BEYOND PURCHASING POWER PARITY:

Nominal exchange rates, output shocks and

non linear/asymmetric equilibrium adjustment in Central Europe

by

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Abstract

This paper models nominal exchange rates in the Czech Republic, Hungary, Poland, Slovakia and Slovenia, five Central European countries that recently joined the EU. We have five main findings. First, equilibrium nominal exchange rates are a function of domestic and EMU price levels, relative output shocks, nominal and real interest rate differentials and real net foreign assets. Second, the widely accepted Balassa-Samuelson effect is not universal. Third, there are non-linear and asymmetric short-run dynamics, where the speed of adjustment depends on both the size and sign of exchange rate misalignment. Fourth, estimated parameters differ between countries to such an extent that the use of panel data estimation can be questioned. Finally, we find little evidence of exchange rate misalignment in recent years, suggesting these countries may be able to meet the Maastricht criterion for exchange-rate stability.

Keywords: Nominal exchange rates, equilibrium, non-linear/asymmetric adjustment

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1. INTRODUCTION

Explaining the movements of nominal exchange rates in Central Europe over the past fifteen years presents a particular challenge to economic researchers. For the majority of the countries in the region, the introduction of market reforms in the early 1990s was accompanied by a pronounced inflation shock and structural changes in the supply side of these economies. Inflation has since been stabilised to single digit levels, however the process of structural change has continued at a high pace. Exchange rate developments have reflected the turbulence of these events. In most countries, the real exchange rate has appreciated over this period. The benchmark model of exchange rate determination, Purchasing Power Parity (PPP), predicts a constant real exchange rate, therefore it is an inadequate framework to describe exchange rate behaviour in Central Europe.

From the theoretical perspective, deviations from PPP occur when the domestic economy is subject to real output shocks relative to its main trading partners. Real and nominal exchange rates will appreciate permanently following positive and permanent relative demand shocks (Rogoff, 1996), of the type identified in Central European economies by Desai (1998) and Dibooglu and Kutan (2001). Exchange rates can also appreciate in the presence of Balassa (1964)–Samuelson (1964) effects arising from sustained increases in relative productivity in the tradeable sector, evidence for which has also been found (Halpern and Wyplosz, 1997, Grafe and Wyplosz, 1999, Krajnyak and Zettelmeyer, 1998, De Broeck and Slok, 2001, Taylor and Sarno 2001, Egert, 2002a, b, Egert and Lahreche-Revil, 2003, and Crespo-Cuaresma et al, 2004). However, the exchange rate will depreciate, if there are economy-wide supply shocks so the overall effect of permanent output shocks is a priori ambiguous.

The task of extracting the underlying equilibrium nominal exchange rate is made even more difficult by possible non-linearities in the process of the short-run exchange rate adjustment. Such non-linearities can be explained by theoretical models assuming limits to arbitrage, through spatially separated markets with transaction costs or sunk costs. In these models, periods of slow adjustment, when the rate of the misalignment of the nominal exchange rate relative to its equilibrium value is small, are followed by periods of much faster adjustment in response to more marked deviations of the exchange rate from the underlying equilibrium (see Dixit, 1989, Dumas, 1992, Uppal, 1993, Sercu et al., 1995, Shleifer and Vishny, 1997, and O'Connell, 1998). A number of empirical studies have found evidence of nonlinear adjustment in the nominal and real exchange rates of the G7 countries (Obstfeld and Taylor, 1997, Michael et al., 1997, Taylor and Peel, 2000, Taylor, Peel and Sarno, 2001, and Baum et al, 2001). In the context of transition economies, evidence of non-linear exchange rate adjustment has been provided by Taylor and Sarno (2001).

Overall, it is clear that empirical models aiming to explain nominal exchange rates in Central Europe must be able to capture changes in the underlying equilibrium nominal exchange by taking into account the full set of explanatory factors, while also modelling possibly nonlinear dynamic adjustment. This paper attempts to provide such an empirical investigation for the Central European countries that have recently joined the EU, namely the Czech Republic, Hungary, Poland, Slovakia and Slovenia.[1] We model the nominal exchange rates of the currencies of these countries against the Euro (the ECU prior to 1998). Studying the exchange rates of these countries is an issue of wider interest, as they all aspire to join the EMU in the foreseeable future, in which case choosing a suitable rate for the irrevocable conversion of their currencies is important

[1] The Baltic states (Estonia, Latvia and Lithuania) are excluded from the analysis due to the lack of sufficiently long data series.

both for these countries as well as for the European Central Bank which will conduct monetary policy on their behalf after accession. Depending on data availability, our sample covers the period from various points in the early 1990s to 2003. We estimate separate models for each exchange rate; this will enable us to capture country-specific elements in the behaviour of exchange rates and thus to test the adequacy of the assumptions underlying the widespread use of panel data estimation techniques (e.g. De Broeck and Slok, 2001, Egert et al, 2003, Crespo-Cuaresma et al 2003, Egert et al, 2004)

We use a behavioural model in which the equilibrium nominal exchange rate responds to a variety of factors including relative domestic and EU price levels, relative permanent output shocks, nominal and real interest rate differentials and net foreign assets. We put particular emphasis on how best to model output shocks. Many studies use productivity measures derived from the CPI-to-PPI ratio. In two recent papers, Egert et al (2003) and Egert et al (2004) argue this leads to overestimates of the Balassa-Samuelson effect, since it does not fully convey the effect of productivity gains and also reflects the effects of other factors omitted by these studies. This highlights the importance of using a properly specified model of the equilibrium exchange rate. To avoid these problems, we measure relative permanent output shocks using a statistical filter that separates underlying permanent changes in output from shorter-term temporary effects. This specification allows us to capture not only the effects of the CPI-to-PPI ratio but also a wider variety of output effects on nominal exchange rates.

