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Abstract

A panel data set for six Central and Eastern European countries (the Czech Republic, Hungary, Poland, Romania, Slovakia and Slovenia) is used to estimate the monetary exchange rate model with panel cointegration methods, including the Pooled Mean Group estimator, the Fully Modified Least Square estimator and the Dynamic Least Square estimator. The monetary model is able to convincingly explain the long-run dynamics of exchange rates in CEECs, particularly when this is supplemented by a Balassa-Samuelson effect. We then use our long-run monetary estimates to compute equilibrium exchange rates. Finally, we discuss the implications for the accession of selected countries to the European Economic and Monetary Union.

Keywords: Exchange rates, monetary model, panel unit root tests, panel cointegration, EMU.

JEL Classification: C33, F31, F36.

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1 Introduction

Applied research on the economics of exchange rates experienced a revival during the 1990s due in part to the application of nonstationary time series methods. One key area of application involved testing the purchasing power parity hypothesis using nonstationary panel methods (see for example Frankel and Rose, 1995, and MacDonald, 1996). In this paper, we use various panel cointegration estimators to estimate a variant of the monetary model of the exchange rate using data from six transition countries (the Czech Republic, Hungary, Poland, Romania, Slovakia and Slovenia). We extend the basic monetary model to capture the Balassa-Samuelson effect, which is generally found to play an important role in transition countries (see for example MacDonald and Wójcik, 2003). Furthermore, we take into account the fulfillment of the uncovered interest parity condition in transition economies, since these countries were characterized by important capital market imperfections during our sample period.

Among our conclusions are the following: we show that the augmented monetary model provides a good description of nominal exchange rates trends and find a significant Balassa-Samuelson effect; although deviations from the uncovered interest parity are also significant, we document that the size of this effect is rather small.

Finally, we consider the issue of the integration of selected transition countries into Economic and Monetary Union (EMU). Fidrmuc and Korhonen (2003) and Fidrmuc (2003) show that the euro area and the CEECs can be increasingly considered an optimum currency area. Furthermore, Kočenda (2001) and Kutan and Yigit (2003) demonstrate increasing similarities in the real and monetary developments between the euro area and the CEECs.

The paper is structured as follows. The next section introduces the monetary model of exchange rate, augmented with a Balassa-Samuelson effect. Section 3 describes our panel data set, while Section 4 contains a set of unit root tests. Section 5 presents several estimates of the monetary model, which are used for simulations of the equilibrium exchange rates in Section 6. Finally, Section 7 concludes.

2 The Monetary Model of the Exchange Rate

The monetary model of the exchange rate has become something of a workhorse in the exchange rate literature. An estimable reduced form is usually generated from an ad hoc framework comprising money demand functions in the home and foreign country. Although this approach has been criticized, we nonetheless follow it here, since it produces a reduced form which is very similar to that derived in an optimizing framework (such as that of Lucas, 1982).

The monetary model is usually presented as a two-country, two-money, two-bonds (where the bonds are assumed to be perfect substitutes) model in which all goods are tradable and the law of one price holds. Money demand relationships are given by standard Cagan-style log-linear relationships:

$$m_t^D - p_t = \beta_0 y_t - \beta_1 i_t, \tag{1}$$

$$m_t^{D*} - p_t^* = \beta_0 y_t^* - \beta_1 i_t^*, \tag{1'}$$

where $\beta_0, \beta_1 > 0$, m^D denotes money demand, p denotes the price level, y is output, i the interest rate, lowercase letters indicate that a variable has been transformed into natural logarithms (apart from the interest rate), and an asterisk denotes a foreign magnitude. For simplicity, we assume that the income elasticity, β_0, and the interest semielasticity, β_1, are equal across countries. If it is additionally assumed that money market equilibrium holds continuously in each country:

$$m_t^D = m_t^S = m_t,$$

$$m_t^{D*} = m_t^{S*} = m_t^*,$$

then using these conditions in (1), and rearranging for relative prices, we obtain

$$p_t - p_t^* = m_t - m_t^* - \beta_0 (y_t - y_t^*) + \beta_1 (i_t - i_t^*). \tag{2}$$

On further assuming that the purchasing power parity (PPP) theory or the law of one price (LOOP), holds for relative prices, we obtain a base-line monetary equation as

$$s_t = m_t - m_t^* - \beta_0 (y_t - y_t^*) + \beta_1 (i_t - i_t^*). \tag{3}$$

In words, the nominal exchange rate, s, is driven by the relative excess supply of money. Holding money demand variables constant, an increase in the domestic money supply relative to its foreign counterpart produces an equiproportionate depreciation of the currency. Changes in output levels or interest rates have an effect on the exchange rate indirectly through their effect on the demand for money. Thus, for example, an increase in domestic income relative to foreign income, ceteris paribus, produces a currency appreciation, while an increase in the domestic interest rate relative to the foreign rate generates a depreciation.

