BALASSA-SAMUELSON EFFECTS, DEMAND SHIFTS
AND REAL EXCHANGE RATE DETERMINATION:
A non-linear analysis for Central European countries

by

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Abstract
This paper models real exchange rate behaviour in the five Central European Economies that recently joined the EU. Our long-run analysis allows for separate modelling of the Balassa-Samuelson (productivity shock) effect and that of real demand shocks on equilibrium real exchange rates. Short-run exchange rate behaviour is modelled using a general model of non-linear adjustment. We find that equilibrium real exchange rates in these countries are pre-dominantly driven by positive productivity shocks vis-à-vis the EMU average, with demand shocks also playing a significant, albeit more limited, role. We also obtain strong evidence of non-linearity, with the speed of real exchange rate adjustment being in all cases a function of the size, and in two cases, the sign of the latter's deviation from its equilibrium level.

Keywords: real exchange rates, Balassa-Samuelson effect, demand shocks, non-linear adjustment

JEL Classification: C52, F31
1. INTRODUCTION

Explaining the real exchange rate in the transition economies of Central and Eastern Europe over the past 15 years presents a particular challenge to economic modellers. Throughout the region, the liberalisation of the early 1990s was marked by extensive structural change, accompanied by rampant inflation and the rapid depreciation of the nominal exchange rate. Although inflation has to some extent been stabilised since the mid 1990s, the process of structural change has continued to transform the economies of these formerly centrally planned countries.

The real exchange rate has reflected the turbulence of these events. Not surprisingly, in most of transition economies the real exchange rate is non-stationary, having appreciated over this period. From the theoretical perspective, non-stationarity has been explained by Balassa (1964) – Samuelson (1964) effects (see e.g. Taylor 1995, Rogoff, 1996), arising from the sustained increase in productivity that has been observed over this period as these economies increasingly orientate themselves towards the market (see Halpern and Wyplosz, 1997, Grafe and Wyplosz, 1999, Krajnyak and Zettelmeyer 1998). Since the equilibrium real exchange rate will adjust in response to these structural changes, it is essential that the latter are included in any empirical model of real exchange rate determination. But this poses a challenge to empirical modellers who must distinguish underlying structural changes from movements in the data that reflect the effect of random, country-specific demand shocks (see e.g. Desai, 1998 and Dibooglu and Kutan (2001)), and the generally observed extreme turbulence observed in transition economies during this period. To date, this has been done in a fairly ad-hoc way, relying on simple trends or using the real interest rate to construct measures of demand shocks, the effects of which are then removed (see e.g. Taylor and Sarno, 2001).
The task of extracting the underlying equilibrium real exchange rate is made even more difficult by the seemingly complex, non-linear behaviour of the real exchange rate in the short-term. Such non-linearities have been 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 disequilibrium is relatively mild, are punctuated by periods of much faster adjustment in response to more marked deviations of the real 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, mainly focusing on the G7 area, including Obstfeld and Taylor (1997), Michael et al. (1997), Taylor and Peel (2000), Taylor, Peel and Sarno (2001) and Baum et al (2001) have found evidence of such non-linear behaviour in the actual movements of real exchange rates.

It is clear that empirical models aiming to explain real exchange rates must account for both factors discussed above, i.e. be able to capture changes in the underlying equilibrium real exchange by extracting the underlying trend from the data, while also modelling the complex, nonlinear short-run movements in the real exchange rate. Inadequate modelling of either of these processes will undermine any econometric examination of real exchange rates. This paper attempts to provide such an empirical investigation for real exchange rates in Central Europe. We focus on those European countries which recently joined the EU, namely the Czech Republic, Hungary, Poland, Slovakia and Slovenia,¹ and model their real exchange rates against the Euro (the ECU prior to 1998), which they all aspire to join within the foreseeable future.

¹ The Baltic states (Estonia, Latvia and Lithuania) are excluded from the analysis due to the lack of sufficiently long data series.
future. Depending on data availability, our sample covers the period from various points in the early 1990s to 2003.

