixth year of the transition process when with respect to the behavior of real exchange rate productivity and demand factors began to affect the real exchange rate more than structural reforms. In the same period, exchange rate regimes switched from less to more flexible framework with respect to the regimes employed at the beginning of 90's.

The transition indicators score from 1 to 4 with a 0.3 decimal points added or subtracted for improvements or declines in ratings. While 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 prices liberalization, especially in the public sector and non-market prices. The shift in exchange rate regimes in Czech Republic and Hungary broadly corresponds to the liberalization of trade and current account convertibility, while in Poland shift toward free floating happened only in 2001. Although Slovenia officially targeted M3 throughout the last decade, tightly managed exchange rate regime was substantially supported by capital controls on short-term capital flows and 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 been on average below 10 percent since 1998. This points to the fact that it is the consistency of country's entire package of economic policy that matters for the macroeconomic performance rather than the exchange rate regime per se. Although the anti-inflationary programs in CEE4 countries 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 to entry the EU, and consequently EMU. As already mentioned, part of the reasons for higher inflation rates could be founded 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 exchange rate regime and monetary policy could substantially contribute to the differences in inflation rates in CEEC-4 as Czech Republic and Poland on average produced lower inflation rates than Hungary and Slovenia in the last three years. As Czech Republic and Poland maintains relatively less managed exchange rate regimes than Hungary and Slovenia, and additionally employ inflation targets, it is believed that the combination of relatively greater flexibility of exchange rate regime and inflation target produces lower inflation.

Table 3: EBRD Transition Indicators for CEEC-4

	1992	1993	1994	1995	1996	1997	1998	1999	2000
Index of Price Liberalization									
Czech Republic	3,0	3,0	3,0	3,0	3,0	3,0	3,0	3,0	3,0
Hungary	3,0	3,0	3,0	3,0	3,0	3,3	3,3	3,3	3,3
Poland	3,0	3,0	3,0	3,0	3,0	3,0	3,3	3,3	3,3
Slovenia	3,0	3,0	3,0	3,0	3,0	3,3	3,3	3,3	3,3
Index of Forex and Trade Liberalization									
Czech Republic	4,0	4,0	4,0	4,0	4,3	4,3	4,3	4,3	4,3
Hungary	4,0	4,0	4,3	4,3	4,3	4,3	4,3	4,3	4,3
Poland	3,0	4,0	4,0	4,0	4,3	4,3	4,3	4,3	4,3
Slovenia	3,0	4,0	4,0	4,0	4,3	4,3	4,3	4,3	4,3

Source: EBRD 2001.

The switch of exchange rate regimes broadly corresponds to the fading of structural reform effects, surge of capital inflows, and further price liberalization. It is commonly believed that more active exchange rate policy could on one hand circumvent the pressing problem of real exchange rate appreciation, and on the other act as a shock absorber against asymmetric real shocks. As discussed in the next section both arguments used to support the switch of exchange rate regime cannot justify greater flexibility of exchange rate regime in CEEC-4.

3. Exchange Rate as Shock Absorber

One of the key arguments for flexibility is that it potentially enables a country to neutralize the adverse effects of asymmetric real shocks. The shocks that are of primary interest in this context are the asymmetric real shock, because asymmetric nominal shocks are effectively removed in a monetary union. In this reasoning, however, lies an important caveat. As Masten (2002) argues, among the asymmetric real shocks we need to distinguish between permanent and transitory shocks. The former are a sign of divergent economic developments (caused by structural changes and catching-up process), and the stochastic trend they induce into the real exchange rate can be seen as an equilibrium driving process. Moreover, this trend cannot be overturned by monetary policy, which, being constrained by real interest rate parity, would indulge into an over-inflationary policy. If a country is subject to permanent asymmetric real shocks, then a monetary union will impose real costs not due to the loss of an important stabilizing tool, but because it will exhibit divergent economic developments relative to other members of the union. The critical point here is that centrally managed monetary policy in such cases amplifies the divergences. The ECB, targeting Euro-area wide aggregates, may in such circumstances act procyclically on the economy of the divergent country. The shock-absorbing role of the real exchange rate should therefore be considered only for the case of transitory real asymmetric shocks.