Short-run exchange rate adjustment is modelled using a nonlinear dynamic model in which adjustment is allowed to vary between an inner regime, where the nominal exchange rate is closer to its equilibrium, and an outer regime, where exchange rate misalignment is more pronounced. This model has two main features. First, it allows the speed of adjustment to vary with the extent of nominal exchange rate misalignment. This aspect has been analysed in all non-linear models of exchange rate adjustment for the G7 countries and the ESTAR model used by Taylor and Sarno (2001) for transition economies. However, our non-linear model extends the existing literature by allowing exchange rates to respond differently to under-valuations and over-valuations of the nominal exchange rate. This type of asymmetry is quite plausible. Consider, for example, exchange rate intervention by a policymaker that assigns greater loss to employment being below the socially desirable level than to employment being too high (such a model has been analysed in a closed economy context by Cukierman and Gerlach, 2003). Such a policymaker may well be more responsive to exchange rate over-valuations than to under-valuations. Avoiding misspecification in the model of the equilibrium exchange rate is also important in the modelling of exchange rate dynamics, since inadequate models of the equilibrium exchange rate may induce spurious estimates of exchange rate dynamics.

We obtain a number of interesting empirical results. First, we find that with the possible exception of real net foreign assets, nominal exchange rates are affected in all countries by all the variables entering our equilibrium exchange rate equation. This is a reflection of the variety of shocks to which Central European countries have been subject to since the early 1990s. Second, the widely accepted Balassa-Samuelson effect is not universal. For the Czech Republic, Hungary and Slovakia, we find that an increase in permanent relative output results in exchange rate appreciation. This is consistent with the Balassa-Samuelson effect. But for Poland and Slovenia we find that relative permanent output gains result in exchange rate depreciation, which is consistent with supply shocks to the whole economy rather than to the tradeable sector alone. This finding overturns the findings of previous studies for these two countries, which may reflect the use of an inadequate measure of shocks or the inappropriate use of panel data techniques.

Third, we find significant evidence of non-linear exchange rate adjustment. In four out of the five countries examined, large exchange rate misalignments are found to be corrected at a rate almost three times faster than small misalignments. For the remaining country (Slovenia), error correction occurs only for large misalignments since the exchange rate is a random walk in the inner regime. Furthermore, we find that in three out of five countries (Czech Republic, Hungary and Slovakia), exchange rate adjustment is asymmetric, as the speed of adjustment in response to an overvaluation is different from that in response to an undervaluation. This is a new finding, which previous studies using symmetry-imposing non-linear models are not in position to capture. Fourth, there are significant differences in the estimated parameters between countries. This suggests the use of panel data estimation may be questionable. Finally, we find that at the end of our sample period (end of 2003), with the exception of the Polish zloty, Central European currencies were close to equilibrium. This is an indication that these countries are in favourable position to meet the exchange-rate stability criterion of by the Maastricht Treaty.

The remainder of the paper is structured as follows: Section 2 models the equilibrium nominal exchange rate. Section 3 uses the results obtained in section 2 to model short-run nominal exchange rate behaviour. In particular, section 3.1 estimates linear error-correction models; section 3.2 tests for the existence of non-linearities in movements of the nominal exchange rate; and section 3.3 estimates non-linear error-correction models of nominal exchange rate adjustment. Section 4 summarises and offers some concluding remarks.

2. MODELLING EQUILIBRIUM NOMINAL EXCHANGE RATES

The benchmark model of exchange rate determination, Purchasing Power Parity (PPP), postulates that nominal exchange rates are a function of the ratio of domestic to foreign price level. This is captured by equation (1) below, where s_t is the log of the market (observed) nominal exchange rate between the domestic currency and the Euro, p and p_{EMU} are the logs of the domestic and EMU average price levels respectively and ε_t a white noise error term.

$$s_t = \alpha + \beta_1 p_t + \beta_2 p_{EMUt} + \varepsilon_t \tag{1}$$

The absolute form of PPP postulates $\alpha = 0$ and $\beta_1 = -\beta_2 = 1$. Relative PPP allows for a non-zero constant, whereas measurement errors in price levels result in weak-form PPP, which only requires $\beta_1 > 0$ and $\beta_2 < 0$ (see Taylor, 1988). PPP captures the long-run exchange rate effects of monetary shocks but cannot capture the effects of permanent relative output shocks, such as those the Central European countries have experienced since the introduction of market reforms back in the early 1990s. Furthermore, PPP does allow for the effects of any other factors that may have an influence on exchange rate behaviour. For these to be captured, (1) should be extended as per equation (2) below, where $(\bar{y} - \bar{y}_{EMU})_t$ is domestic supply relative to that of the Eurozone area, $\mathbf{Z}_t$ is a $(k \times 1)$ vector of other variables relevant to exchange rate determination and B a $(1 \times k)$ vector of parameters.

$$s_t = \alpha + \beta_1 p_t + \beta_2 p_{EMUt} + \beta_3 (\bar{y} - \bar{y}_{EMU})_t + \mathrm{B}\,\mathbf{Z}_t + \varepsilon_t \tag{2}$$

Equation (2) is an equilibrium model of the exchange rate (see Stockman, 1980 and Lucas 1982), a simple version of which is discussed by Taylor (1995, pp. 24-26). Equilibrium exchange rate models consider two countries under full price flexibility, so $\bar{y}$ and $\bar{y}_{EMU}$ represent full-employment output levels. In this context, an increase in domestic productivity, captured by an increase in $(\bar{y} - \bar{y}_{EMU})_t$, has two analytically separate effects. The first, defined by Taylor as a relative price effect, involves a reduction in the relative price of domestic output, leading to a depreciation of s_t. The second, defined as a "money demand effect", will tend to appreciate s_t, as an increase in domestic supply increases the transactions demand for money, thus leading to a reduction in the equilibrium level of domestic prices. Whether the exchange rate rises or falls as a result of an increase in domestic productivity is ultimately a function of the degree of substitutability between domestic and foreign goods: an appreciation (depreciation) is more likely the higher (lower) the degree of substitutability, the smaller (higher) the relative price effect.

Nominal appreciation is also more likely if the domestic economy is subject to Balassa (1964) – Samuelson (1964) effects. These occur, in the context of a two-sector (traded / non-traded) economy, if there are productivity shocks specific to the traded sector. Through labour mobility, these shocks induce wage increases across the economy that create pressure for an appreciation. Since the effect of output shocks will reflect these offsetting pressures, there is no theoretical expectation for the sign of β_3.

