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Exchange Rate Pass-Through in Acceding Countries: the Role of Exchange Rate Regimes[†]

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Abstract

This paper analyzes the link between the choice of exchange rate regime and inflationary performance in countries acceding to the EU. Estimation of pass-through effect of exchange rate changes to CPI inflation is complemented by I(2) cointegration 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. The size of the pass-through effect can be associated with the degree of accommodation in the exchange rate policy. The paper concludes that an accommodative exchange rate policy is one of the main sources of inflationary pressures in accession countries.

JEL codes: E42, E52, E58, C32

Keywords: EMU accession, pass-through effect, cointegration analysis, real exchange rate targeting, policy accommodation

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1. Introduction

Following accession to the European Union, candidate countries (CEECs) will eventually 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 Euro zone. The length of the transition period to the Euro is therefore an important policy choice for CEECs. The chosen strategy crucially depends on the role policy-makers attribute to the exchange rate for macroeconomic performance. In particular, a key question is the capability of affecting the real exchange rate through changes in the nominal exchange rate. This, in turn, depends on the impact that nominal exchange rate fluctuations have on domestic inflation, in other words, on the magnitude of the pass-through effect. In this respect, the aim of the paper is to empirically analyze the role of exchange rate regimes in overall inflationary performance of a subset of acceding countries: Hungary, the Czech Republic, Poland and Slovenia (CEEC-4 hereafter).

The interplay between the exchange rate regime and the speed of convergence of inflation rates between CEECs and the Euro zone is studied by estimating the 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 Slovakia). By contrast, Slovenia continued to maintain managed float, while others, the Baltic states 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 shocks. To answer this question it is important to distinguish between cases where higher exchange rate flexibility reflects an exchange rate policy geared at achieving a certain inflation target, and cases where exchange rate policy is accommodative, i.e. tries to neutralize the effects of adverse shocks on the real exchange rate. In the latter case, when taking the form of a reaction function to perceived disequilibria in the real exchange rate, a systematic component is induced into the dynamics of the nominal exchange rate. Such a policy is likely to be incorporated into the pricing decisions of economic 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 CEEC-4 and find that regimes with a more accommodative stance of exchange rate policy generate higher pass-through. 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.

With exchange rate pass-through we denote a change in the selected price index caused by the change in the nominal exchange rate. [1] Its empirical importance has been analyzed in a 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 on pass-through for CEEC-4 for the period 1993 to 2000.

Our study differs substantially from previous literature in the estimation methodology employed. Using a cointegrated vector autoregressive model we estimate the pass-through from exchange rates to prices and estimate the importance of shocks to the nominal exchange rate in the movements of domestic inflation for the CEEC-4. In addition, we invoke theoretical results from Johansen (2002) to address the issue of identification of pass-through effect, which has not been achieved in previous studies using cointegration analysis (e.g. Kim, 1998). Thus, we are not the first to use cointegration analysis to estimate exchange rate pass-through, but the first to solve the identification problem within this framework. This qualifies our estimates as the system generalization of single equation estimates of pass-through effect.

The empirical analysis indicates that pass-through is highly significant in the four candidate countries examined, although important differences emerge, which can be associated with differences in exchange rate regimes. 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. In this respect, it is not surprising that we found a perfect pass-through from exchange rate growth to domestic inflation for Slovenia and Hungary. A much 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. Even though it has never been officially declared, 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 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 is the worst inflation performer among the Acceding Countries in recent 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 the Czech Republic more to Poland. In general 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.

The analysis has a number of clear policy implications, all based on empirical fact that exchange rate changes importantly affect domestic inflation and hence in any disinflation experiment the central role should be given to the path of the nominal exchange rate. Even abstracting from the issue of propagation of exogenous shocks originating in international financial markets, flexible exchange rates are not an effective instrument for absorbing asymmetric real shocks (Masten, 2002). Large pass-through to import price is an incentive for policy-makers to attempt ex-post to drive the exchange rate in a way that improves external competitiveness. Large pass-through to CPI, however, reduces the expenditure switching effect and benefits of exchange rate flexibility. 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

[1] 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).

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.

The paper proceeds as follows. Section 2 presents stylized facts on inflation and exchange rate behavior in CEEC-4. After briefly discussing 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 offers a simple theoretical framework that links the size of pass-through exchange rate regimes and persistence of exchange rate shocks in particular. Section 4 contains the main empirical analysis of the paper, focusing on the pass-through. 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. Reduction to single-digit inflation was much faster in Slovenia and the Czech Republic, countries less affected by large stocks of debt and the attendant need to finance large debt service payments. However, inflation rates seem to be more stubborn in Slovenia and Hungary than in the Czech Republic and Poland. In the period 1998-2002, inflation hovered around 8-9 % in Slovenia and Hungary, with signs of decline after the second half of 2002, in a period of economic slowdown. The sharp decline in 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. 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 and relative price convergence (on the latter see Čihak and Holub, 2002). However, there is, in addition, a dynamic process connecting exchange rates and inflation.

Table 1 reports simple correlation coefficients between CPI-inflation rates, rates of change of nominal exchange rates and rates of change of real exchange rates. Data refer to 12-month changes to eliminate seasonal effects. In Hungary, Poland and Slovenia, the correlation between inflation rates and changes in nominal exchange rates is high. By contrast, the correlation between inflation and exchange rate changes is low in the Czech Republic. As a result, there was a high correlation between movements of nominal and real exchange rates. Interestingly, Hungary, Poland and Slovenia have kept positive depreciation rates on average. While Hungary broke up with this practice in second half of 2001 and Poland in 2000[2] Slovenia maintained a positive depreciation rate, which is entirely policy induced. It is especially notable for the Czech Republic that its nominal exchange rate growth never deviated permanently from zero, which considerably contributed to its most favorable inflationary performance in the CEEC-4 group. Furthermore, the two countries that used more actively the exchange rate as a policy tool, Hungary and Slovenia, displayed the smaller volatility in the real exchange rate, suggesting the presence of some form of real exchange rate targeting.