However, the PPP assumption necessary to derive (3) is clearly not tenable given the extant empirical evidence, which suggests the mean reversion of real exchange rates is too slow to be consistent with PPP (see, for example, Froot and Rogoff, 1995, and MacDonald, 1995). One important explanation for the persistence in real exchange rates is the existence of real factors, such as the Balassa-Samuelson effect, which drive the nominal exchange rate away from its PPP-defined level. Indeed, MacDonald and Ricci (2001) have demonstrated the importance of this effect in explaining the persistence of the real exchange rates of a group of industrialized countries. Since such real effects are likely to be at least as important for the current group of accession countries, we incorporate a Balassa-Samuelson effect into the monetary equation.

Following Clements and Frenkel (1980), a Balassa-Samuelson effect may be incorporated into the monetary equation in the following way. Assume that overall prices in the home and foreign country are a weighted average of the price of traded and nontraded prices:

$$p_t = \alpha p_t^T + (1-\alpha) p_t^{NT} \tag{4}$$

$$p_t^* = \alpha p_t^{T*} + (1-\alpha) p_t^{NT*} \tag{4'}$$

where p now represents overall prices, incorporating both traded and nontraded components, p^T represents the price of traded goods, p^{NT} is the price of nontraded goods and α denotes the weight (for simplicity we assume the same weights in both countries). Consider the definition of the real exchange rate (LOOP holds in the tradable sector), defined with respect to overall prices (i.e. the CPI):

$$q_t \equiv s_t - p_t + p_t^*, \tag{5}$$

where q is the real exchange rate. We define a similar relationship for the price of traded goods as:

$$q_t^T \equiv s_t - p_t^T + p_t^{T*}. \tag{6}$$

Using (4), (5) and (6), the following expression may be obtained for the real exchange rate

$$q_t = q_t^T - (1-\alpha)\left[\left(p_t^{NT} - p_t^T\right) - \left(p_t^{NT*} - p_t^{T*}\right)\right]. \tag{7}$$

Using expression (7) in (2), we may obtain the following equation,

$$s_t = m_t - m_t^* - \beta_0 (y_t - y_t^*) + \beta_1 (i_t - i_t^*) - (1-\alpha)\left[\left(p_t^{NT} - p_t^{T}\right) - \left(p_t^{NT*} - p_t^{T*}\right)\right], \quad (8)$$

where the nominal exchange rate is predicted to appreciate as the relative price of nontraded to traded goods rises.

3 Data Description

Although we have access to monthly data for the period January 1993 to December 2002, our analyses will concentrate on the subperiod September 1994 to March 2002. This allows us to estimate the monetary model with panel cointegration methods and a balanced sample.[1]

We have included six Central and Eastern European countries in our data sample: the Czech Republic, Hungary, Poland, Romania, Slovakia and Slovenia.[2] Rahn (2003) and Šmídková et al. (2002) use similar panels to estimate BEER (behavioral equilibrium exchange rates) and FEER (fundamental equilibrium exchange rates) models of real exchange rates. It is important to bear in mind that several of the countries in our panel moved from adjustable pegged exchange rates to a managed or free-floating regime during the sample period, so that our sample period does not represent a homogeneous exchange rate regime. The official changes took place in 1997 in the Czech Republic, in 1998 in Slovakia and in 2000 in Poland. In all these cases, however, the official change followed after previously widening the fluctuation bands to up to ±15%. The introduction of floating exchange rates was necessitated by currency crises in the Czech Republic (see Horvath and Jonas, 1998) and Slovakia. However, the time series on nominal exchange rates do not seem to display a structural break related to the exchange rate regime change, although the variance of several variables was higher around periods of currency crises in the case of the Czech Republic and Slovakia.

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[1] Estimations with the longer, unbalanced sample were used to check the robustness of the parameter estimates to the inclusion of earlier transition periods. Although the parameters remain in the range of those presented for the balanced sample, for some countries the use of the sample back to 1993 affects the conclusions on the current position of the nominal exchange rate with respect to the equilibrium rate.

[2] Although we have data on all ten candidate countries, in this paper we focus on countries with relatively flexible exchange rate regimes.

While the exchange rate regimes of our group of CEECs were relatively flexible during the whole period, Hungary followed a narrow-band crawling peg system up to May 2001 (that is, during the whole analyzed period). Therefore, it could be argued that Hungary should be excluded from our data sample. However, our robustness analyses do not indicate that this is necessary.

The variables in our data set comprise the nominal exchange rate vis-à-vis the euro (expressed as local currency units per euro), the money stock (M2) and industrial production. Furthermore, we include deposit interest rates and the ratio of consumer prices to producer prices to capture the deviations from the uncovered interest parity and the Balassa-Samuelson effect, respectively. All conditioning variables are defined as deviations from the corresponding variables for the euro area.[3] In instances where we introduce time dummies into our models, the euro numeraire is of course removed. All variables except interest rates (see the definition of interest rates below) were indexed as 100 to the base year 1995 and are converted into logs. As far as possible, data on the CEECs are taken from the IMF's International Financial Statistics. This database is complemented by national sources and publications of The Vienna Institute for International Economic Studies (WIIW).