We include three further explanatory variables in $\mathbf{Z}_t$. We use nominal interest rate differentials $(i\text{-}i_{EMU})_t$ to capture the effects of expectations of regarding future macroeconomic fundamentals and real interest rate differentials $(r\text{-}r_{EMU})_t$ since these affect capital inflows in Dornbuch (1976)-like sticky price models (see e.g. Hallwood and MacDonald 2000, pp. 175-209). We also include real net foreign assets $(rnfa)_t$ aiming to capture the wealth effects of accumulated current account imbalances. In portfolio balance models of the exchange rate (Branson, 1983), an increase in the stock of foreign assets held by domestic agents generates future capital inflows leading to an appreciation of the domestic currency. However, the negative relationship between

$(rnfa)_t$ and s_t may be reversed within the context of stock-flow models of the exchange rate (see Faruqee, 1995). As Egert et al (2004, p.20) argue, the relationship between $(rnfa)_t$ and the real exchange rate may be positive in the medium-run because the higher growth potential of transition economies cannot be financed by domestic savings only, but necessitates increased foreign borrowing leading to an increase in foreign liabilities. However, in the long-run, payments on the existing stock of foreign liabilities would restore the negative relationship: the higher the stock of foreign liabilities (i.e. the lower $(rnfa)_t$ is), the higher the need for real exchange rate depreciation to service the debt through an improved current account. [2]

Our model of the equilibrium nominal exchange rate is therefore

$$s_t = \alpha + \beta_1 p_t + \beta_2 p_{EMUt} + \beta_3 (\bar{y} - \bar{y}_{EMU})_t + \beta_4 (r - r_{EMU})_t + \beta_5 (i - i_{EMU})_t + \beta_6 rnfa_t + \varepsilon_t \quad (3)$$

Equation (3) resembles Frankel's (1979) "hybrid" monetary model equation, adjusted to allow for Balassa-Samuelson effects and the exchange rate effects of net foreign assets. From that point of view, it can be seen as a Behavioural Equilibrium Exchange Rate (BEERs) equation similar to those discussed by MacDonald (2000) where, and in line with our discussion above, we expect $\beta_1 > 0$, $\beta_2 < 0$, $\beta_4 < 0$, $\beta_5 > 0$, while β_3 and β_6 can take both a positive and a negative sign. The fitted values of (3), denoted by $\hat{s}*$, provide an estimate for a behavioural equilibrium value of the nominal exchange rate, while the estimated error term $\hat{\varepsilon}_t$ is measure of the deviation of the market nominal exchange rate s_t from its estimated equilibrium $\hat{s}*$.

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[2], Rogoff (1996, p. 663) eloquently describes the ambiguity discussed above by stating that, "from a theoretical perspective, virtually any correlation between the current account and the real exchange rate can be easily rationalise. Ultimately, the correlation between the current account and the real exchange rate is an empirical matter, one that remains the subject of debate".

We estimate (3) using monthly data for nominal exchange rates against the ECU (the Euro since 1999), taken by the Eurostat database provided by Datastream. Domestic and EMU price levels are approximated by the series of producer price index taken from the International Financial Statistics (IFS) databank provided by the same source.[3] We use an index of industrial production as the best available measure of output. To obtain a measure of ($\bar{y}$ - $\bar{y}_{EMU}$), we fit in the domestic and Eurozone output series a Hodrick-Prescott filter (Hodrick and Prescott, 1997)[4]; we then define the fitted trend values to be equal to $\bar{y}$ and $\bar{y}_{EMU}$ respectively.[5] Nominal interest rates, i_t, are approximated by the series on central bank discount rates provided by the IFS databank.[6] Finally, real net foreign assets, which are expressed in terms of log-levels, have been calculated using the nominal net foreign asset series provided by IFS.[7]

Data availability allows for the following sample periods: 1993(2)-2003(11) for the Czech Republic; 1992(1)-2003(10) for Hungary; 1991(6)-2003(9) for Poland; 1993(3)-2004(10) for Slovakia; and 1992(3)-2003(11) for Slovenia.[8] Augmented

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[3] The PPI index for the EMU average is taken by the Eurostat Database available by Datastream. The Slovenian PPI index is taken by the Vienna Institute for International Economic Studies (WIIW) database provided by the same source.

[4] To obtain the Hodrick-Prescott trend, the investigator must specify the value of a smoothing parameter that regulates the smoothness of the series; we use the recommended value for monthly data of 14400, but we investigate the effect of other values of the smoothing parameter below.

[5] Working with finite date sets implies that in addition to permanent relative supply shocks, ultimately transitory yet relatively persistent demand shocks may also have an effect on the fitted output series. This implies that our permanent output component may capture the effects of both permanent supply and demand shocks dying out at a slow rate. As a highly persistent positive demand shock would result in real appreciation for the domestic country, this increases the possibility of obtaining a negative coefficient for β_3.

[6] We would have preferred to use a series on long-term government bond yields, or a Treasury Bill rate, however time series for these variables were not available for the whole of the sample periods considered by our analysis. For the pre-Euro period, the EMU discount rate is approximated by the Germany discount rate. This assumption is plausible, given that Germany is the single most important trading partner of the countries examined by our analysis, and also due to Germany's central role in overall exchange rate developments in Europe throughout the 1990s. Real interest rates were calculated using the formula $1+r = (1+i)/(1+\pi)$, where π is the rate of PPI inflation over a period of twelve months,

[7] We would have preferred to measure net foreign assets as a percentage in GDP. However, due to the lack of monthly-frequency data on the GDP series of the countries considered by our analysis, this was not possible.

[8] The starting points of our samples are justified as follows: for the Czech Republic and Slovakia, February 1993 is the month of the monetary "divorce" between these countries and the introduction of two separate national currencies. For Poland and Hungary, although data availability goes back to 1990,

Dickey-Fuller (ADF) tests (Dickey and Fuller, 1979) [9] show that all variables entering equation (3) are non-stationary.[10] We exploit non-stationarity by estimating (1) as a cointegrating relationship. Following, Michael et al (1997), we estimate (3) using the Engle-Granger (1987) methodology.