Table 1: Exchange rates and inflation, 1994:1-2002:5 (*)

	Czech Republic	**Hungary**	**Poland**	**Slovenia**
Correlation between changes in nominal exchange rate and inflation(**)	-0.2	0.9	0.8	0.5
	Mean			
-Nominal exchange rate	2.2	11.0	7.4	6.4
-Real exchange rate	-4.4	-4.5	-7.1	-3.6
-Consumer prices	7.1	16.0	15.5	10.4
	Standard Deviation			
-Nominal exchange rate	5.7	10.8	11.4	5.2
-Real exchange rate	6.6	4.9	6.1	4.5
-Consumer prices	3.2	7	9.5	4.7

(*) 12-month percentage changes
(**) Correlation coefficient between 3-month moving average of 12-month changes
Source: Authors' calculations on data from Datastream.

It is worth stressing that official changes in the exchange rate regimes need not be fully informative about the effect of exchange rate policy on inflation. As shown in the empirical part of the paper, models for these countries do not exhibit signs of structural parameter breaks even though three out of four countries introduced official changes in their exchange rate regimes. The degree of monetary policy accommodation can be revealed only by looking closely at actual path of the exchange rate and results of empirical analysis. For Slovenia it is clear that tight management was oriented towards sustaining a depreciation rate, which can be understood as a sign of accommodative monetary policy. In the Czech Republic this has not been the case even after the move to a more flexible arrangement. Thus even though two countries could be officially characterized as having a

[2] Poland adopted a flexible exchange rate system with inflation targeting in 2000, while Hungary switched to an inflation targeting regime in October 2001

similar exchange rate regime, it is important to take into account whether exchange rate management is used as a key tool for achieving a preset inflation target, or to accommodate any shock to the real exchange rate. While in the former exchange rate does not add to inflationary pressures, it does so in the latter.

3. Real Exchange Rate Targeting and the Size of Pass-Through Effect

Consider a small open economy with a continuum of monopolistically competitive firms that set their prices in a staggered manner as in Calvo (1983). Each firm resets its price in any period with fixed probability $1-\theta$, independently of other firms' price decisions and memoryless. By the law of large numbers, a fraction $1-\theta$ of firms resets their prices every period, while a fraction θ keeps their prices unchanged. Let pt denote the log of the aggregate price level, and p_t* the log of optimal price level of firms that receive a price-change-signal in period t. This implies that the current price level is described by the following law of motion (in log-linear form)

$$p_t = \theta p_{t-1} + (1-\theta) p_t^* \tag{3.1}$$

The firms maximize the expected discounted value of profits for given demand schedule, technology (common across firms), factor prices and the restriction on price adjustment. Log-linearization of the firm's first-order conditions around a zero-inflation steady state yields the following expression for optimal price level (see also Gali, 2001)

$$p_t^* = \mu + (1-\beta\theta)\sum_{k=0}^{\infty}(\beta\theta)^k E_t mc_{t+k} \tag{3.2}$$

where μ is the log of the optimal frictionless mark-up and mc_t the log of nominal marginal cost. $E_t(\cdot)$ is the rational expectations operator. Let us assume that nominal marginal costs are indexed to a linear combination of domestic prices and import prices

$$mc_t = (1-\gamma) p_t + \gamma(e_t + p_t^f) \tag{3.3}$$

where γ is the share of imported inputs in domestic production, e_t is the log of nominal exchange rate and p_t^f is the log of import prices denominated in foreign currency. We assume perfect exchange rate pass-through to import prices. For simplicity and without loss of generality assume that $p_t^f = 0$. Monetary policy follows a real exchange rate targeting rule, which is formulated as in Uribe (2003)

$$\Delta e_t = g(e_t + p_t^f - p_t), \qquad g'(\cdot) < 0 \tag{3.4}$$

This formualtion postulates that nominal depreciation increases in response to real apreciation. It also implies that transitory shocks to the real exchange rate combined with monetary policy rule induce a unit root in the level of (the log of) nominal exchange rate. Consider a shock to the real exchange rate at time t and no shocks afterwards. This induces a permanent shock to the nominal exchange rate: $e_{t+k} = e_t; \forall k$. It also implies that $p^*_{t+k} = p^*_t; \forall k$ along the perfect foresight path. By solving (3.1) forward and inserting into (3.3) we obtain the following expression for the nominal marginal costs

$$mc_t = (1-\gamma) p^*_t + \gamma e_t \tag{3.5}$$

Inserting (3.3) into (3.2) yields the expression for optimal price level in the case of real exchange rate targeting after a transitory shock to the real exchange rate

$$p^*_t = \mu + (1-\gamma) p^*_t + (1-\beta\theta) \sum_{k=0}^{\infty} (\beta\theta)^k E_t e_{t+k} = \frac{1}{\gamma}\mu + e_t \tag{3.6}$$

From (3.6) it follows that permanent shocks to the nominal exchange rate are associated with perfect pass-through to the optimal price level and by (3.1) also to the CPI in equilibrium. Albeit in a less general framework Taylor (2000) showed that persistence of shocks is a key determinant of the size of the pass-through. It has been shown above that shocks to the nominal exchange rate are perceived as permanent when agents incorporate a real exchange rate targeting policy into their pricing behavior. Again from (3.1) it follows that in such a case not only will equilibrium pass-through be perfect, but also that short-run pass-through will be higher. As discussed in the previous section, exchange rate policy in CEECs can to a great extent be characterized by elements of real exchange rate targeting (see also Laxton and Pesenti, 2003). Unlike in developed countries for which this practice cannot be confirmed, we would expect to find estimates of high pass-through. The extent to which CEECs have resorted to real exchange rate targeting has also differed substantially across countries. In those where it has been most systematic (Hungary and Slovenia) finding estimates of perfect equilibrium pass-through should not be surprising.

4. Estimation of Exchange Rate Pass-Through

This section presents the empirical analysis of pass-through effect of exchange rate growth to CPI inflation within a cointegration framework. Compared to existing studies of pass-through our empirical analysis introduces a number of novelties. First, within I(1) cointegration analysis the paper offers a formal discussion of identification of pass-through effect conditional on cointegration rank.