Our methodology develops the existing literature in two ways. First, our model of the equilibrium real exchange rate is based on the view that the equilibrium real exchange rate responds differently to demand shocks and supply shocks. We therefore use a statistical filter to separate the underlying longer-run movements in the data from shorter-term effects, assuming that the long-run components of the real exchange rate represent supply shocks and that the short-term components measure demand shocks. We then model the equilibrium real exchange rate as the fitted values from a regression of the real exchange rate on these components. We expect a stronger response to supply shocks since these capture the effects of structural change. Second, we estimate a nonlinear model of the exchange rate in the short run that allows for non-linear and asymmetric adjustment.

We obtain several interesting results. First, we find clear evidence in favour of Balassa-Samuelson effects since supply shocks are the main influence on real exchange rates over the longer run. Demand shocks play a significant but less pronounced role in the majority of the countries examined. Second, we obtain strong evidence of non-linear adjustment in the short-run. We find that the real exchange rate adjusts more rapidly in an outer regime where exchange rates are away from equilibrium, compared to much slower adjustment in an inner regime where disequilibrium is less marked (we find that in the inner regime the real exchange rate is a random walk in the Hungary and the Czech Republic). We also find that exchange rate adjustment is more complex in Hungary and the Czech Republic, where thresholds of the inner regime may be asymmetric. This may be related to asymmetric policy objectives in the loss functions of these countries’ authorities.
The remainder of the paper is structured as follows: Section 2 models the equilibrium level of the real exchange rate using variables capturing the separate effect of relative demand and supply shocks. Section 3 uses the results obtained in section 2 to model short-run real 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 real exchange rate; and section 3.3 estimates non-linear error-correction models of real exchange rate adjustment. Finally, section 4 summarises and offers some concluding remarks.

2. MODELLING EQUILIBRIUM REAL EXCHANGE RATES

We use monthly data for real exchange rates against the ECU (the EURO since 1999), calculated using spot exchange rates and producer price indexes taken by the IMF databank provided by Datastream. Data availability allows for the following sample periods: 1993(1)-2004(2) for the Czech Republic; 1990(1)-2003(12) for Hungary; 1990(1)-2004(1) for Poland; 1993(1)-2004(1) for Slovakia and 1992(3)-2003(12) for Slovenia. The calculated real exchange rates (in logs) are depicted in Figure 1. With the exception of Slovenia, they present a strong appreciation pattern during the period under examination. Preliminary analysis of their order of integration using the augmented Dickey-Fuller (ADF) tests (Dickey and Fuller, 1979) confirm that the series are integrated of order 1, with the exception of Slovenia whose real exchange rate appears to be stationary.

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2 The real exchange rate is defined as the product of the nominal exchange rate by the ratio of foreign (European) to domestic price levels. An increase (reduction) in the value of the real exchange rate denotes a real depreciation (appreciation) of the domestic currency against the Euro.

3 These tests are not reported but are available by the authors upon request.

4 Neither these tests, nor subsequent estimates of cointegrating relationships are affected by non-linearities (Michael, et al, 1997).
A non-stationary real exchange rate is theoretically explained by shocks in the real sector of the economy. Positive relative supply shocks (e.g. an increase in domestic productivity) give rise to Balassa-Samuelson effects leading to real currency appreciation. By contrast, an increase in relative demand for domestic products is associated with real currency depreciation. Taylor and Sarno (2001) modelled these two separate effects in transition economies using a time trend as a proxy of relative productivity gains and the real interest rate differential as a proxy of relative demand shocks. Our modelling approach uses a statistical procedure enabling us to extract and investigate the separate effects of each of these two kinds of shocks in a more direct way. More specifically, we fit a Hodrick-Prescott (1997) filter in the (log of) index of industrial production, which we use as a proxy for output. For every period in time, we define the value of the fitted series ($\bar{y}$) to be the equilibrium output level. The difference between the observed and the fitted series is defined to be the residual (excess) demand output component ($y^R$). We define the difference between the equilibrium levels of domestic industrial output and that of the EMU to be a proxy of equilibrium domestic supply relative to that of the EMU average ($\bar{y} - \bar{y}_{EMU}$). Changes in ($\bar{y} - \bar{y}_{EMU}$) are then defined to be a measure of relative supply shocks. In a similar fashion, the difference between the domestic and EMU excess demand components ($y^R - y^R_{EMU}$) is defined to be a proxy of relative demand shocks. We then model real exchange rates on the calculated relative supply and relative demand as suggested by equation (1) below, where $\alpha$ is a constant term and $\varepsilon$ denotes a white noise random error term.