The empirical results are reported in Table 1.[11] The ADF tests show that the estimated equation residuals are stationary, implying that all models reported in Table 1 are cointegrated. Our empirical findings confirm our expectations for the variables whose sign is theoretically unambiguous: the price level is statistically significant with β_1 being positive and β_2 negative. Real and nominal interest rate differentials are also in all (but one) cases statistically significant and display the expected negative and positive sign for β_4 and β_5 respectively. Permanent shocks to relative output are significant in all cases but of varying sign. For the Czech Republic, Hungary and Slovakia, β_3 is negative. This suggests that the Balassa-Samuelson and money demand effects dominate, particularly in the cases of the Czech Republic and Slovakia. By contrast, for Poland and Slovenia β_3 is positive, suggesting that the relative price effect discussed by Taylor (1995) prevails. This is a new finding that supports the arguments of Egert et al (2003) and Egert et al (2004) that Balassa-Samuelson effects have been overstated. Finally, in four out of five countries, net foreign assets are reported to be statistically insignificant. For the remaining country, Hungary, in line with the stock-flow argument discussed above, β_6 has a small positive sign,

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we chose to start our analysis later, as the initial phases of transition in these countries were accompanied by wild fluctuations in interest rates and real exchange rates. In Hungary, these fluctuations became less dramatic towards the end of 1991 whereas in Poland relative stability was introduced following the devaluation of the zloty in May 1991. Finally, our sample for Slovenia starts in March 1993 as in February 1993 this country's currency, the tolar, was de-linked from the Yugoslav dinar and was devalued against the German Mark by 20 per cent.

[9] These tests are not reported but are available by the authors upon request.

[10] Neither these tests, nor subsequent estimates of cointegrating relationships are affected by non-linearities (Michael, et al, 1997).

Overall, our findings suggest that nominal exchange rates in Central Europe over the last fifteen years have been influenced by both monetary and output shocks, as well as additional factors captured by the nominal and real interest rate differentials. This is a reflection of the variety of shocks these countries have been subject to following the introduction of market reforms in the early 1990s. They also suggest that despite the existence of strong similarities, each of the countries examined maintains idiosyncratic elements in the process of exchange rate determination.

In particular, and with regards to output shocks, our findings show that these affect nominal exchange rates in different countries in varying forms and at different degrees. For the Czech Republic, Hungary and Slovakia, relative output gains lead to currency appreciation, suggesting that the Balassa-Samuelson and money-demand effects discussed above outweigh the relative price effect in these countries. By contrast, for Poland and Slovenia, the relative-price effect appears to be stronger, leading relative output gains to result in nominal exchange rate depreciation. From a policy-making perspective, this suggests that costs of EMU entry may be larger for Poland and Slovenia than for the remaining countries since the adverse effects of asymmetric relative output shocks may be larger.

3. MODELLING SHORT-RUN NOMINAL EXCHANGE RATE ADJUSTMENT

3.1. Linear short-run adjustment models

In this section we model the short-run adjustment of the nominal exchange rate towards its equilibrium level described by (3). We start by estimating a standard error-correction equation given by (4) below:

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[11] As a robustness check, we experimented with models in which the smoothing parameter on the Hodrick-Prescott filter was adjusted by up to 50% in either direction. This made little difference to our results.

$$\Delta s_t = \beta(L)\,\Delta\, s_{t-1} + \gamma(L)\,\Delta\, \hat{s}^*{}_t + \delta(s - \hat{s}^*)_{t-1} + \varepsilon_t \tag{4}$$

In (4), s is the log of the actual (observed) nominal exchange rate, $\hat{s}^*$ is the log of the derived equilibrium nominal exchange rate, $\beta(L)$ and $\gamma(L)$ are polynomials in the lag operator, L, ε is a white noise error term and Δ is the first difference operator. The mechanism through which the actual nominal exchange rate converges to its equilibrium value is the error correction term $(s-\hat{s}^*)_{t-1}$, which measures the deviation of the nominal exchange rate from its equilibrium value. According to the Granger representation theorem, if this is statistically significant, there exists a cointegrating relationship between nominal exchange rates and right-hand side terms in equation (3), with δ measuring the speed of adjustment towards equilibrium.[12]

Table 2 presents the estimates of the linear error correction equations in (4). We report estimates of parsimonious models obtained using a general-to-specific specification search on a baseline model using eighteen lags of all variables.[13] For all countries the error correction term is statistically significant at the 5 per cent level, further confirming that we have estimated cointegrating relationships. The estimated error correction terms imply fast reversion to equilibrium in the Czech Republic and a moderate speed of adjustment for the remaining countries. The equations are generally well-specified, but for some countries (especially for Poland and Slovakia) their explanatory power is rather low. This may reflect the well-established stylised fact of high short-run volatility in the movements of nominal exchange rates. However, it may

[12] An alternative modelling approach would be to substitute (1) into (2) and estimate the resulting equation $\Delta s_t = \beta(L)\,\Delta\, a_{t-1} + \gamma(L)\,\Delta\,(\pi' z_t) + \delta(a - \pi' z)_{t-1} + \varepsilon_t$, with z'_t defined as $[p_t, p_t^*, (\bar{y} - \bar{y}_{EMU})_t, (y^T - y^T{}_{EMU})_t]$. We prefer equation (2) to this alternative because it requires estimation of a smaller number of parameters, an important consideration when estimating non-linear models using relatively short samples.

[13] The equations reported in Table 2 include intercept dummies for periods of particular turbulence (e.g. devaluations and changes in monetary policy framework). For each equation we investigated the significance of a number of dummies corresponding to such events and kept those that proved to be statistically significant. The dates for which these dummies are defined appear in the note accompanying

also be due, at least to some extent, to the existence of non-linearities in the process of nominal exchange rate determination. We test this hypothesis formally in the following section.