We show that reported estimates of equilibrium pass-though effect are actually identified and can thus be interpreted as elasticities in a structural sense. The issue of identification of pass-through effect has not yet been addressed in the existing literature and as is turns out this is of crucial importance when conclusions from empirical analysis are used for policy implications. Second, we discussed the necessary contemporaneous correlation of reduced form shocks that yields the empirically observed equilibrium pass-through effect. Finally, in order to provide better grounds for proposed policy implications, the analysis of common nominal I(2) trends is performed using I(2) cointegration analysis. The I(2) framework offers, first, useful tools for direct analysis of relative importance of shocks to different variables for overall inflationary performance in the economy, and, second, distinguish between nominal and real trends in the economy. All these features, combined with actual estimates, lead to important implications for the choice of exchange rate regime.

4.1. Methodological Issues and Limitations of Existing Empirical Studies

The measure of pass-through estimated in the paper is the effect of changes in nominal exchange rates on CPI inflation, a common final target variable of monetary authorities. In particular, the main focus of the paper is the equilibrium effects of exchange rate changes on inflation, real interest rates and output. In the literature there is no uniform approach to the analysis of pass-through. Some authors attempt to measure the pass-through directly; others use empirical results to investigate the underlying economic mechanisms (e.g. Choudhri, Faruqee and Hakura, 2002).

In terms of methodology, structural VAR is the most common in the literature, of which McCarthy (2000) is a very notable example. In that approach, pass-through is measured by means of impulse responses of different price series to an identified structural exchange rate shock. The problem with this approach is that it is not entirely consistent with the simplest notion of pass-through: any type of shock can cause the co-movement between the exchange rate and prices. In principle this would imply that we could observe as many measures of (short-run) pass-through as there are identified structural shocks. In addition, exchange rate changes need not occur only as consequences of stochastic shocks, but they can also reflect systematic changes in policy, that can be particularly important in countries adopting some form of real exchange rate targeting. It could occur also as the change in the inflation target of a central bank running disinflation policies. All such changes are not accounted for in a typical SVAR analysis. Proposed estimates of exchange rate pass-through could in such a case be severely biased and, not surprisingly, underestimated (see section 4.3 for a further discussion).

This problem does not arise in Campa and Goldberg (2002) who estimate a simple single-equation model for OECD countries and measure the pass-through effect (to import prices in their case) with the coefficient on the nominal exchange rate. Single equation approach is used also by Darvas (2001) for the group of Acceding Countries. A different use of SVAR analysis is found in Choudhri, Faruqee, and Hakura (2002). Their empirically observed impulse responses of various price indexes to an exchange rate shock are used not to measure pass-through effect directly but as a benchmark for simulated responses obtained from calibrated theoretical model under different assumption about nominal rigidities in the economy.

A common drawback of all SVAR-based studies is that they do not account explicitly for the possibility of cointegration. 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 long-run co-movement of prices and exchange rate seems very plausible theoretically. Neglecting cointegration when it is genuinely present means neglecting the intrinsic meaning of equilibrium long-run relationship between the nominal exchange rate and prices. In order to maintain comparability with typical SVAR-based studies this paper at the end of the analysis estimates also a structural vector error-correction (VEC) model (see Warne, 1993 for a detailed discussion).

Our analysis improves over existing studies of pass-through in three ways. First, the analysis is conducted within the framework of cointegrated vector autoregression model (CVAR). CVAR has been used in previous studies (Kim (1998); Billmeier and Bonato (2002)). As shown below, however, the estimates of the pass-through presented in these two studies are not identified.

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 justifications. We use instead the theoretical framework of Johansen (2002) to determine required contemporaneous changes in endogenous variables – contemporaneous pass-through effect – that yield equilibrium (long-run) responses. In other words, cointegration analysis effectively gives information on the contemporaneous correlation between variables that supports a permanent effect on variables of interest. This correlation is not estimated by imposing non-testable restrictions on reduced-form parameter space of a VAR, but is obtained directly from parameters of cointegrated VAR without imposing any additional restrictions apart from those required for identification of the cointegrating vectors. Moreover, the procedure used in the paper directly distinguishes between permanent and transitory shocks. For the analysis of pass-through this is a very important distinction, since, as shown in Section 3, only permanent exchange rate shocks can have a non-zero equilibrium pass-through effect and hence cause a different change in pricing behavior of economic agents. It is unlikely that transitory exchange rate shocks induce significant short-run changes in pricing behavior if firms face costs associated with frequent price changes. If the analysis is to be used for policy implications about disinflation policies and the choice of exchange rate regime, tracing the effects of permanent shocks separately becomes even more important.

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 (2002), 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. As a consequence, nominal shocks have a persistent effect on the level of inflation. Treating inflation as stationary results in invalid statistical inference.[3] Thus, all results obtained without testing for I(2)-ness in the price level should be interpreted with caution. We find that prices (as well as nominal wages and the nominal exchange rate) also in the present sample of countries – CEEC-4, can be better described as variables integrated of order two. The aim of the I(2) analysis is not to directly estimate the pass-through effect, which is done in I(1) framework, but to analyze the sources of nominal stochastic trends.

4.2 I(2) Analysis of Nominal Stochastic Trends

Detailed steps of testing and estimation are reported in Coricelli, Jazbec and Masten (2003). Here we present only the final results. 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 (all variables are in logs):

$$X_t = \left(y_t, e_t, cpi_t, ppi_t, w_t \right)' \quad t = 1993:1,...,2002:5 , \tag{4.1}$$

[3] 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 requires the use of monthly data.

where y_t denotes the index of total industrial production, e_t the nominal exchange rate (units of domestic currency per Euro), cpi_t the consumer price index, ppi_t the producer price index, and w_t average nominal wages.[4]

Table 2 contains estimates of matrices $\alpha'_{\perp 2}$ and $\beta_{\perp 2}$ of the I(2) cointegration model that describe the second order stochastic trends (see Johansen, 1995 for technical details of notation, here we concentrate on economic interpretation). $\alpha'_{\perp 2}$ matrix contains the coefficients (or weights) that shocks to each equation of system (4.1) have in the I(2) trend. We can write this as $\alpha'_{\perp 2}\sum_{s=1}^{t}\sum_{i=1}^{s}\varepsilon_i$, where ε_t is the vector of reduced-form residuals. $\beta_{\perp 2}$ measures the corresponding loading coefficients. The I(2) trend dominates the long-run stochastic behavior of the economy. If it turns out that this trend is nominal, we can conclude that it determines the predominant inflationary pressures in the economy.