$$q_t = \alpha + \beta_1 (\bar{y} - \bar{y}_{EMU})_t + \beta_2 (y^R - y^R_{EMU})_t + \varepsilon_t \quad (1)$$
The results of estimating (1) are reported in Table 1. With the exception of Slovenia, we obtain strong evidence in favour of the Balassa-Samuelson hypothesis, with long-term real exchange rates predominantly determined by positive supply (productivity) shocks vis-à-vis the EMU average. Demand shocks are also found to play a significant, albeit more limited, role, in the majority of the countries examined (Hungary, Poland and Slovakia). In the case of Slovenia, both terms are insignificant, a finding consistent with the Purchasing Power Parity hypothesis, according to which the real exchange rate is represented, in the long-run, by a constant term. The reported ADF tests suggest that all models are cointegrated at the 5 per cent level, with the exception of the Czech Republic. However, failing to accept cointegration is not necessarily a reflection of model misspecification: it can also reflect the effects of non-linear adjustment to the equilibrium level defined by (1), a hypothesis which is we test formally in the following section.

3. MODELLING SHORT-RUN REAL EXCHANGE RATE ADJUSTMENT

3.1. Linear short-run adjustment models

In this section we model the short-run adjustment of the real exchange rate towards its equilibrium level described by (1). We start by estimating a standard error-correction equation given by (2) below:

\[ \Delta q_t = \beta (L) \Delta q_{t-1} + \gamma (L) \Delta \hat{q}^*_t + \delta (q - \hat{q}^*)_{t-1} + \epsilon_t \] (2)

In (2), \( q \) is the log of the actual (observed) real exchange rate, \( \hat{q}^* \) is the log of the derived equilibrium real exchange rate (the fitted values of equation (1)), \( \beta (L) \) and \( \gamma (L) \) are polynomials in the lag operator, \( L \), \( \epsilon \) is a white noise error term and \( \Delta \) is
the first difference operator. The mechanism through which the actual real exchange rate converges to its equilibrium value is the error correction term \((q-q^*)_{t-1}\), which measures the deviation of the real exchange rate from its equilibrium value. If this is statistically significant, there exists a long-run (cointegrating) relationship between real exchange rates and the calculated supply and demand shock terms, with \(\delta\) measuring the speed of adjustment to the equilibrium level defined by (1).

Table 2 presents estimates of the linear error correction equations in (2). We report estimates of parsimonious models obtained using a general-to-specific specification search on a baseline model using twelve lags of all variables. In all cases, including the equation referring to the Czech Republic, the error correction term is statistically significant at the 5 per cent level. This is consistent with our cointegration findings in Table 1. However, the estimated speed of adjustment present a rather mixed picture: in two countries (Slovakia and Slovenia), reversion to equilibrium is rather fast; in Poland it is rather moderate; and in the cases of the Czech Republic and Hungary it is slow. The equations are generally well-specified, but their explanatory power is rather low. This may reflect the well-established stylised fact of high short-run volatility in the movements of real exchange rates. However, it may also be due, at least to some extent, to the existence of non-linearities in the process of real exchange rate determination. We proceed to test this hypothesis formally in the following section.

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5 An alternative modelling approach would be to substitute (1) into (2) and estimate the resulting equation \(\Delta at = \beta (L) \Delta at_{-1} + \gamma (L) \Delta (\pi' z)_{t-1} + \delta (a - \pi' z)_{t-1} + \epsilon_{t}\), with \(z'\) defined as [(\(y - EMU_y\)) \((yR - EMU_{yR})\)]. 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.