3.2. Linearity tests

The hypothesis of linear adjustment can be tested using the procedure described in Saikonnen and Luukkonen (1988), Luukkonen et al (1988), Granger and Teräsvirta (1993) and Teräsvirta (1994). More specifically, to test linearity we estimate

$$(s-\hat{s}^*)_t = \gamma_{00} + \sum_{j=1}^{\phi}\{\gamma_{0j}(s-\hat{s}^*)_{t-j} + \gamma_{1j}(s-\hat{s}^*)_{t-j}(s-\hat{s}^*)_{t-d} + \gamma_{2j}(s-\hat{s}^*)_{t-j}(s-\hat{s}^*)^2_{t-d}$$
$$+\gamma_{3j}(s-\hat{s}^*)_{t-j}(s-\hat{s}^*)^3_{t-d}\} + \gamma_4(s-\hat{s}^*)^2_{t-d} + \gamma_5(s-\hat{s}^*)^3_{t-d} + v(t) \tag{5}$$

In (5), $(s\text{-}\hat{s}\,^*)_t$ is the deviation of the nominal exchange rate from its estimated equilibrium, measured by the residual term obtained from (3), d is the delay parameter of the transition function to be used and $v(t) \sim niid\ (0,\sigma^2)$. Linearity implies the null hypothesis H_0: $[\gamma_{1j} = \gamma_{2j} = \gamma_{3j} = \gamma_4 = \gamma_5 = 0]$ for all $j \in (1,2...\phi)$. This can be tested using an LM-type test. Having determined ϕ through inspection of the partial autocorrelation function,[14] (5) can be estimated for all plausible values of the delay parameter d. The correct value of d is that which yields the largest value of the test statistic.

Table 3 presents the results of our non-linearity tests. In all cases we reject the null hypothesis of linear adjustment at the 5% level or better and conclude that the short-run adjustment of the nominal exchange rate to its equilibrium level is a non-linear process. We proceed to model this non-linearity formally below.

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the Table. The omission of these dummies does not change the nature of the results but results in problems of residual non-normality.

3.3. Non-linear short-run adjustment models

We model the non-linear adjustment of the nominal exchange rate identified in the previous section using the Quadratic Logistic Smooth Transition Error Correction Model (QL-STECM). This is described by equations (6) to (9) below: [15]

$$\Delta s_t = \theta_t M_{It} + (1-\theta_t)\, M_{Ot} \tag{6}$$

$$M_{It} = \beta_I (L)\, \Delta s_{t-1} + \gamma_I (L)\, \Delta \hat{s}^* + \delta_I (s-\hat{s}^*)_{t-1} + \varepsilon_{It} \tag{7}$$

$$M_{Ot} = \beta_O (L)\, \Delta s_{t-1} + \gamma_O (L)\, \Delta \hat{s}^* + \delta_O (s-\hat{s}^*)_{t-1} + \varepsilon_{Ot} \tag{8}$$

$$\theta_t = pr\ \{ \tau^L \leq (s-\hat{s}^*)_{t-d} \leq \tau^U \} = 1 - \frac{1}{1+e^{-\sigma[(s-\hat{s}^*)_{t-d}-\tau^L][(s-\hat{s}^*)_{t-d}-\tau^U]}} \tag{9}$$

Equation (6) models exchange rate changes as a weighted average of the linear models M_I and M_O, where M_I represents the inner regime and M_O the outer regime. Equations (7) and (8) describe M_I and M_O as linear error-correction models, similar to (2). Equation (9) specifies the regime weight θ as the probability that the transition variable $(s-\hat{s}^*)_{t-d}$ lies within the "regime boundaries" τ^L and τ^U, where the probability is described using a quadratic logistic function. We expect $\tau^L < 0$ and $\tau^U > 0$. Nominal exchange rates are mainly determined by M_I (the inner regime) when the nominal exchange rate is close to its equilibrium value described by (3) and mainly by M_O (outer regime) in periods of significant exchange rate misalignment, with σ denoting the speed of transition between the two regimes.

The speed of adjustment of the exchange rate differs between regimes if $\delta_I \neq \delta_O$. If $\delta_I = 0$ and $\delta_O < 0$, the nominal exchange rate only adjusts towards its fundamental

[14] Granger and Teräsvirta (1993) and Teräsvirta (1994) advise against choosing ϕ using an information criteria such as the Akaike since this may induce a downward bias. The estimated Partial Autocorrelation functions underlying the results reported in Table 3 are available upon request.
[15] For a detailed discussion of this model see van Dijk et al. (2002)

value in the outer regime, evolving as a random walk in the inner regime. If $\tau^U + \tau^L = 0$, the model is in effect equivalent to the ESTAR model, similar to the one used by Sarno and Taylor (2001), in which the speed of adjustment depends only on the size of the deviation of exchange rates from fundamentals. If, on the other hand, $\tau^U + \tau^L \neq 0$, the model is more general than the ESTAR model since the speed of adjustment depends both on the size and on the sign of the deviation from equilibrium. In particular, if $\tau^U + \tau^L > 0$, the nominal exchange rate responds more vigorously to under-valuations, while $\tau^U + \tau^L < 0$ indicates a stronger response to over-valuations.

Table 4 presents the estimates of our non-linear QL-STECM models. The reported equations are again obtained using a general-to-specific specification search.[16] The econometric properties of the models reported in Table 4 are generally superior to their linear counterparts in Table 2, as they all pass the misspecification tests at the 5 per cent level and produce lower regression standard errors.[17] We obtain significant differences between the speeds of adjustment in the inner and the outer regime. In four out five cases (the Czech Republic, Hungary, Poland and Slovenia), exchange rates revert to the estimated equilibrium in both regimes but where the estimated speed of reversion in the outer regime is almost three times higher than that of the inner. In the remaining country (Slovakia), the error correction coefficient is insignificant in the inner regime, implying that the exchange rate is a random walk in the inner regime, but significant in the outer regime. Finally, the symmetry restriction $\tau^U + \tau^L = 0$ is rejected in three out of five cases (Czech Republic, Hungary and Slovakia), with the difference between the two threshold values being more pronounced in the Czech Republic and

[16] These models include the statistically significant crisis dummies included in their linear counterparts mentioned in the notes accompanying Table 2. This ensures that our non-linear findings apply to the whole of our samples and do not simply pick up the influence of these one-off events.

[17] However, in all cases σ is imprecisely estimated as the likelihood function is very insensitive to this parameter (see the detailed discussion on this point in van Dijk et al., 2002).

Slovakia. This asymmetry cannot be captured by the TAR or ESTAR models that are extensively used in the literature.