An extended time series for the euro was obtained by using the so-called synthetic euro, that is, the ECU excluding the currencies of those countries which did not introduce the euro in 1999 (or 2000 in the case of Greece): Denmark, Sweden and the UK. Given this definition, there should be no structural break in 1999 for any of the countries.

Nominal exchange rates in our group of CEECs fluctuated significantly during the sample period. In general, the currencies of CEECs depreciated during the first part of the sample, and we can see a stabilization of nominal exchange rates (with the exception of Romania and Slovenia) in some countries around 1998. Thereafter, nominal exchange rates started to appreciate in the Czech Republic (in 2000), Hungary (2001), Poland (2001) and Slovakia (2002).

[3] We used data for Germany as a proxy for the euro area as well. The results, which are available from the authors on request, do not substantially differ from presented results.

4 Panel Unit Root Tests

Given the long-run positive inflation differential between the euro area and the CEECs, we would expect all nominal variables to display a clear trend pattern. A similar feature is expected for industrial production, given the real convergence of CEECs to the EU's income level. Standard unit root tests for single time series confirm that the majority of the individual time series are I(1) processes.[4] As is now well known, adding a cross-sectional dimension to unit root tests can potentially improve the quality of these tests significantly by increasing their power.[5] Furthermore, an important contribution of panel unit root tests is that the resulting test output can be normalized to statistics that have limiting standard normal distributions. According to Baltagi and Kao (2000), this phenomenon is due to the fact that individual data units along the cross-sectional dimension can act as repeated draws from the same distribution.

Quah (1992 and 1994) and Levin and Lin (1992 and 1993) have significantly influenced the discussion of panel unit root tests for a panel of individuals $i = 1, \ldots, N$, where each individual contains $t = 1, \ldots, T$ time series observations. Quah (1992) proposed a panel version of the Dickey-Fuller test (*DF* test) without fixed effects.[6] Levin and Lin extended this test for fixed effects, individual deterministic trends and serially correlated errors. The resulting test is a panel version of the *DF*-test

$$\Delta y_{i,t} = \rho y_{i,t-1} + \alpha_{mi} d_{mt} + \varepsilon_{i,t}, \tag{9}$$

where d_m stands for the set of deterministic variables (fixed effects or joint intercept, individual deterministic trends and time dummies) with coefficient vectors α_m. Levin and Lin show that their test statistic (t-statistic) converges to standard normal distribution as $T \to \infty$, and $N \to \infty$ with $N/T \to 0$. However, it was found that the asymptotic mean and variance of the unit root test statistic vary under different specifications of the regression equation. Therefore, the majority of applications (see for

[4] The results of the Augmented Dickey Fuller test (*ADF* test) and the test according to Kwiatkowski et al. 1992) are available from the authors on request.

[5] Baltagi and Kao (2000) and Banarjee (1999) provide detailed surveys of panel unit root tests.

[6] This model specification corresponds fully to income convergence to the group's average analyzed in Quah's application. The test proposed by Quah (1992), however, is meant to be used in what he calls 'data fields,' that is, panels with large N and large T.

example Kočenda, 2001) used Monte Carlo simulations to compute critical values which corresponded fully to the analyzed panels. This also represented an important limit to general empirical applications.

Based on this criticism, Levin et al. (2002) proposed a new test (*LLC* test) based on orthogonalized residuals and the correction by the ratio of the long-run to the short-run variance of *y*. The calculation of the *LLC* test involves three steps. In the first step, two regressions are run to generate orthogonalized residuals

$$\Delta y_{i,t} = \sum_{l=1}^{P_i} \pi_{1,il} \Delta y_{i,t-l} + \alpha_{1,mi} d_{mt} + e_{i,t}\,, \tag{10a}$$

$$y_{i,t} = \sum_{l=1}^{P_i} \pi_{2,il} \Delta y_{i,t-l} + \alpha_{2,mi} d_{mt} + v_{i,t}\,, \tag{10b}$$

where d_m again stands for the set of deterministic variables with coefficient vectors α_1 and α_2 in the specifications (10a) and (10b), respectively. The lag order P_i, which may be different for individual cross-section units, is specified in individual *ADF* regressions

$$\Delta y_{i,t} = \delta_i y_{i,t-1} + \sum_{l=1}^{P_i} \theta_{il} \Delta y_{i,t-l} + \alpha_{mi} d_{mt} + \varepsilon_{i,t}\,. \tag{11}$$

The residuals from regressions (10a) and (10b) have to be normalized by regression standard errors estimated for (11) to control for heterogeneity between the panel units. These adjusted residuals, denoted by $\tilde{e}$ and $\tilde{v}$, are finally used to estimate the panel *t*-statistic as

$$\tilde{e}_{i,t} = \delta \tilde{v}_{i,t-1} + \tilde{\varepsilon}_{i,t}\,. \tag{12}$$

The conventional *t*-statistic for the coefficient δ has a standard normal limiting distribution if the underlying model does not include fixed effects and individual trends. Otherwise, this statistic has to be corrected using the first and second moments tabulated by Levin et al. and the ratio of the long-run variance to the short-run variance, which accounts for the nuisance parameters present in the specification. The limiting distribution of this corrected statistic is normal as $N \to \infty$ and $T \to \infty$, while $\sqrt{N}/T \to 0$ or $N/T \to 0$, depending on specified models. Furthermore, the Monte Carlo simulation shows that the test is appropriate also for panels of moderate size (*N* between 10 and 250 individuals and *T* between 25 and 250 periods), which are close to our panel.