For each of the four countries 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 equation most strongly contribute to the trend or whether this trend is equally or more strongly affected by shocks to the CPI, PPI or to nominal wage equation. In the first case this would imply that most of the inflationary pressures come via the exchange rate, due to the pass-through effect. In the second case most of the inflationary pressures would come 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. And finally, a high share of nominal wages would imply that important inflationary pressures come from aggressive trade unions.

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. With this we explore whether the index of industrial production is also an I(2) variable and whether the I(2) stochastic trend is predominately a nominal one.

The I(2) model is estimated with the 2-step procedure proposed by Johansen (1995b). When statistically supported (see test statistics below Table 2), α and β matrices obtained in the first step

[4] Wages in manufacturing sector only for the Czech Republic.

enter the second step restricted.[5] One common feature of the results is that shocks to output and nominal wages equations 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. This qualifies the I(2) trend as nominal for every country. The share of shocks to nominal exchange rate equation in the nominal stochastic trend is the highest in Slovenia, approximately twice as large as the corresponding shares of the CPI and PPI shocks, which are roughly equal. This implies that overall price inflation in Slovenia is most strongly affected by shocks to the nominal exchange rate, and much less from the autonomous pricing behavior of imperfectly competitive firms.

Table 2: 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.

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

[5] 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)).

shares of two price indexes. From the two, the share of PPI is higher. For Hungary the situation is different in the sense that shocks to the PPI contribute little to inflation, which is mostly affected by shocks to the exchange rate, and more importantly, by 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.

Examination of the $\beta_{\perp 2}$ vectors, measuring the loadings to second order stochastic trends, can identify candidate I(2) variables in our system. A common feature of Table 2 is that the loading coefficient of the I(2) trend into output is considerably smaller than other coefficients. This 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.

To summarize, in this section we have established that prices can be better treated 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 to be 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 firms to incorporate expected depreciations into their pricing behavior. Compared with a regime that on average maintains the level of the exchange rate stable the nominal exchange rate in such a regime could be less volatile (around a trend); however, a given change, more likely perceived to be permanent, would feed into prices more strongly and faster. 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 the 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.[6]

4.3 Estimating Pass-Through Effect in I(1) Framework

In this section we analyze the following I(1) system:[7]

$$X_t = \left(y_t, \Delta e_t, \pi_t - \pi_t^*, i_t - i_t^* \right)', \tag{4.2}$$

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

The inclusion of industrial production permits to control for the cyclical position of the economy. Without such controls the estimate of pass-through in an economy facing real appreciation will necessarily be larger than unity, that is also what Campa and Goldberg (2002) report for some countries. But such a coefficient clearly does not take into account only the effects of the nominal exchange rate on the price level and hence cannot be interpreted as the pass-through effect. For the same reason the system also includes foreign inflation and interest rate spread in order to account for interest rate parity effects.

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6 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.

7 We do not use a direct reduction from a I(2) to I(1) model, because formal testing revealed that it would not yield a nominal to real reduction that would enable identification of pass-through coefficient in line with Proposition 1 (see below). 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.

Exploring cointegration relations among the variables of the system enables us to estimate the equilibrium pass-through effect of (permanent) nominal exchange rate changes into the spread between domestic and foreign inflation. At first sight the measure of pass-through would seem to be the coefficient between inflation differential and exchange rate growth in a cointegration relation that contains both these variables and potentially other variables of the system. However, closer inspection of admissible long run equilibrium changes in variables combined with basic economic laws reveals that the measure of pass-through depends on cointegrating rank. Consider an I(1) system written as

$$\Delta X_t = \Pi X_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \Phi D_t + \varepsilon_t \qquad (4.3)$$

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

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

where $\alpha_\perp$ and $\beta_\perp$ are the orthogonal complements to α and β respectively. In the following we invoke results in Johansen (2002). From the solution of the error-correction model is follows that the long run value $X_{\infty/t}$ as a function of current values $(X_t, X_{t-1}, ..., X_{t-k+1})$ is given by

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

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

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

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

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

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

Economic meaning of Proposition 1 and its corollary can be seen from the following example. Let's look at the following example how Proposition 1 can be applied to this case.

Example: Let's write $X_t = \left(\pi_t - \pi_t^*, \Delta e_t, i_t - i_t^*, y_t\right)^{'}$. In our four-dimensional system cointegration rank is found to be three (*r=3*). In such a case identified cointegration relations can be written as

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

with corresponding cointegrating vectors: $\beta_1 = (1, -\lambda_1, 0, 0)^{'}$, $\beta_2 = (1, 0, -\lambda_2, 0)^{'}$ and $\beta_3 = (0, 0, 1, -\lambda_3)^{'}$. The orthogonal complement to these vectors (normalized on the second element) is

$\beta_{\perp} = (\lambda_1, 1, \lambda_2^{-1}, \lambda_2^{-1}\lambda_3^{-1})^{'}$, a one-dimensional space. In this case the pass-through effect is identified and directly measured by λ_1^{-1}. Note that rank three implies also that the long-run relation between exchange rate changes and inflation is supported by a non-zero effect on real output, whereas this is not necessarily the case for lower rank orders. We can combine the conclusions from these examples and the corollary to Proposition 1 to the following definition:

Definition 1: *Equilibrium pass-through effect* is measured by the coefficient of nominal exchange rate growth on the difference between domestic and foreign CPI inflation in a cointegrating relationship that contains no other variables.

In line with Proposition 1 identification of long-run or equilibrium pass-through effect depends on cointegrating rank. When identification can be achieved (cases with *r=2* or *3*), we can also determine what is the corresponding contemporaneous pass-through effect – a short-run change – that supports a given long-run change. As seen above, a given long-run change $k \in sp(\beta_\perp)$ can be achieved by adding Γk to X_t. We can interpret a change in X_t by Γk also as the effects of shocks that clearly have permanent effects on (some) variables in X_t. In fact, this is a restriction any type of shock in structural sense (real and nominal) that economic theory can justify to have a permanent change given by $k \in sp(\beta_\perp)$. In addition, short-run changes Γk need not occur only due to exogenous shocks (ε_t in expression 4.3), but also as a consequence of systematic policy reactions and changes in policy targets.