The shock-absorbing role of the real exchange rate is traditionally analyzed within a structural VAR framework. An important drawback of this approach is that it neglects the possiblity of cointegration and thus it is unable to distinguish permanent from transitory shocks. Namely, in a cointegrated VAR with dimension p and cointegrating rank r we can apply the Granger Representation Theorem to decompose the system into r stationary and hence transitory components, and p-r nonstationary that have permanent effects on variables of interest. This decomposition is based on identified contegrating relation and hence supported by the data. Usual non-testable identifying restrictions that need to be motivated by economic theory can then be imposed separately on the transitory and permanent part of the system in order to identify permanent and transitory structural shocks. In a traditional SVAR exercise, however, variables are conformably differenced to achieve I(0) properties, at the expense of losing the posiblity to perform the permanent-transitory decomposition. The identified strucutral shocks are for this reason a linear combination of firs-differenced permanent shocks but also first-differenced transitory shocks, such that it is impossible to discriminate between the two. Masten (2002) makes a step further in this direction and estimates the common trends model (proposed by King, Stock Plosser and Watson (1991)) in order to discriminate between permanent and transitory shocks. Using the stricter definition for shock-absorbtion – shock-absorbing role considered only for the case of asymmetric and transitory real shocks – no shock-absorbing role of the real exchange rate is found in the Czech Republic, Hungary, Slovenia, Denmark and the United Kingdom. This is the first study of shock-absorbing role of the real exchange rate that considers also some of the Accesion Countries, but its findings are in line with the findings of other authors for different sets of European countries. Very limited shock-absorbing role of the real exchange rate is, for example, found by Canzoneri et al. (1996) for Autria, the Netherlands, France, Italy, Spain and the UK; and then by Thomas (1997) and Funke (2000) for the UK and Sweden respectively. The analysis by Artis and Ehrman (2000) that distinguishes between symmetric and asymmetric shocks also comes to a similar conclusion for Denmark, Sweden, Canada and the UK.

An interesting finding in Masten (2002) is that it identifies significant divergent economic movements in the three ACs that are likely to persist in the future. The latter finding complements the findings of Coricelli and Jazbec (2001), using a different methodology. From this point of view, the benefits of exchange rate flexibility seem very limited and the costs of higher inflation due to pass-through effect should deserve a more important consideration.

4. 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 from pressure caused by a surge in foreign capital inflows (Corker et al., 2000). In so doing, CEEC-4 added a potential new source to higher inflation rates in addition to the working of the Balassa-Samuelson effect. The evidence on selected transition economies could partially support this kind of argument although the extent of the pass-through cannot be firmly established (Darvas, 2001; Campa and Goldberg, 2002). With all the caveats, Darvas (2001) finds short-run estimates of pass-through of nominal exchange rate to fundamental prices (food, energy, and administered prices were excluded from CPI) in 2000 higher in Hungary and Slovenia than in Poland and the Czech Republic. He tentatively concludes that part of the difference in the pass-through estimates could be attributed to the exchange rate regime, as Hungary and Slovenia had a managed exchange rate regime opposed to Poland and the Czech Republic, which had a floating regime in 2000. Although Darvas (2001) takes into account the change of the exchange rate regime in Hungary, the Czech Republic, and Poland during the transition process, the main concern explaining results for pass-through in transition economies is still the shortness of time series for the exchange rate and prices if one seriously considers the importance of the initial period of the labor reallocation process as explained above. Interestingly, the timing of the change of exchange rate regimes in Hungary, the Czech Republic, and Poland vaguely corresponds to the periods in which the process of structural reforms proxied by labor reallocation settled down.

Although the existence of the Balassa-Samuelson effect and potential exchange rate pass-through could provide an explanation for the real exchange dynamics in CEECs on average, it is in the Baltics where both effects had a rather modest occurrence. On one hand, all Baltic countries have currency boards, which offset the exchange rate pass-through, while on the other hand, it seems that the increase in productivity differential was rather small after the initial labor reallocation. For those reasons, real exchange rate appreciation in the Baltics could mostly be attributed to demand factors. The dynamics of relative wages in Latvia could provide justification for this kind of reasoning. Also, wages in the public sector have been increasing more in the Baltics than in other transition economies, with the exception of Slovenia and Romania (Coricelli and Jazbec, 2002).