The generality of the Levin-Lin type tests has made them a widely accepted panel unit root test. However, Levin and Lin have an important homogeneity restriction in their tests, namely the null assumes that $\rho_i = \rho = 0$ against the alternative $\rho_i < 0$ for all individual units *i*. As far as this result also reflects the possible speed of convergence, the Levin and Lin type tests are likely to reject the panel unit root.

Im et al. (2003) address this homogeneity issue, proposing a heterogeneous panel unit root test (*IPS* test) based on individual *ADF* tests. They propose average *ADF* statistics for fixed *T*, which is referred to as the $\tilde{t}-bar$ statistic

$$\tilde{t}-bar_{NT} = \frac{1}{N}\sum_{i=1}^{N}\tilde{t}_{iT} \ . \tag{13}$$

Furthermore, they show that this statistic can be normalized by tabulating the first two moments of the distribution of $\tilde{t}$. The resulting standardized $\tilde{t}-bar$ statistic, denoted by $Z_{\tilde{t}bar}$, has *N*(0,1) distribution as $T \to \infty$ followed by $N \to \infty$. By construction of the heterogeneous panel unit root test, the rejection of the null of panel unit root does not necessarily imply that the unit root is rejected for all cross-sectional units, but only for a positive share of the sample. The *IPS* test does not provide any guidance on the size of this subgroup.

Finally, Hadri (2000) presents an extension of the test of Kwiatkowski et al. (1992) to a panel with individual and time effects and deterministic trends (*PKPSS* test), which has as its null the stationarity of the series. Similarly to the time-series framework, the *PKPSS* test is based on a decomposition of cross-sectional series into the following components (for simplicity, we exclude the deterministic trend from the discussion here)

$$y_{it} = r_{it} + \varepsilon_{it}\,, \tag{14}$$

where the first term,

$$r_{it} = r_{it-1} + u_{it}\,, \tag{15}$$

is a random walk for cross-sectional units that is reduced to fixed effects under the null of stationarity. This implies that $\sigma_{ui} = 0$ under the null of stationarity. Following Kwiatkowski et al. (1992), Hadri defines Lagrange multiplier test (*LM*),

$$LM = \frac{\frac{1}{N}\sum_{i=1}^{N}\frac{1}{T^2}\sum_{t=1}^{T}S_{it}^2}{\hat{\sigma}_{\varepsilon}}\,, \tag{16}$$

where S_i is defined as the partial sum of the residuals in a regression of y on fixed effects.

$$S_{it} = \sum_{j=1}^{t} e_{ij} \text{ and } t = 1, 2, \ldots, T. \tag{17}$$

The denominator of the *LM* statistic is the long-run variance of the residuals, ε_{it}. Under no serial correlation, it can be estimated simply by the variance of the residuals from the *KPSS* equation. However, the long-run variance has to be estimated separately in the more common cases of serial correlation using a number (which can also be determined endogenously) of covariances of the residuals and their weights. Unfortunately, the outcome of the *KPSS* test may be relatively sensitive to this lag truncation. As in the previous tests, the panel version of the *KPSS* test can be normalized to $N(0,1)$ as $T \to \infty$ and $N \to \infty$.

In general, our estimates of the panel unit root tests confirm that the variables contain a unit root (see Table 1). The panel version of the *KPSS* is perhaps most clear-cut on this issue, as it rejects the null of stationarity for exchange rates money supply, real industrial production and the CPI - to - PPI ratio. A similar result applies to the *IPS* test, although there is some evidence with this test that the money supply is stationary when time dummies are not included. However, their inclusion would seem to be important for our sample, given the importance of events like the Russian crisis.[7] Although the *LLC* test produces a rejection of the unit root hypothesis for exchange rates and M2, as we have pointed out, the homogeneity assumption of this test means that its small sample properties are not as appealing as those of the other tests, and we therefore conclude that our variables are I(1).

[7] Backé and Fidrmuc (2000) find significant effects of the Russian crisis especially on Slovakia, Hungary and Poland.