Our focus is on permanent effects of changes in exchange rate growth on inflation, because exchange rate shocks that are only transitory do not necessarily induce significant changes in pricing behavior of firms Γk contains the contemporaneous effect of a permanent exchange rate change on inflation and other variables. This characterization of "shocks" is not based on any non-testable restriction imposed on reduced form parameters of the (cointegrated) VAR. Indeed, any alternative identification scheme that can be found in structural VAR literature should be able to replicate this contemporaneous effects otherwise it would fail to identify the true permanent exchange rate shock in structural sense. All other shocks (three in our case) cause only transitory fluctuations around long-run equilibrium values. This example is also instructive for how misleading would be the structural VAR analysis without initially performing the permanent-transitory decomposition, which effectively boils down to testing for cointegration.

4.4 Results of I(1) Analysis

Lag length of the systems has been chosen by complementary use of standard information criteria and the usual Wald-type tests for a valid lag reduction. It proved sufficient to include two endogenous lags for Slovenia, three for Hungary and the Czech Republic and four for Poland.[8] Tests 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 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.[9]

The choice of cointegration rank is 3, uniformly across all four countries. Both the asymptotic and bootstrap versions of the trace test indicate this very clearly for Slovenia, Hungary and Poland, while rank 2 is also possible for the Czech Republic. We have nevertheless chosen rank 3 also for the latter country as the systems show significant and strong equilibrium correction to the third cointegrating relation, and because we wanted to maintain direct comparability of results between all four countries. Visual inspection of the estimated third cointegrating vector presented in Figures A1 – A4 also show no obvious signs of non-stationarity for all countries and the Czech Republic in particular.

The left panel of Table 3 presents the estimates of just-identified cointegrating vectors. The right panel reports the corresponding adjustment coefficients. Signs of all coefficients are consistent with economic theory. The most informative for the analysis of pass-through effects is the second cointegrating vector β_2 and, in particular, its first coefficient. As explained in previous section, it can be interpreted as long run or equilibrium pass-through effect. We can observe that it is the largest in Slovenia and practically identical to 1. For Hungary it is only marginally different, and for both countries the restriction that it is actually equal to 1 cannot be rejected, (see the corresponding likelihood ratio tests reported under Table 2) with corresponding p-values above 0.90. For Poland the point estimate of this coefficient is 0.86; 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

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

[9] The parameter stability of the VARs has been tested with recursive 1-Step and Break-Point Chow tests. The tests reveal no signs of parameter instability for all countries. The same turns out to be the case for the Hansen-Johansen test of constancy of cointegration coefficients. These test are available upon request.

hypothesis that the pass-through coefficient is equal to 0.5, we cannot reject the restrictions (corresponding p-value is 0.09).

Table 3: 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	-	**-0.64** (0.06)	-	**-62.06**	**-1.03**	**-1.28**
$\pi - \pi^*$	-	**1.00**	**-1.28** (0.19)	**-20.83**	0.02	**-0.79**
i-i^*	**0.03** (0.001)	-	**1.00**	-0.68	-0.00	0.02
			Hungary[b] (1993:2 – 2002:7)			
y	**1.00**	-	-	**-0.09**	0.00	0.00
Δe	-	**-0.97** (0.10)	-	-4.87	**-0.88**	**-0.98**
$\pi - \pi^*$	-	**1.00**	**-1.49** (0.11)	2.62	0.01	**-0.73**
i-i^*	**0.03** (0.004)	-	**1.00**	**-1.10**	0.00	0.01
			Poland[c] (1993:1 – 2002:4)			
y	**1.00**	-	0.00	-0.022	0.00	**0.001**
Δe	-	**-0.86** (0.10)	0.00	**-59.79**	**-1.23**	**-1.02**
$\pi - \pi^*$	-	**1.00**	**-0.84** (0.08)	**-25.37**	-0.01	**-0.78**
i-i^*	**0.03** (0.006)	-	**1.00**	-0.64	-0.00	0.01
			Slovenia[d] (1993:3 – 2002:3)			
y	**1.00**	-	-	**0.03**	-0.00	-0.00
Δe	-	**-1.01** (0.10)	-	**-26.49**	**-0.62**	**-0.78**
$\pi - \pi^*$	-	**1.00**	**-2.32** (0.20)	-2.66	0.02	**-0.75**
i-i^*	**0.01** (0.001)	-	**1.00**	**-18.51**	0.00	**0.38**

Bold indicates significance. Standard errors in parentheses.

[a] *H₀*: $\beta_{21} = -1$; $\chi^2(1) = 2.52$, p-val.=0.11. Weak exogeneity of y: $\chi^2(3) = 3.03$, p-val.=0.39. Weak exogeneity of i-i^*: $\chi^2(3) = 9.42$, p-val.=0.02. Weak exogeneity of y + H_0: $\beta_{21} = -0.5$; $\chi^2(4) = 8.07$, p-val.=0.09.

[b] H_0: $\beta_{21} = -1$; $\chi^2(1) = 0.01$, p-val.=0.92

[c] H_0: $\beta_{21} = -1$; $\chi^2(1) = 0.24$, p-val.=0.61, H_0: $\beta_{32} = -1$; $\chi^2(1) = 0.76$, p-val.=0.38

[d] H_0: $\beta_{21} = -1$; $\chi^2(1) = 0.00$, p-val.=0.99

To complete the exposition of I(1) analysis it is also important to look at the corresponding α coefficients in Table 3, 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 β_{21} 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. In the interpretation of β_{21} coefficient this allows us to abstract from output movements that cause trend movements in the real exchange rate, namely real appreciation.