Figure 1 presents the estimated nominal exchange rate misalignment term against the estimated inner regime thresholds for each of the countries examined. Three interesting observations emerge: First, nominal exchange rate misalignments are normally low and fall within the inner regime, although larger misalignments that fall in the outer regime are not uncommon, particularly in the Czech Republic and Slovakia. Second, most of the major discrete nominal devaluations in the 1990s were preceded by exchange rate overvaluations that were large enough to fall outside the inner regime.[18] Finally, we observe that at the end of 2003, Central European nominal exchange rates were close to their equilibrium values, suggesting that these countries are in a favourable position to meet the criterion of exchange rate stability necessary for gaining access to the EMU within the foreseeable future. The only exception appears to be Poland, whose currency appears to have been undervalued by approximately 4 percent, placing the misalignment term in the outer regime.

4. SUMMARY AND CONDLUDING REMARKS

This paper has examined nominal exchange rate behaviour in the five Central European countries that have recently joined the EU, namely the Czech Republic, Hungary, Poland, Slovakia and Slovenia, against the Euro (the ECU prior to 1998). We modelled equilibrium nominal exchange rates used a behavioural equilibrium model. We found that their exchange rates of Central European countries are explained by a variety of factors, including movements in relative prices, relative output shocks, as

[18] Dates of major devaluations within our sample periods are May 1997 for the Czech Republic; August 1994 and May 1995 for Hungary (for this country the early 1990s saw a series of small pre-announced devaluations, typically within the range of 2 to 3 percent); February 1992 and August 1993 for Poland; and July 1993 for Slovakia.

well as nominal and real interest rate differentials. We also find that individual countries maintain elements of idiosyncratic behaviour in the process of exchange rate determination, some of which overturn the findings of previous studies according to which Balassa-Samuelson effects have taken place in all Central European countries. Finally, we obtain strong evidence of non-linearity, with the speed of nominal exchange rate adjustment being in all cases a function of the size, and in three countries (the Czech Republic, Hungary and Slovakia) the sign of exchange rate misalignment.

Our work can be extended towards a number of directions. In the empirical level, our methodology could be applied to explain exchange rate behaviour in a number of emerging economies, where the process of transition has not been as rapid as in the case of the countries examined here. In the theoretical level, it would be useful to develop a formal model of non-linear/asymmetric exchange rate behaviour, perhaps drawing on the recent literature on non-linear policy rules, which would provide a clearer theoretical grounding for empirical work in this area.

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Table 1

Long-run models

	Czech Republic	Hungary	Poland	Slovakia	Slovenia
constant	3.577	2.448	3.458	2.553	3.942
	(0.430)	(0.453)	(0.586)	(0.492)	(0.240)
p	0.692	0.996	1.015	0.910	1.438
	(0.119)	(0.044)	(0.112)	(0.077)	(0.047)
p_{EMU}	-1.693	-1.182	-2.417	-1.381	-2.491
	(0.291)	(0.232)	(0.241)	(0.293)	(0.158)
$(\bar{y} - \bar{y}_{EMU})$	-0.728	-0.167	0.498	-0.723	0.528
	(0.068)	(0.091)	(0.350)	(0.093)	(0.100)
$(r\text{-}r_{EMU})$	-0.564	-0.061	-0.407	-0.229	-0.531
	(0.084)	(0.050)	(0.092)	(0.064)	(0.036)
$(i\text{-}i_{EMU})$	0.489	0.308	0.483	0.406	0.397
	(0.061)	(0.040)	(0.121)	(0.152)	(0.047)
rnfa	-0.009	0.088	-0.015	0.001	-0.012
	(0.015)	(0.013)	(0.053)	(0.006)	(0.014)
rnfl [1]				-0.002	
				(0.007)	
R^2	0.75	0.99	0.98	0.79	0.99
Cointegration ADF	-6.08**	-4.98**	-3.74**	-4.59**	-5.46**

Notes: Standard errors in parentheses; Critical values for cointegration ADF test: 95 per cent: -2.88; 99 per cent: -3.48.

[1] For a short period of time in the early 1990s, Slovakia is recorded by IFS to have had a negative net foreign assets position. As a result, we define two variables relating to the foreign assets of this country: real net foreign assets (*rnfa*), is defined as the log of net foreign assets for periods of positive asset position, zero otherwise; and real net foreign liabilities (*rnfl*), which is defined the log of the absolute value of net foreign liabilities for periods of negative asset position; zero otherwise.

Table 2

Linear Error Correction Models

	Czech Republic	Hungary	Poland	Slovakia	Slovenia
constant	0.000 (0.0008)	0.000 (0.0009)	0.0009 (0.0011)	0.0005 (0.0007)	0.0005 (0.0002)
Δs_{t-1}					
Δs_{t-2}				0.152 (0.083)	
Δs_{t-3}	0.121 (0.062)				
Δs_{t-7}					0.230 (0.044)
Δs_{t-8}					-0.153 (0.052)
Δs_{t-9}					0.098 (0.044)
Δs_{t-14}					-0.091 (0.036)
Δs_{t-17}		0.195 (0.075)			
Δs_{t-18}					-0.068 (0.031)
Δs^*_t		0.543 (0.156)	0.771 (0.206)		0.253 (0.039)
Δs^*_{t-1}	-0.978 (0.371)				
Δs^*_{t-2}	1.262 (0.375)				-0.151 (0.058)
Δs^*_{t-3}					-0.106 (0.042)
Δs^*_{t-9}					0.130 (0.042)
Δs^*_{t-10}	0.573 (0.257)				
Δs^*_{t-14}					0.180 (0.047)
$(s-s^*)_{t-1}$	-0.382 (0.073)	-0.206 (0.053)	-0.174 (0.050)	-0.262 (0.057)	-0.191 (0.024)
R^2	0.62	0.50	0.19	0.16	0.79
Regression Std Error	0.00876	0.00850	0.01177	0.00858	0.00153
F-AR	0.66 [0.71]	1.46 [0.19]	0.43 [0.88]	0.89 [0.52]	0.51 [0.82]
F-ARCH	0.67 [0.70]	1.14 [0.34]	0.68 [0.69]	0.87 [0.53]	2.384 [0.03]
Chi sq. Normality	5.73 [0.06]	0.82 [0.67]	3.37 [0.19]	3.10 [0.21]	3.49 [0.17]
F-Het	0.56 [0.91]	1.23 [0.28]	0.61 [0.69]	0.15 [0.96]	0.703 [0.85]
RESET	0.01 [0.92]	1.03 [0.31]	0.17 [0.68]	1.59 [0.21]	2.59 [0.11]

NOTES: All models have been estimated including dummy variables for periods of particular exchange rate turbulence. These are defined: For the Czech Republic 1997-May, 1998- Aug, 1999-Jan, 2002-June, 2002-July, 2003-May, 2003-June. For Hungary: 1994-Aug, 1995-Mar, 2001-June, 2003-Jun. For Poland: 1998-Aug. For Slovenia, 1995-Feb, 1995-March, 1996-July, 1998-April. F-AR is the Lagrange Multiplier F- test for residual serial correlation of up to seventh order. F-ARCH is an F-test for Autoregressive Conditional Heteroskedasticity and general mispecification. Chi-sq. normality is a Chi-square test for residuals' normality. F-Het is an F-test for residuals heteroskedasticity. F-RESET is an F-test for functional form. The numbers reported in square brackets are p-values.