Table 1: Panel Unit Root Tests, 1994:9-2002:3

	Exchange Rate	Money (M2)	Industrial Production	Interest Rates	Price Ratio (CPI to PPI)
IPS test	-0.928	-7.092***	-0.116	-0.131	0.608
IPS^{TD} test	0.595	-1.535	-0.367	-5.252***	-1.506*
LLC test	-2.512***	-7.516***	-0.189	0.361	-0.625
LLC^{TD} test	-3.187***	-3.360***	-0.354	-2.742***	-0.153
PKPSS test	14.301***	18.513***	10.361***	8.413***	14.509***
$PKPSS^{TD}$ test	15.136***	16.720***	13.243***	5.372***	6.207***

Note: *TD* denotes the inclusion of time dummies. *IPS* test with 2 lags (based on the maximum number of lags implied by SIC for the individual tests); *PKPSS* test with lag truncation of six lags. The panel includes the Czech Republic, Hungary, Poland, Romania, Slovakia and Slovenia. All explanatory variables are defined as deviation of individual countries from the euro area time series. All variables except interest rates are in logs. Variables are seasonally adjusted where necessary (money supply, industrial production). */**/*** denote significance at the 10%/5%/1% level.

5 Estimation of the Long-Run Monetary Model

The empirical work on exchange rate determination has been strongly influenced by Meese and Rogoff (1983), who compared the predictive abilities of a variety of exchange rate models. The key result of this paper was that structural models are generally not able to outperform simple naïve forecasts as made for example by a random walk. Although the subsequent research has produced some better results (see MacDonald and Taylor, 1993 and 1994), the generally accepted view is that (nominal) exchange rates cannot be robustly modeled in the short run. Furthermore, tests of purchasing power parity have also cast significant doubt on the behavior of real exchange rates (see Rogoff, 1996). However, new hopes emerged in the 1990s with the application of panel unit root tests and panel cointegration. Testing purchasing power parities for various panels has become one of the major application fields of these methods. Husted and MacDonald (1998) and Groen (2000) have shown that the monetary model has good in-sample properties in panel data sets for industrialized countries. Here we apply panel econometric methods to estimate the monetary model for a group of CEEC countries.

Following our discussion in Section 2, equation (8) may be expressed in a form suitable for econometric estimation as

$$s_{it} = \mu_i + \theta_t + \psi\left(m_{it} - m_t^*\right) - \delta\left(y_{it} - y_t^*\right) + \eta\left(i_{it} - i_t^*\right) - \pi\left(p_{it} - p_t^*\right) + \varepsilon_{it}, \qquad (18)$$

where m, y and i were defined before as money supply, output and interest rates. Price indices, p, are defined as differentials between CPI and PPI, and ε is the disturbance term. Various specifications of the model include fixed and/or time effects (denoted by μ and θ, respectively) or a common intercept. The coefficient of money supply, ψ, is expected to be close to unity, but we do not impose this condition in the estimations.

There appears to be a significant Balassa-Samuelson effect in the CEECs, corresponding to the catching-up process (see Égert, 2002, and MacDonald and Wójcik, 2003). The Balassa-Samuleson effect is proxied by including the ratio of consumer prices to producer prices into (18). If consumer prices are assumed to be a composite of tradable and nontradable prices, and producer prices areidentified with tradables, the ratio proxies the development of nontradable prices in the economy. As can be seen in Table 2, this variable has a very significant effect on nominal exchange rate in various specifications.

The previous section showed that the exchange rates and the right-hand side variables are I(1). Furthermore, the monetary model predicts that these variables should be cointegrated. Therefore, we consider several approaches to estimating the long-run (cointegrating) relationship between the variables. Kao and Chen (1995) show that the panel ordinary least squares (OLS) estimator is asymptotically normal, but it is still asymptotically biased. Although they propose a correction for this bias, it has been found that this correction does not tend to perform very well in reducing the bias in small samples. Therefore, some authors have proposed alternative methods of panel cointegration estimation.

Pedroni (1996 and 2001) proposes the fully modified OLS estimator (FMOLS), while Kao and Chiang (2000) recommend the dynamic OLS (DOLS). Pedroni's FMOLS corrects for endogeneity and serial correlation to the OLS estimator. Similarly, DOLS uses the future and past values of the differenced explanatory variables as additional regressors.

Kao and Chiang show that both estimators have the same (normal) limiting properties, although they are shown to perform differently in empirical analyses. The FMOLS does not improve the properties of the simple OLS estimator in finite samples.

Correspondingly, Baltagi and Kao (2000) consider DOLS to be more promising for the estimation of panel cointegration.

As an alternative to the previous methods, Pesaran et al. (1999) propose a pooled mean group estimator (PMGE). A particular advantage of the PMGE is that it also provides estimates of the short-run dynamics, which is ignored by simple OLS, FMOLS and DOLS.

The results for the individual estimators of the monetary model of exchange rates are listed in Table 2 with and without fixed effects and time dummies. Furthermore, we present a DOLS specification accounting for the contemporaneous correlation in the errors across countries by a seemingly unrelated regression (SUR). The long-run elasticities for the PMGE estimator corresponding to the columns PMGE and PMGE-T (including time dummies) are based on the estimates from a partial adjustment model of the type

$$\Delta s_{it} = \mu_i + \zeta_i \left[s_{it} - \psi\left(m_{it} - m_t^*\right) + \delta\left(y_{it} - y_t^*\right) - \eta\left(i_{it} - i_t^*\right) + \pi\left(p_{it} - p_t^*\right)\right] + \varepsilon_{it}, \quad (19)$$

where the correction to equilibrium (given by the parameter ζ) is allowed to differ across countries.[8] Furthermore, we also estimated the cross-section specific short-run dynamics (not reported in Table 2).