It is surprising at first sight that only the exchange rate adjusts strongly and significantly to the second cointegrating relation. One would expect this also for the inflation rate differential. However, corollary to Proposition 1 (see also the left panel of Table 4 below) implies that any permanent increase in exchange rate depreciation that also increases inflation must be matched with a corresponding increase in the interest rate spread. We can observe from the vector of adjustment coefficients to the third cointegrating relation α_3 a very strong equilibrium adjustment of inflation (positively in response to a positive exchange rate shock that increases the interest rate spread). Therefore, shocks to the second cointegrating relation alone cannot be interpreted as exchange rate changes 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.

4.5 Short-run and long-run changes

The orthogonal complement to β, denoted by $\beta_\perp$, is a 4x1 vector for each country. Estimates are reported in the left panel of Table 4. Admissible long-run co-movements of the variables analyzed are thus summarized by a one-dimensional space, which has qualitatively the same structure across countries. In the explanation lets consider the case of Slovenia: $\beta_\perp$ is (-0.025, 1,1,2.5)'. This means that an equilibrium or permanent change in exchange rate growth accompanied by an equivalent increase in inflation differential, a disproportionate increase in interest rate spread, and consequently a lower level of output. The effect on the level of output is significant for all countries. Because the change in interest rate spread is more than proportional, ex-post real interest rate spread also increases, which is most likely the cause of lower output.[10] A disproportionate effect on the interest rate differential is a consequence of non-stationarity of the risk premium and the real interest rate spread. This non-stationarity is observed for all countries and is not surprising as these countries had gone through the process of transition to market economies. It is important to note that any other vector linearly independent of (-0.025, 1,1,2.5)' will violate the orthogonality with respect to β and hence cannot span the equilibrium long-run changes in variables of interest. In particular, monetary authorities cannot permanently depreciate the currency without causing an equal change in the

[10] To see this, note that we can rewrite the third cointegrating relation generically denoted as $\pi - \pi^* = \beta_{32}^{-1}(i - i^*)$ as $r - r^* = (1 - \beta_{32}^{-1})(i - i^*)$, where r denotes the ex-post real interest rate.

difference in CPI inflation rates. In other words, we must note that $(-0.025, 1, \lambda, 2.5) \notin sp(\beta_{\perp})$ for any $\lambda \neq 1$. This means that we can in line with Proposition 1 indeed interpret the coefficient as a measure of equilibrium pass-through into CPI inflation. Moreover, as noted above, any permanent rise in the rate of depreciation results in a rise in real interest rate differential and a negative effect on output.

Table 4: Orthogonal complements to cointegrating space - $\beta_{\perp}^{'}$ and Γk [11]

	$\beta_{\perp}^{'}$				Γk			
	y	Δe	$\pi - \pi *$	i-$i*$	y	Δe	$\pi - \pi *$	i-$i*$
Czech Republic	-0.018	1	0.5	0.625	-0.038	7.87	-0.59	0.39
Hungary	-0.045	1	1	1.5	-0.092	1.00	1.05	1.01
Poland	-0.020	1	0.8	0.67	-0.047	0.33	0.63	0.49
Slovenia	-0.025	1	1	2.5	-0.019	0.84	0.84	1. 58

Hungary again shows very similar properties to Slovenia, the only difference being that a policy of further exchange rate stabilization while yielding the same gain in inflation reduction yields a somewhat smaller, but still more than proportional, reduction in interest rate spread, and a larger positive effect on output. Poland and the Czech Republic share many similarities and appear distinct from Hungary and Slovenia. Their point estimate of pass-through is smaller than one, and consequently, there is also a smaller negative effect on output and a smaller required increase in interest rate spread to support a potential policy of accelerated exchange rate depreciation.[12] Again this fits with our priors about the nature of exchange rate policy in these two countries.

Right panel of Table 4 reports required initial (short-run) changes (see Johansen, 2002) in current values of endogenous variables Γk that yield a long-run response of the system spanned by $\beta_{\perp}$. These results, combined with the conclusions of I(2) analysis about the most important sources of nominal pressures in the economies, show that results on pass-through effect cannot be discussed independently of the specific type of the exchange rate regime in a certain country. It can be observed

[11] The asymptotic distribution of parameters in Table 4 is mixed Gaussian. For this reason, the corresponding standard errors are not reported. The figures in right panel of Table 4 were calculated using significant parameters in matrix Γ only.

[12] Note that for Poland even though the point estimate of β_{32} in Table 3 is larger than 1, it is not significantly different from 1. This means that Polish risk premium and real interest rate differential have been stationary in the period under analysis. A change in the rate of depreciation of zloty thus does not yield lower equilibrium real interest rates.

that in Hungary and Slovenia, two countries with most accommodative exchange rate regimes in the past, a "shock" that permanently increases the growth of the exchange rate is accompanied with practically equal change in inflation differential already on impact. This is a sign of very high short-run pass through. The hypothesis that these contemporaneous changes can be induced by economic policy cannot be accepted based only on these results. However, it is also true that the hypothesis that this reflects a high degree of policy accommodation cannot be entirely rejected. It can be at least considered as an indicator of accommodative exchange rate policy that aims to stabilize short-run fluctuations in the real exchange rate. Very high contemporaneous correlation between changes in exchange rate growth and inflation can thus be seen as a sign of the interplay between policy accommodation and forward-looking pricing behavior that induces an important systematic component into domestic inflation. In such circumstances, if the central bank operating a managed float regime does not internalize the inflationary consequences of its past actions, might find itself in a vicious circle where trying to respond to inflationary pressures leads instead to a full accommodation of these. Inflation thus becomes more persistent not because of various forms of nominal rigidities but because of accommodative monetary policy. A fully forward-looking and accommodative monetary policy, even though re-optimizing in every decision cycle and committing to minimization of the social loss function, lacks necessary history dependence for optimality and determinacy as discussed in Svensson and Woodford (2003). Moreover, such a policy can easily violate the Taylor principle, which implies that disinflation period without a regime change can be quite lengthy, if possible at all.[13]

Exactly the opposite is the case of the Czech Republic, where instantaneous pass-through effect is practically zero (even with a negative point estimate). This is consistent with observed path of Czech exchange rate growth, which, even though non-stationary, does not show any tendency to drift away from zero even though it exhibits large and persistent swings, to a large extent caused by speculative capital flows. For Poland the results are in line with the findings of I(2) analysis, where we observe that shocks to exchange rate do not contribute to the nominal stochastic trend. The right panel of Table 4 shows a very similar picture. Consistently with imperfect, but still large, pass-through effect, exchange rate policy seems to act as to gradually and imperfectly accommodate inflationary pressures, hence not acting as a string source of inflationary pressures itself.