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

The hypothesis of linear adjustment of the real exchange rate towards its equilibrium level can be tested using the procedure described in Saikonnen and Luukkonen (1988), Luukkonen et al (1988), Granger and Teräsvirta (1993) and Teräsvirta (1994). This involves the estimation of equation (3) below:

\[
(q - \hat{q}^*), = \gamma_{00} + \sum_{j=1}^{\phi} \{\gamma_{0j} (q - \hat{q}^*)_{t-j} + \gamma_{1j} (q - \hat{q}^*)_{t-j} (q - \hat{q}^*)_{t-d} + \gamma_{2j} (q - \hat{q}^*)_{t-j} (q - \hat{q}^*)_{t-d}^2 + \\
\gamma_{3j} (q - \hat{q}^*)_{t-j} (q - \hat{q}^*)_{t-d}^3 + \gamma_{4} (q - \hat{q}^*)_{t-d}^2 + \gamma_{5} (q - \hat{q}^*)_{t-d}^3 + \nu(t)\}
\]

\[
(3)
\]

In (3), \((q - \hat{q}^*)_t\) is the estimated deviation from equilibrium measured by the estimated residual term obtained from (1), \(d\) is the delay parameter of the transition function to be used and \(\nu(t) \sim niid (0, \sigma^2)\). 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 \(\phi\) through inspection of the partial autocorrelation function,\(^7\) (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. Therefore, we conclude that the hypothesis of non-linear adjustment of the real exchange rate to its equilibrium level cannot be rejected.

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\(^7\) Granger and Teräsvirta (1993) and Teräsvirta (1994) advise against choosing \(\phi\) using an information criteria such as the Akaike since this may induce a downward bias.
3.3. Non-linear short-run adjustment models

We now model the previously detected non-linearity in the adjustment of the real exchange rate using the Quadratic Logistic Smooth Transition Error Correction Model (QL-STECM). This model, which is a generalisation of the ESTAR model, is specified as follows:

\[
\Delta q_t = \theta_t M_{It} + (1 - \theta_t) M_{Ot} + \epsilon_t \tag{4}
\]

\[
M_{It} = \beta_{I1} (L) \Delta q_{t-1} + \gamma_{I1} (L) \Delta \hat{q}^* + \delta_I (q - \hat{q})_{t-1} + \epsilon_{It} \tag{5}
\]

\[
M_{Ot} = \beta_{O1} (L) \Delta q_{t-1} + \gamma_{O1} (L) \Delta \hat{q}^* + \delta_O (q - \hat{q})_{t-1} + \epsilon_{Ot} \tag{6}
\]

\[
\theta_t = \text{pr} \{ \tau^L \leq (q - \hat{q}^*)_{t-1} \leq \tau^U \} = \frac{1}{1 + e^{-\sigma(q - \hat{q}^*)_{t-1} - \tau^L \sqrt{\tau^U - \tau^L}}} \tag{7}
\]

Equation (4) 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 (5) and (6) describe \( M_I \) and \( M_O \) as linear error-correction models, similar to (2). Equation (7) specifies the regime weight \( \theta_t \) as the probability that the transition variable \( (q - \hat{q}^*)_{t-1} \) lies within the “regime boundaries” \( \tau^L \) and \( \tau^U \), where the probability is described using a quadratic logistic function. We expect \( \tau^L < 0 \) and \( \tau^U > 0 \). Real exchange rates are mainly determined by \( M_I \) (the inner regime) when the real exchange rate is close to its equilibrium value described by (1) and mainly by \( M_O \) (outer regime) in periods of significant exchange rate misalignment, with \( \sigma \) 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 real exchange rate only adjusts towards its

\[\text{For a detailed discussion of this model see van Dijk et al. (2002)}\]
fundamental value in the outer regime, evolving as a random walk in the inner regime. In the case where $\tau^U + \tau^L = 0$, the model is in effect equivalent to the ESTAR model since the speed of adjustment depends only on the size of the deviation of exchange rates from fundamentals. If $\tau^U + \tau^L \neq 0$, 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 real 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. 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 regression standard errors smaller than those of the corresponding linear models. In line with our previous findings on non-linearity, we obtain significant differences between the speeds of adjustment in the inner and the outer regime. More specifically, in the cases of the Czech Republic and Hungary, the error correction coefficient is insignificant in the inner regime, implying no convergence towards equilibrium (random walk behaviour), but highly significant in the outer regime. For the remaining countries, the error-correction term is statistically significant in both regimes, but reversion to equilibrium is faster in the outer rather than the inner. These results are consistent with those obtained by Taylor and Sarno (2001), who modeled non-linearities in real exchange rates using the ESTAR model.