As real appreciation in transition economies resulted in higher inflationary pressure rather than nominal appreciation, part of the inflationary pressure could derive from goods and labor market rigidities. For that reason, it is 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.

4.1. Limitations of Previous Empirical Studies

Empirical studies of exchange rate pass-through in transition countries have two major shortcomings. Firstly, they do not perform the analysis within the framework of cointegrated vector autoregression model (CVAR). This means that these studies neglect the intrinsic meaning of equilibrium long-run relationship between the nominal exchange rate and prices identified in a cointegrated VAR. Consequently, they cannot analyze the adjustment to equilibrium relations, which is the most important dynamic aspect of price adjustment, and also distinguish it from short-run adjustment of the system. Secondly, existing studies do not address the possibility of prices, the nominal exchange rate and nominal wages being integrated of order 2, which is a more and more common finding in the literature (see for example Banerjee et al. (2001), Juselius (1999, 2001), Coenen and Vega (2001) and Ericsson et al. (1998). This 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 trend stationary has as a consequence invalid statistical inference.[2] Thus, all results obtained without testing for I(2)-ness in the price level before treating inflation as trend stationary can be seriously questioned.

For these reasons, the analysis of pass-through here is presented in a cointegrated VAR using monthly data on nominal exchange rate and a set of selected price indexes for the Czech Republic, Hungary, Poland and Slovenia. In particular, for every country we consider two systems of variables.

[2] On the other hand it is true that econometric investigation of the pass-through effect on quarterly data in the 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.

First, a system of industrial production index, nominal exchange rate, consumer prices (CPI), producer prices (PPI) and nominal wages in I(2) framwork gives a general insight into the relative importance that shocks to different variables have in the identified nominal stochastic trend of order 2. Secondly, based on the results of I(2) analysis we analyze an I(1) a 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 solve almost completely the problem of identification of what we can interpret as an estimate of pass-through effect.

Empirical investigation of the pass-through effect on aggregated data suffers from an identification problem of pure effects of exchange rate changes on prices. The cointegration approach is very natural as it captures what we are primarily interested in: an equilibrium relation between the nominal exchange rate and prices. However, the 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). In particular, without controlling, for example, for the state of the business cycle, productivity developments, short-run aggregate demand effects, etc., the measure of pass-through in an economy facing real appreciation will necessarily be larger than one, which is 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 of other economic variables, and hence cannot be interpreted as pass-through effect in the usual sense. Moreover, 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. The reductions that are supported by the I(2) cointegration analysis are not always economically meaningful (Kongsted, 2002). As an illustration, consider a system that contains nominal wages and price level. If the I(2) trend would feed proportionally into both variables, the ratio of two variables, which yields an economically meaningful quantity i.e. real wages, could cointegrate down to I(1) and thus could be used in typical I(1) cointegration framework. If the I(2) trend does not feed proportionally into two variables, then the reduction could include for example also the nominal exchange rate (possibly with a weight different than one), but the resulting variable would not make sense from the economic point of view. As a result, one has to 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. This point will be briefly illustrated also for the four countries we analyze here.

4.2 Brief Description of I(2) Model

In this section we give a minimal theoretical exposition of I(2) model that we think is still 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 $\Psi_i = -\sum_{j=i+1}^{k-1} \Gamma_i, i = 1,\dots,k-2$. 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 $s \le p-r$. (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 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 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, 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 obeys a special notation $\theta = \Gamma\overline{\beta}\,\overline{\alpha}'\Gamma + \sum_{i=1}^{k-1} i\Gamma_i$, $\overline{\beta} = \beta(\beta'\beta)^{-1}$ and analogously for $\overline{\alpha}$.

From the moving average 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 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 is this trend 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.[3] The second case would correspond to 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 nontradable and service sector.[4] 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 and thus indicates 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 is the I(2) stochastic trend predominately a nominal one. [5]

It is worth repeating that the aim of the I(2) analysis presented in this section is not to directly estimate the pass-trough effect,[6] but to highlight the relative importance of shocks to different nominal variables in the identified nominal I(2) trends.[7] For this reason we have considered here a broad system of

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[3] 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).

[4] A formal analysis of this notion would need to explicitly distinguish between tradable and nontradable prices. Unfortunately we have those indexes available only for Slovenia.