The reasoning based on the results in right panel of Table 4 is confirmed also with a structural VEC analysis (see Warne, 1993). For all four countries we have identified the structural parameters by

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

assuming a recursive ordering of three transitory shocks. With this we obtain also the estimates of structural contemporaneous effects of the permanent shock (only one in our systems), which is of our primary concern. The impulse responses of nominal exchange rate growth and inflation differential to the permanent shock – single stochastic driving trend - are plotted in Figure 5. As could be anticipated from previous discussion for Hungary and Slovenia we can clearly observe considerably smaller short-run fluctuation and much smoother transitional dynamics to the new steady state. Moreover, Hungary and Slovenia have a much faster pass-through effect. Most of the permanent (!) change in exchange rate growth is transferred into inflation differential within one year, which is in line with the theoretical analysis in section 3. From the theoretical model it also follows that if we would restrict ourselves to analyzing only the responses to transitory shocks, as is most commonly done in SVAR-based studies of pass-through, the observed speed and magnitude of pass-through would be smaller. However, effects of transitory shock are of relatively small importance when we want to address the necessary disinflation policies necessary to obtain a given policy objective i.e. entry to the Euro area. For this, the effects of permanent changes are of central importance.

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, two countries with most accommodative exchange rate policy. Moreover, from I(2) analysis for Slovenia it follows that innovations to the exchange rate are transferred most strongly to domestic inflation. 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 in both price indexes are more important components of the I(2) nominal stochastic trend.

5. Concluding Discussion

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

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seriously questioned, however.

working of the Balassa-Samuelson effect and the process of overall relative price convergence in CEEC-4 on average. However, it is argued in this paper 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.

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. Empirical analysis of this paper shows that inflation rates in CEEC-4 are non-stationary, which implies that these economies exhibit full inflation inertia and not only price inertia. Calvo, Celasun and Kumhof (2002) have recently proposed a theoretical framework that incorporates inflation inertia into the framework of staggered price setting of firms in a monopolistically competitive market. The intuition of this 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 internalize 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. As a result, the model contains inflation inertia in addition to price-level inertia. A simulation of a disinflation policy implemented through a reduction of the rate of depreciation of the exchange rate shows that output in the non-tradable sector temporarily declines. Under the assumption of monopolistic competition also in the tradable sector the output decline, even though smaller, would have occurred also in the tradable sector. Nevertheless, 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 exchange rate policy in candidate countries, because it shows that in presence of high and fast pass-through effect, which has been found for two countries with most accommodative exchange rate policies: Hungary and Slovenia, disinflation can be achieved by a different choice of exchange rate policy at low costs in terms of output decline and with potential gains in welfare.

What are the policy implications that can be drawn from presented empirical results? The most important conclusion is that in any policy design the important effect of the nominal exchange rate on prices should not be underestimated. It has been show theoretically that accommodative exchange rate policy is associated with high pass-though effect. This has been confirmed also in the empirical analysis. The path of nominal exchange rate within a more general exchange rate regime arrangement during disinflation should thus be given an important role.

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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 (asymptotic)	1.29	12.78	44.90	101.32		
p-value (asymptotic)	0.26	0.14	0.00	0.00		
p-value (bootstrap)	0.45	0.18	0.00	0.00		
r	3	2	1	0		
Hungary (1993:8 – 2002:7)						
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		
p-value (bootstrap)	0.36	0.04	0.00	0.00		
r	3	2	1	0		
Poland (1994:1 – 2002:4)						
Res. autocorr. 1-7	F(112,132) = 1.28			p-val.=0.09		
Normality	$\chi^2(8)$ = 19.51			p-val.=0.01		
Heteroscedasticity	F(320,247) = 0.37			p-val.=1.00		
	Modulus of 6 largest characteristic roots					
Unrestricted VAR	0.98	0.91	0.91	0.76	0.76	0.62
r =3	1.00	0.89	0.67	0.67	0.59	0.59
Trace test	5.42	20.91	57.02	221.56		
p-value	0.02	0.01	0.00	0.00		
p-value (bootstrap)	0.10	0.03	0.00	0.00		
r	3	2	1	0		
Slovenia (1993:3 – 2002:3)						
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		
p-value (bootstrap)	0.41	0.06	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 1: Cointegrating Vectors (unconcentrated) for the Czech Republic