Table 3

Linearity tests

	ϕ	d	F-test [p-value]
Czech Republic	1	11	5.322 [0.00]
Hungary	1	6	4.385 [0.01]
Poland	1	15	5.067 [0.00]
Slovakia	1	1	3.942 [0.01]
Slovenia	2	14	3.081 [0.00]

NOTES: ϕ is the order of the autoregressive component and d the order of the delay parameter in the artificial regression given by equation (5).

Table 4

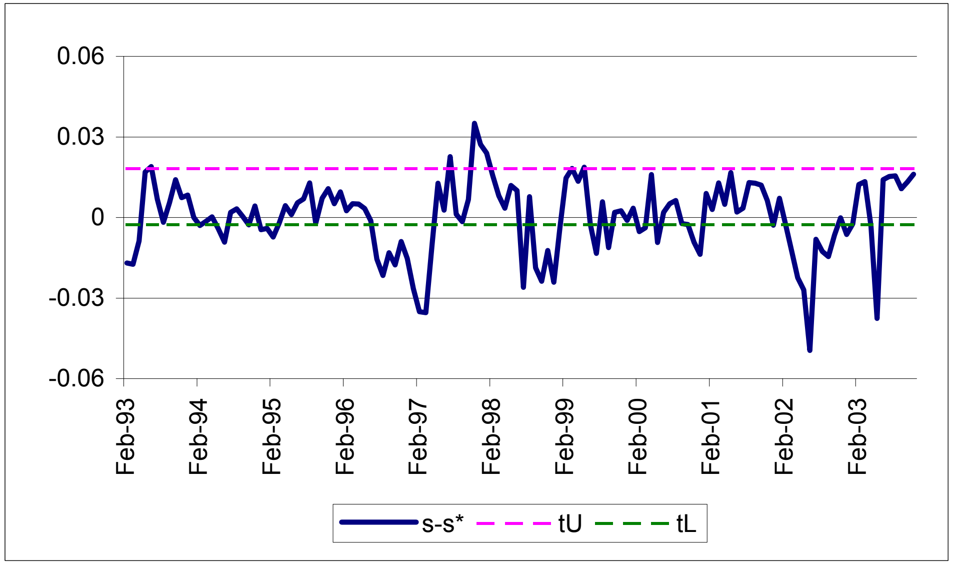
Nonlinear Error Correction Models

	Czech Republic	Hungary	Poland	Slovakia	Slovenia
M_I (Inner regime)					
Constant	-0.0015 (0.0011)	-0.0007 (0.0009)	-0.001 (0.001)	0.0027 (0.0013)	0.0005 (0.0002)
Δs_{t-1}	-0.358 (0.105)				0.359 (0.079)
Δs_{t-8}		0.160 (0.080)			
Δs_{t-11}		0.169 (0.088)	0.745 (0.249)		
Δs_{t-17}		0.164 (0.073)			
Δs^*_t		0.478 (0.156)			0.193 (0.055)
Δs^*_{t-2}	1.063 (0.505)				
Δs^*_{t-3}	1.189 (0.467)				
Δs^*_{t-7}	-0.911 (0.472)				
$(s-s^*)_{t-1}$	-0.205 (0.106)	-0.191 (0.053)	-0.167 (0.059)	0.039 (0.200)	-0.089 (0.033)
M_O (Outer regime)					
constant	0.0032 (0.0015)	0.0102 (0.003)	0.0044 (0.0026)	-0.0008 (0.011)	0.0003 (0.0004)
Δs_{t-1}		-0.844 (0.329)			0.484 (0.077)
Δs_{t-2}				0.281 (0.130)	
Δs_{t-6}					0.224 (0.055)
Δs_{t-10}					0.148 (0.050)
Δs_{t-14}	0.172 (0.093)				
Δs_{t-15}			-0.481 (0.158)		
Δs^*					0.207 (0.0600
Δs^*_{t-3}					-0.189 (0.066)
Δs^*_{t-5}			1.292 (0.530)		
Δs^*_{t-9}					0.168 (0.065)
Δs^*_{t-10}	1.307 (0.572)				
Δs^*_{t-18}					-0.151 (0.047)
$(s-s^*)_{t-1}$	-0.541 (0.108)	-0.536 (0.205)	-0.430 (0.130)	-0.300 (0.063)	-0.226 (0.036)
σ	152.1 (1268)	50.0 (158.6)	10.8 (18.0)	19.01 (52.58)	17.64 (37.16)
τ^U	0.018 (0.006)	0.022 (0.0002)	0.0241 (0.0052)	0.007 (0.001)	0.0066 (0.0009)
τ^L	-0.003 (0.0007)	-0.027 (0.0004)	-0.0235 (0.0018)	-0.016 (0.008)	-0.0080 (0.0009)
R^2		0.56			0.80
Regression Std Error	0.00859	0.00807		0.00842	0.00148
F-AR	0.66 [0.70]	1.35 [0.24]	0.45 [0.87]	0.66 [0.70]	1.27 [0.27]
Chi sq. Normality	0.81 [0.67]	0.77 [0.68]	2.91 [0.23]	3.10 [0.21]	5.54 [0.07]
F-Het	0.85 [0.68]	0.98 [0.49]	0.80 [0.70]	0.30 [0.99]	0.83 [0.72]
F-Test H_O: $\tau^L + \tau^U = 0$ (95 pc CV: 3.84]	7.435	16.198	1.432	9.138	3.45