[5] We choose not to report other results of I(2) analysis 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).

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

[7] 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.

domestic nominal variables, to which also the index of industrial production has been added in order to explore a potential presence of I(2) trends also in this variable.

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

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

y_t - index of total industrial production,[8]

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

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

ppi_t - producer price index, and

w_t - average nominal wages.

4.3 I(2) Analysis of Nominal Stochastic Trends

As a first step in the analysis unrestricted VAR models are fitted for each country. In each system impulse dummies have accounted for outliers. 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 likelihood-based test for a valid lag reduction. 3 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.[9]

Table A1 reports also the roots of the companion matrices for unrestricted VARs and with rank restrictions imposed. High number of large roots in each system is already a sign of I(2) trends. This is

[8] This is the only measure of real output available on monthly frequency.

[9] 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).

confirmed in Table 4 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 $r \leq p$. [10] Based on the results one I(2) trend is the preferred choice for all countries.[11] The ranks chosen are 1 for Hungary and 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 a correct one, as in all cases there is only one very large root left in the system.

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

[11] 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 4: 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.*

Table 5 reports the estimates of $\alpha_{\perp 2}$ and $\beta_{\perp 2}$ vectors.[12] 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 5), α and β matrices obtained in the first step enter the second step restricted. 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.[13]

[12] The estimates have been obtained with computer code written by Clara Jorgensen for CATS in RATS.

[13] 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 5: Second Order Stochastic Trends and Corresponding Loadings Probabilities

	y	*E*	*cpi*	*ppi*	*W*
The Czech Republic[a]					
$\alpha_{\perp 2}$	-0.0032	-0.0127	0.0193	-0.0248	-0.0019
$\beta_{\perp 2}$	-0.6751	1.6631	1.9289	1.3574	2.6669
Hungary[b]					
$\alpha_{\perp 2}$	0.0003	0.0173	-0.0310	0.0016	0.0032
$\beta_{\perp 2}$	0.2223	-1.9661	-1.3737	-1.5566	-0.8164
Poland					
$\alpha_{\perp 2}$	0.0006	0.0003	-0.0098	0.0102	0.0043
$\beta_{\perp 2}$	0.9539	1.8258	1.2064	1.2102	1.5751
Slovenia[d]					
$\alpha_{\perp 2}$	0.0027	0.0104	0.0062	0.0054	-0.0026
$\beta_{\perp 2}$	-0.4740	-1.3874	-1.5923	-1.4651	-2.1627

[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.

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 main inflationary pressures in Slovenia come from shocks to the nominal exchange rate, and much less from autonomous pricing behavior of imperfectly competitive firms. Just the 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 reflect also prices of nontradable goods, we could infer that in Hungary an important share of inflationary pressures comes from the nontradable sector. This could arise from a combination of monopolistic pricing, wage pressure and administrative price changes in nontradable sectors.

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 5 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 inters in the same ways the models in the next section, where we consider I(1) systems.

The second common feature is that nominal wages (with 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 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). This ratio is a measure of relative price index that under assumption that all products in the PPI are tradable roughly corresponds to a ratio of nontradable versus tradable prices. However, such assumption is indeed very rough and the proposed "nominal to real" reduction would be valuable if we really disposed with tradable and nontradable price indexes for all countries. Moreover, using CPI/PPI ratio in I(1) systems did not lead to identification of I(1) nominal trend that could be attributed to exchange rate policy and its effects on prices and output.

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 holds also 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 relative importance of shocks to different variables in I(2) trends. A central question of this paper is how do 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 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.[14]

4.4 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).

[14] 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.