It can be seen that the basic features of the monetary model (the sign and absolute value range) are very robust to the estimation method. All variables have the correct signs and are highly significant. The performance of panel methods is much better than estimations using standard vector error correction models (available upon request).

The coefficient on the money supply term is close to unity in all specifications, with the exception of the estimates derived using the PMGE and FMOLS. Also, the effect of the interest rate is estimated uniformly between the various specifications. Although the uncovered interest parity condition does not seem to hold for the CEECs, the resulting effect of the interest rate remains very low. Given the definition of the interest rate and the fluctuation of the dependent variable, the interest rate has a negligible effect on exchange rates. As expected, real industrial production enters with a

[8] All estimates of ζ_i in the specification are negative and significant, providing evidence that the long-run equilibrium implied by the monetary model actually behaves like an attractor for nominal exchange rates.

negative sign. Although the coefficient is highly significant for all specifications, the DOLS specification with time dummies reduces the coefficient by one half, and both FMOLS specifications yield very low coefficients. By contrast, the coefficient on industrial production is close to –1 for the PMGE specification.

The price ratio is found to have a very important effect on the exchange rates. In the majority of specifications (DOLS and PMGE, but not FMOLS), the estimated elasticity is larger than one. Thus, a one percentage point increase in nontradable prices (consumer prices above producer prices) leads to a nominal exchange rate appreciation of about 1.5 percentage points, although the FMOLS estimates suggest a smaller slope of only 0.5 percentage points or even 0.2 percentage points. Thus, the DOLS estimates seem to be consistent with available estimates of the Balassa-Samuelson effect (see Halpern and Wyplosz, 2001, and Égert, 2003).

Finally, we test whether the estimated relationships are true cointegrating vectors in Table 3. Following the Engle and Granger's approach, Kao (1999) proposed several tests based on a homogenous panel version of the residual Dickey-Fuller test. First tests are based on a Dickey-Fuller-type equation for residuals estimated in the above specifications

$$\hat{\varepsilon}_{it} = \rho\hat{\varepsilon}_{it-1} + \nu_{it}\,, \tag{20}$$

where $\hat{\varepsilon}_{it}$ are residuals computed from the various specifications of (18) and (19). Kao's panel cointegration tests are based both on the autoregressive coefficient, ρ, (denoted by DF_ρ) and on the corresponding *t*-statistic (DF_t). Furthermore, they consider the endogeneity relationship between the regressors and residuals, which is adjusted by the long-run conditional variance of the residuals (see Kao et al., 1999). The corresponding test statistics for the autoregressive coefficients and the *t*-statistics are denoted by DF_ρ^* and DF_t^*, respectively.

Furthermore, Kao proposes a panel version of the residual *ADF* test based on

$$\hat{\varepsilon}_{it} = \gamma\hat{\varepsilon}_{i,t-1} + \sum_{j=1}^{p}\pi\Delta\hat{\varepsilon}_{i,t-j} + \nu_{it}\,. \tag{21}$$

The *ADF* test uses the *t*-statistic on the autoregressive coefficient, γ, which is again corrected for a possible endogeneity relationship between the regressors and the residuals.

Table 2: Panel Cointegration Estimation of the Monetary Model, 1994:9-2002:3

	OLS	FE	FE-T	FMOLS	FMOLS-T	DOLS	DOLS-T	DOLS-SUR	PMGE	PMGE-T
Money Supply	0.815	0.817	0.874	0.459	0.975	0.860	0.886	0.844	0.567	0.300
	(76.156)	(80.021)	(53.868)	(22.273)	(7.075)	(72.910)	(51.346)	(116.189)	(5.870)	(1.780)
Industrial Production	-0.403	-0.477	-0.329	-0.010	-0.074	-0.388	-0.250	-0.487	-1.106	-0.323
	(-10.390)	(-11.364)	(-6.888)	(-12.979)	(-14.632)	(-8.498)	(-4.713)	(-17.908)	(-2.914)	(-3.349)
Interest Rates	0.001	0.002	0.002	0.007	0.009	0.004	0.005	0.003	0.008	0.003
	(4.252)	(4.569)	(5.316)	(10.572)	(14.534)	(5.364)	(6.068)	(4.815)	(2.609)	(2.023)
Price Ratio	-1.843	-1.408	-1.405	-0.534	-0.199	-1.555	-1.632	-1.392	-1.049	-1.306
	(-18.471)	(-15.870)	(-11.351)	(-13.500)	(-8.480)	(-16.870)	(-11.334)	(-24.711)	(-2.3207)	(-3.861)
No. of obs. per country	91	91	91	91	91	91	91	91	91	91
Total no. of observations	546	546	546	546	546	546	546	546	546	546
Fixed effects	no	yes	yes	yes	yes	yes	yes	yes	yes	yes
Time effects	no	no	yes	no	yes	no	yes	no	no	yes