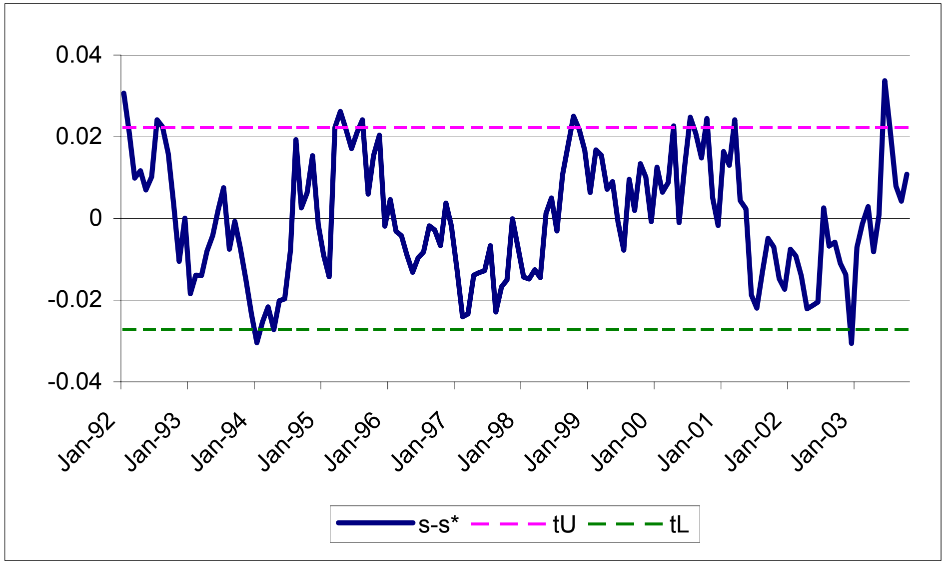
NOTES: All models have been estimated including dummy variables for periods of particular exchange rate turbulence. These are defined: For Czech Republic 1998-July, 2002-June, 2003-May, 2003-June, 2003-July. For Hungary: 1994-Aug, 1995-Mar, 2001-June, 2003-Jun. For Poland: 1998-Aug. F-AR is the Lagrange Multiplier F- test for residual serial correlation of up to seventh order. Chi-sq. normality is a Chi-square test for residuals' normality. F-Het is an F-test for residuals heteroskedasticity. The numbers reported in square brackets are p-values.

Figure 1: Nominal exchange rate misalignment against inner regime thresholds

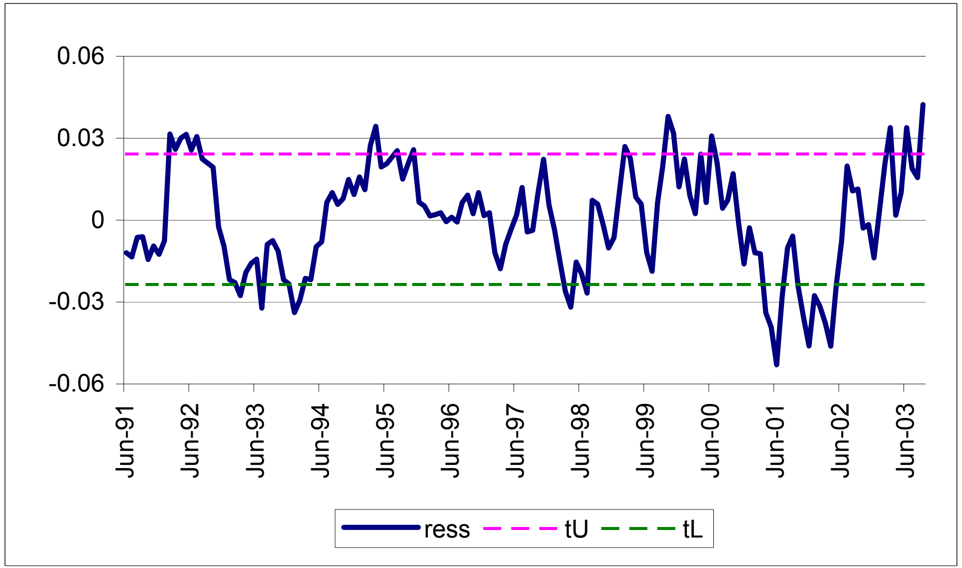
Czech Republic



Hungary



Poland



Slovakia

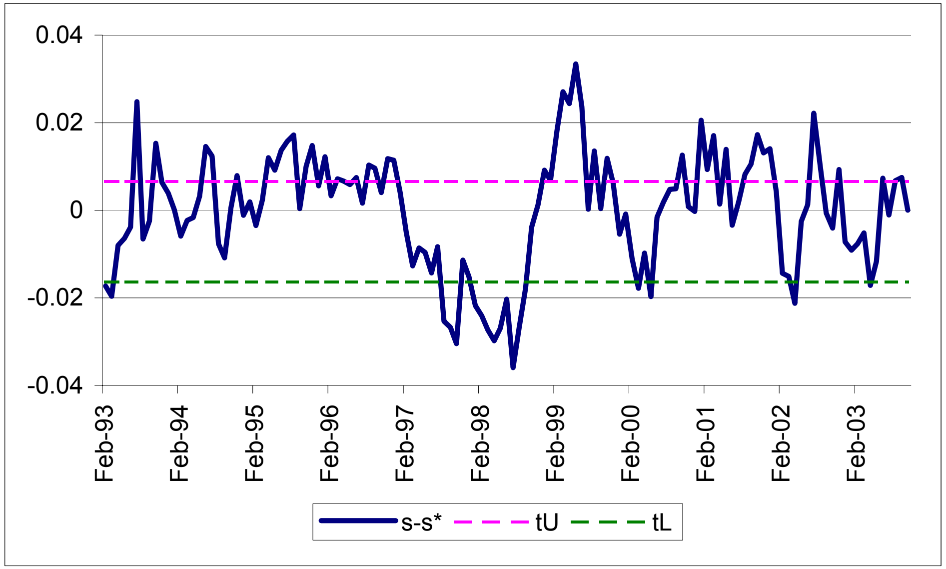


Figure 1: Nominal exchange rate misalignment against inner regime thresholds

Slovenia