We have chosen to use the levels 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. Test for model misspecification are presented in Table A2 in the 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 we wish to emphasize that the key assumption for the validity and robustness of cointegration analysis is that the residuals be stochastically independent, but this is confirmed by the absence of residual autocorrelation.[15]

[15] 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 6: 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
Δe	-	**1.00**	-	**-62.06**	**-1.03**	**-1.28**
$\pi - \pi *$	-	**-2.17**	**1.00**	**-20.83**	0.02	**-0.79**
i-$i*$	**0.03**	-	**-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**	-	-	-0.01	0.00	**0.001**
Δe	-	**1.00**	-	-36.95	**-0.96**	-0.72
$\pi - \pi *$	-	**-1.25**	**1.00**	**-27.35**	-0.00	**-0.80**
i-$i*$	**0.03**	-	**-1.19**	**-0.85**	-0.00	0.01
	Slovenia[d] (1993:3 – 2002:3)					
Y	**1.00**	-	-	**0.03**	-0.00	-0.00
Δe	-	**1.00**	-	**-26.49**	**-0.62**	**-0.78**
$\pi - \pi *$	-	**-0.99**	**1.00**	-2.66	0.02	**-0.75**
i-$i*$	**0.01**	-	**-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, in addition to first restriction 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: $\beta_{22} = 1$; $\chi^2(1) = 0.36$, p-val.=0.55

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

The cointegration rank chosen 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. Moreover, visual inspection of the estimated third cointegrating vector presented in Figures A1 – A4 show no obvious signs of non-stationarity also for the Czech Republic and Poland.

The left panel of Table 6 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 conditionally 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-based tests reported in Table 6) 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 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).

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 cointegrating relation. 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 (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. It is less surprising if we note that the differential contains also German inflation, which does not respond to inflationary developments and exchange rate policy in the four small economies considered here. However, we need to keep in mind that the dynamics of overall adjustment in proportions presented in Table 6 depends also on short-run adjustment coefficient (again we refer the reader to a detailed exposition in Johansen (2002)). Using this, we can see from α_3 vectors that a very strong equilibrium adjustment of inflation (positively in response to a positive exchange rate shock that increases the interest rate spread) occurs through the third cointegrating relation, which again supports our uniform choice of rank 3.

Empirical results can be summarized as follows. A higher growth rate of nominal exchange rate transfers one to one to the 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 that 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.

5. 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 inflation rates still higher than in EU could be found in the working of Balassa-Samuelson effect and the remaining process of relative price convergence in CEEC-4 on average. However, we argued 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 turn out to be a policy of accommodation of inefficiencies and monopoly power in non-tradable sectors.

In the last three years, inflation rates were lower in Czech Republic and Poland than in Hungary and Slovenia. As the Czech Republic and Poland maintains relatively less managed exchange rate regimes than Hungary and Slovenia, and additionally employ inflation targets, it is believed that the combination of relatively greater flexibility of exchange rate regime and inflation target 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 it 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 pre-adoption period may generate persistent inflationary pressure as candidate countries will probably try to maintain external competitiveness and use exchange rate as shock abosorber. For that reason one can 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 should 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, or eventually a currency board regime.

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

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(112,193) = 1.02			p-val.=0.45		
Normality	$\chi^2(8) = 26.20$			p-val.=0.00		
Heteroscedasticity	F(260,428) = 0.55			p-val.=1.00		
	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	2.28	7.07	55.71	229.10		
p-value	0.13	0.58	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 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 Czech Republic