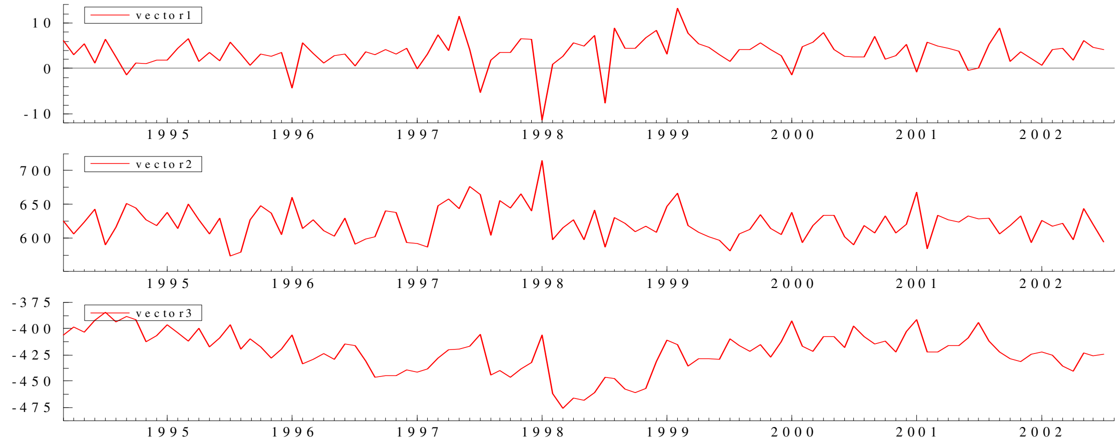

Figure 2: Cointegrating Vectors (unconcentrated) for Hungary

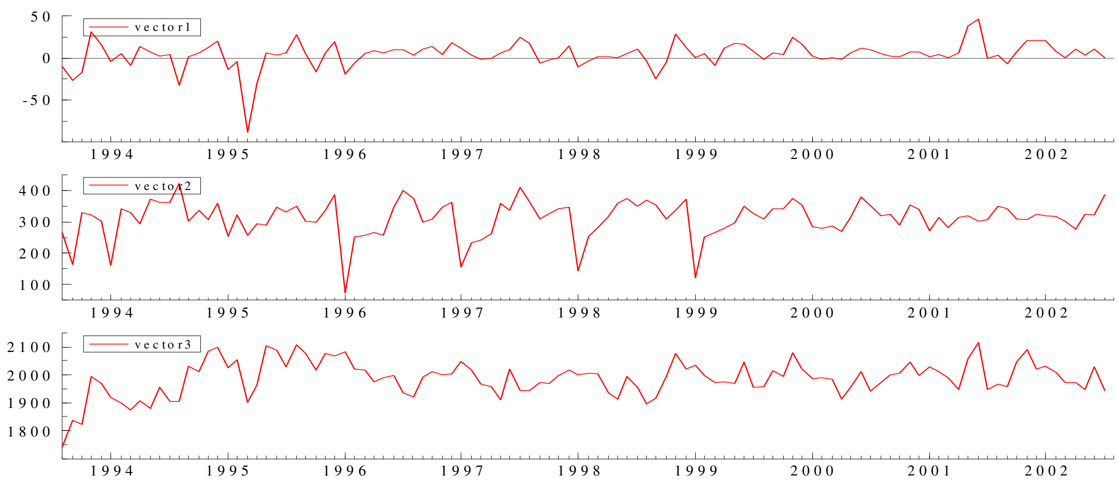

Figure 3: Cointegrating Vectors (unconcentrated) for Poland

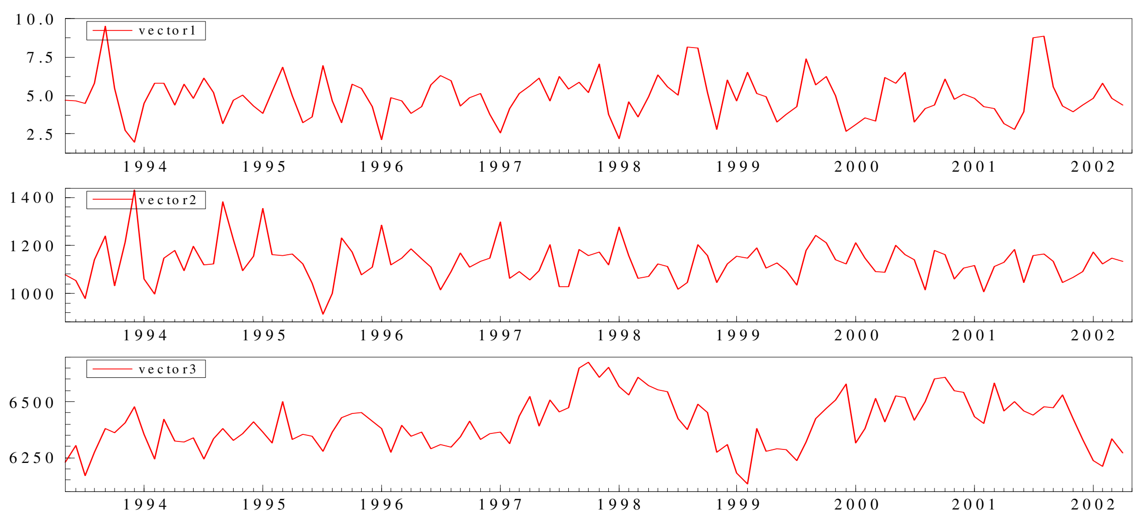

Figure 4: Cointegrating Vectors (unconcentrated) for Slovenia

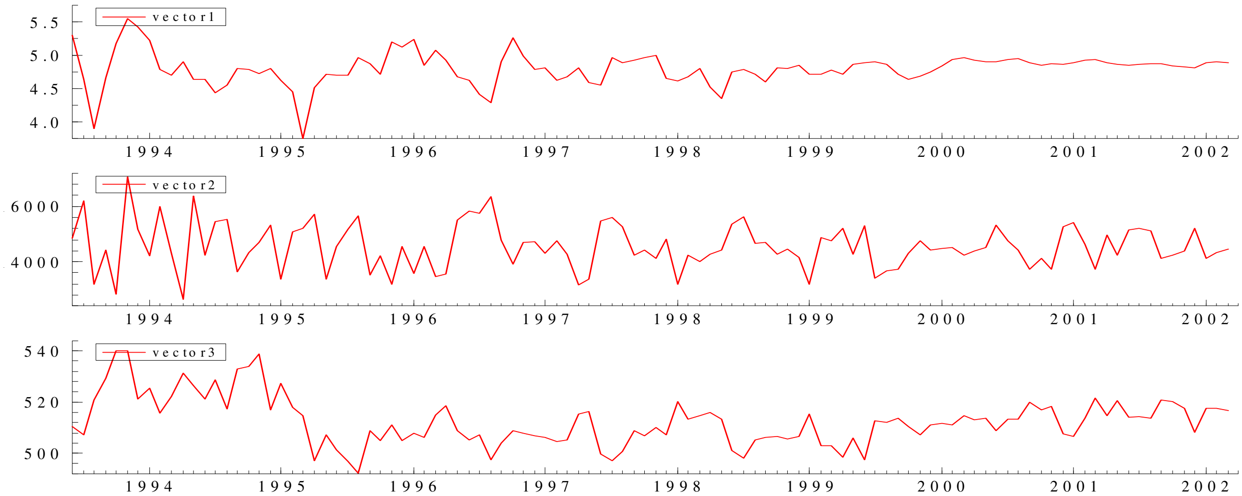

Figure 5: Impulse responses of nominal exhcange rate growth and inflation differential to the permanent shock (with 95 % asymptotic confidence inntervals)