Notes: The panel includes the Czech Republic, Hungary, Poland, Romania, Slovakia and Slovenia. All explanatory variables are defined as deviation of individual countries from the euro area time series. All variables except interest rates are in logs. Variables are seasonally adjusted if necessary (money supply, industrial production). *t*-statistics are in parentheses. The PMGE column corresponds to the estimates of the long-run elasticities in a partial adjustment monetary model. The PMGE and PMGE-T columns correspond to the long-run elasticities in the error correction representation of an $ARDL(p_i, q_i, r_i, s_i)$ model for the nominal exchange rate, where the lag length is chosen through AIC.

Table 3: Residual Panel Cointgation Tests, 1994:9-2002:3

	OLS	FE	FE-T	FMOLS	FMOLS-T	DOLS	DOLS-T	DOLS-SUR	PMGE	PMGE-T
DF_ρ-Test	-2.890***	-2.949***	-2.261**	0.452	-3.093***	-3.682***	-3.842***	-3.366***	-2.807***	1.226
DF_t-Test	-2.290**	-2.352***	-1.616*	1.338	-2.506***	-3.128***	-3.296***	-2.795***	-2.202**	2.184
DF_ρ^*-Test	-8.317***	-8.391***	-7.177***	-2.464*	-8.459***	-9.569***	-9.777***	-9.080***	-8.129***	-1.159
DF_t^*-Test	-0.771	-0.884	-0.372	3.641	0.740	-1.190	-1.109	-1.092	1.055	5.016
Panel *ADF*-Test	-2.256**	-2.254**	-2.307**	-1.072	-2.992***	-2.737***	-3.045***	-2.451***	-2.033**	-0.679

Notes: See Table 2. */**/*** denote significance at the 10%/5%/1% level.

With the exception of the DF_t^* test, which is insignificant for all specifications, the remaining statistics show nearly the same picture.[9] On the one hand, the panel cointegration tests for DOLS, DOLS with time dummies and DOLS with SUR errors confirm the stationarity of the residuals. We should recall here that these specifications are also closer to the theoretical predictions on the coefficients than the other formulations. On the other hand, the tests reject a cointegrating relationship for fully modified OLS and pooled mean group estimators with time dummies. There are mixed results for the remaining specifications.

6 Equilibrium Exchange Rates in Selected Acceding Countries

In this section, we use the long-run relationship between exchange rates, money supply and industrial production to discuss the development of equilibrium exchange rates, which we define as potential levels corresponding to the development of money supply and real growth of industrial production in the EU and selected acceding countries. Given that our group of CEECs is still catching up, we also include a Balassa-Samuelson effect and the interest rate differential between the CEECs and the EU, although the latter is negligible.

Kim and Korhonen (2002) and Šmídková et al. (2002) discuss BEER and FEER models of exchange rates, respectively. A particular advantage of monetary models compared to these models is that the nominal exchange rate can be directly computed, both in sample and in out-of-sample forecasts. Thus, no further assumptions are necessary to derive nominal exchange rates which can be used for policy discussion. We do not discuss whether the monetary policy of our group of selected countries was in fact appropriate during the sample period.

The Maastricht exchange rate criterion of the Treaty on European Union foresees a participation in the ERM II of at least two years. Rahn (2003) argues that the current euro participants previously used ERM parities to determine the conversion rates for entry in monetary union. As a result, the setting of exchange rate parities in the ERM II possibly as soon as in the course of 2004 may have important long-run effects

[9] We used NPT 1.3 for panel cointegration tests (see Chiang and Kao, 2002), reflecting the comments on potential errors in this program by Hlouskova and Wagner (2003).

for the current accession countries. Čech et al. (2003) stress that an undervalued exchange rate may cause inflationary pressure in an economy, which could possibly delay the fulfillment of the inflation criterion. In contrast, overvalued exchange rates are likely to have deleterious effects on the competitiveness and prospects of real convergence.

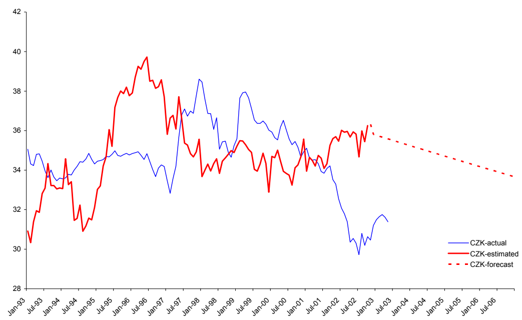
Thus, the path to the euro area should be optimally characterized by two features. First, realized exchange rates should be close to the equilibrium levels. As part of the convergence process, markets are likely to converge to the ERM II parities, which should be set at the appropriate equilibrium level. Second, the equilibrium exchange rates should be stable given reasonable expectations of the economy in the medium and long run, in order to avoid later exchange rate misalignments.