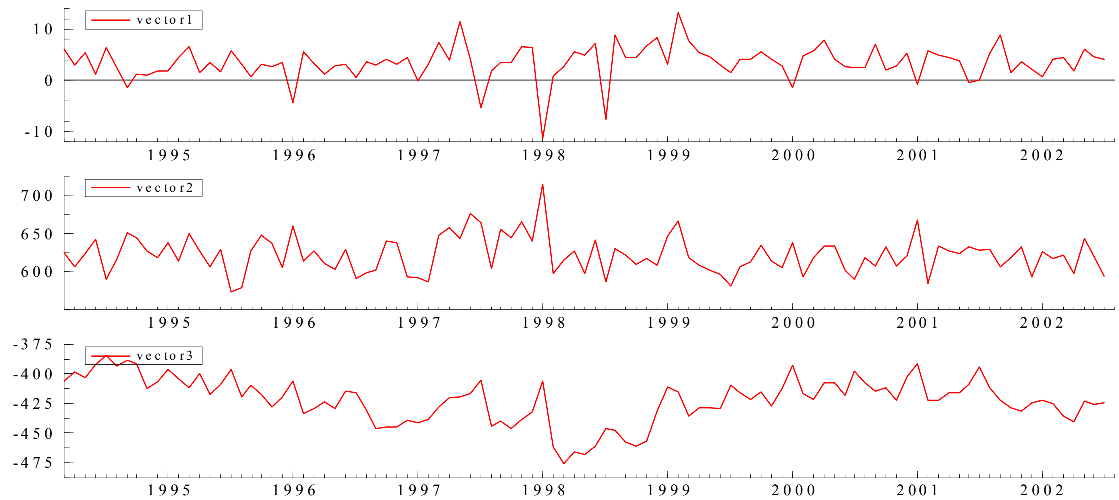

Figure A2: Cointegrating Vectors (unconcentrated) for Hungary

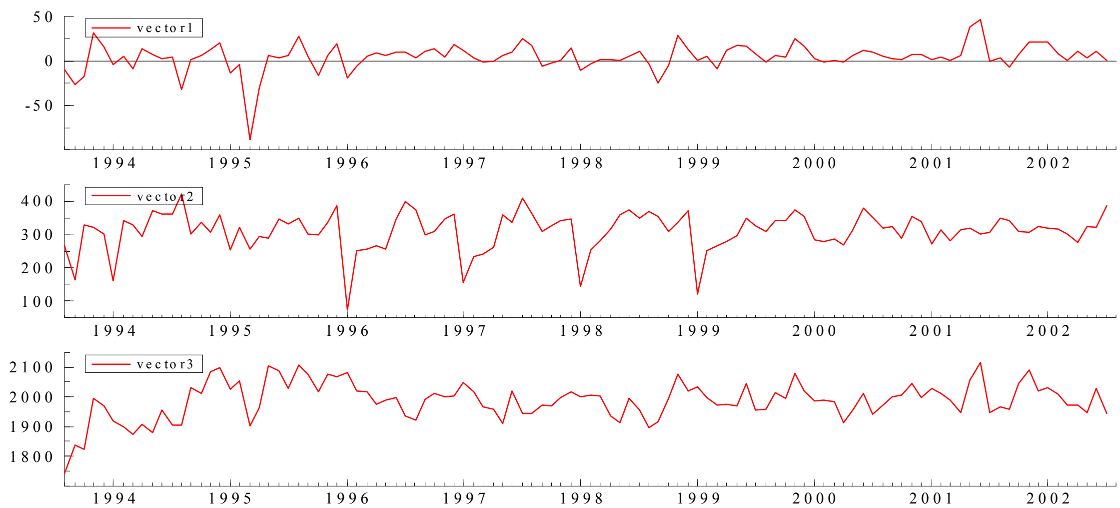

Figure A3: Cointegrating Vectors (unconcentrated) for Poland

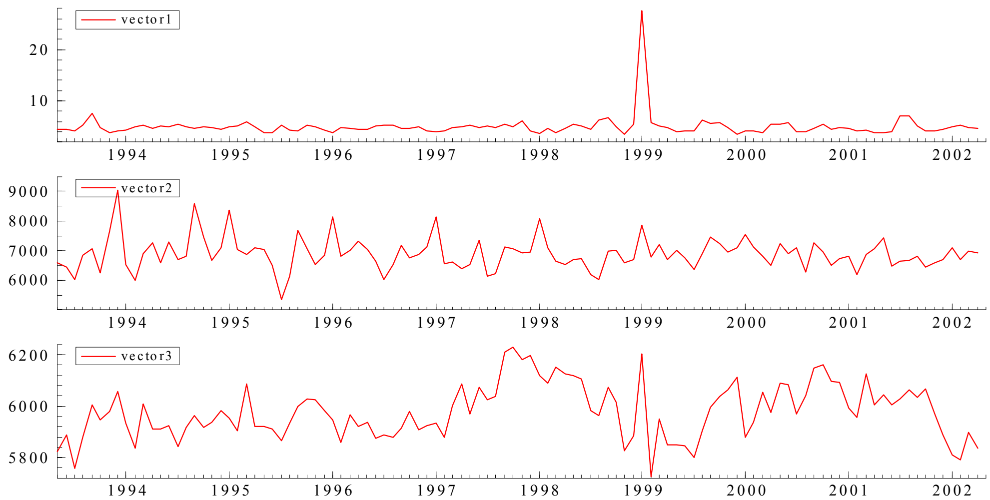

Figure A4: Cointegrating Vectors (unconcentrated) for Slovenia