Within our sample (1994 to 2003), estimates show that, in general, the quality of the fit is relatively good given the standard of the exchange rate forecasts (see the example of DOLS-T specification in Figure 1). The market exchange rates have nearly always moved in the direction determined by the variables of the monetary model. The deviations between the predicted and the market exchange rates were relatively small during the whole analyzed period. However, we can see that the deviations have increased at the end of the sample. The Czech Republic appears to have a significant currency overvaluation of close to 15%. Our finding thus largely confirms earlier results e.g. by Šmídková et al. (2002).

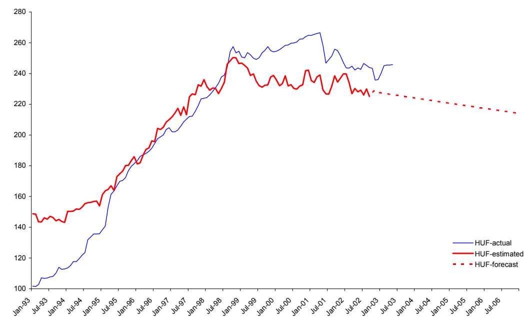
Finally, we simulate some possible trends of exchange rate development between 2003 and 2006 (see Figure 1), in order to evaluate the requirement of an early ERM II participation for various macroeconomic scenarios. As a result, we can see that there will be a slight tendency for exchange rate appreciation in the CEECs. Thus, the deviations from the equilibrium exchange rates as computed for 2002 will tend to decline during the next few years. Also, the equilibrium exchange rates of the CEECs display a relatively low variance. Actually, we can easily formulate an acceptable scenario for monetary policy if we keep the equilibrium exchange rates exactly constant.

Figure 1: Equilibrium Exchange Rates in CEECs, DOLS-T

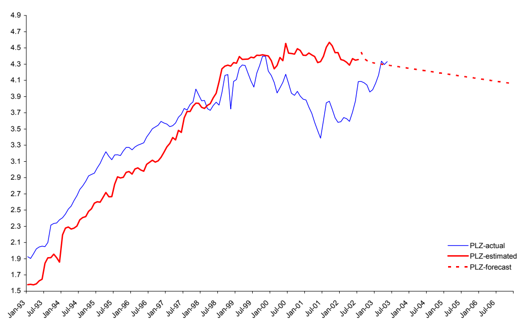
Czech Republic



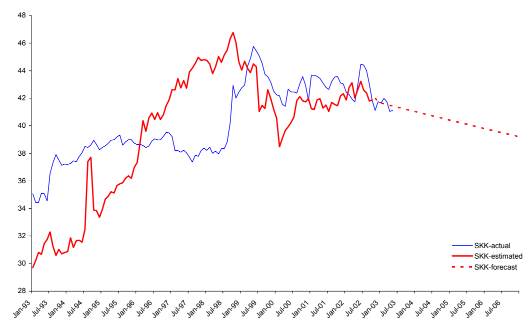
Hungary



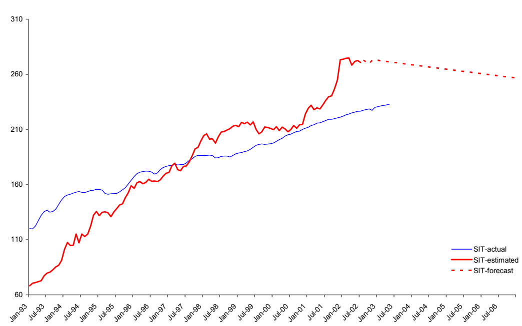
Poland



Slovakia



Slovenia



Note: The scenario for 2003 to 2006 assumes a growth differential for industrial production between the CEECs and the EU of 3 percentage points, a money supply growth differential of 3 percentage points and a rise in nontradable prices in comparison to the EU of 2 percentage points.

7 Conclusions

In this paper we have shown that the monetary model of exchange rate provides a relatively good explanation of the behavior of nominal exchange rates in a panel of six Central and Eastern European transition countries (Czech Republic, Hungary, Poland, Romania, Slovakia, and Slovenia) between 1994 and 2002. Since nominal exchange rates as well as our set of explanatory variables were found to be nonstationary, we use various panel cointegration estimators (OLS, DOLS, FMOLS, PMGE).

During the analyzed period, nominal exchange rates can be described mainly by the trend in money supply and real industrial production. We also find a significant Balassa-Samuelson effect, to which we can attribute about 2 or 3 percentage points of the annual exchange rate appreciation. This is comparable to the estimated effects available in the literature (see Halpern and Wyplosz, 2001, and Égert, 2003). We expect a decline of the Balassa-Samuelson effect after the accession to the EU. Although we find some evidence for interest rate determination of exchange rates, the size of this effect is generally not important.

Finally, we compute the equilibrium exchange rates based on monetary and real development in the CEECs and in the EU. For 2002, these results show that the nominal exchange rates against the euro could be overvalued to some degree especially in the Czech Republic and Slovenia.

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