9 These models include the statistically significant crisis dummies included in their linear counterparts mentioned in the note accompanying the Table. 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.

10 However, in all cases $\sigma$ 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).
Finally, in two out of five cases (the Czech Republic and Hungary), the regime thresholds appear to be asymmetric. In the case of Hungary, the value of the upper regime threshold is higher in absolute value rather than the lower while the opposite is true for the Czech Republic. This asymmetry, which cannot be captured by the TAR or ESTAR models extensively used in the literature, may be related to asymmetric policy objectives in the loss functions of these countries’ authorities.\textsuperscript{11}

Figure 2 presents the estimated disequilibrium term (solid line) vis-à-vis the estimated regime thresholds (dotted lines). Three interesting observations emerge: First, in three out of five countries, namely the Czech Republic, Hungary and Poland, the introduction of the transition process was accompanied by significant undervaluation of the real exchange rate relative to its equilibrium value. This is not surprising, given the very pronounced nominal currency devaluations experienced by these countries in the early 1990s. Second, deviations of the real exchange rates from equilibrium have predominantly been taking values within the inner regime, although values in the outer regime are not uncommon, particularly in the case of the Czech Republic. Finally, at the end of 2003 real exchange rates in the countries of Central Europe seem to have been close to their equilibrium value. The only exception is Poland, whose real exchange rate seems to have been undervalued by 7.5 percentage units, placing the disequilibrium in the outer regime. The reported close proximity to equilibrium is an indication that in the absence of unexpected events or major country-specific policy shifts, joining the single currency in the foreseeable future is a realistic proposition for the countries examined.

\textsuperscript{11} Consider, for example, real 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, as an overvalued exchange rate may lead to short-run output losses and an increase in unemployment.
4. SUMMARY AND CONCLUDING REMARKS

This paper has examined real exchange rate determination in the five Central European countries which have recently joined the EU, namely the Czech Republic, Hungary, Poland, Slovakia and Slovenia. Modelling real exchange rates in these countries presents a challenging task for economists as the latter have been subject to a variety of effects including the rise in productivity observed during the transition process; random demand shocks; and non-linearities in the process of short-run towards equilibrium.

Our econometric approach enables us to extract measures of relative demand and supply shocks vis-à-vis the EMU average and then model the equilibrium real exchange rate as the fitted values from a regression of the real exchange rate on these components. We have also estimated a nonlinear error-correction model allowing for non-linear and asymmetric adjustment of the real exchange rate towards the extracted equilibrium level. We found clear evidence of Balassa-Samuelson effects, with supply shocks being the main determinant of real exchange rates during the transition period. Relative demand shocks are also found to play a significant, but less pronounced, role in real exchange rate determination. Finally, we found strong evidence of non-linear real exchange rate adjustment, with exchange rate correcting more rapidly large rather than small deviations from equilibrium. We also found that exchange rate adjustment may be even more complex in Hungary and the Czech Republic, where the thresholds of the inner regime may be asymmetric.

Our work can be extended in several ways. In the empirical level, one direction on which the authors are currently working, is to investigate the effects of supply and demand shocks on the movements of the nominal rather than the real exchange rate. Such an analysis would have important policy implications for the set
of countries we have examine, given the importance of an accurate estimate of their
equilibrium nominal exchange rate against the Euro before the latter’s irrevocable
fixing which will accompany their accession to the EMU. In the theoretical level,
another extension also currently pursued by the authors, is to develop a formal model
of non-linear exchange rate behaviour, perhaps drawing on the recent literature on
non-linear policy rules, which would provide a clearer theoretical grounding for our
empirical work.

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Figure 1: Real exchange rates in Central Europe

Czech Republic

Note: Real exchange rates are defined as the product of spot exchange rates times the ratio of foreign to domestic price levels.
Figure 1 (continued): Real exchange rates in Central Europe

Slovenia

Note: Real exchange rates are defined as the product of spot exchange rates times the ratio of foreign to domestic price levels.
Table 1

Long-run models

<table>
<thead>
<tr>
<th></th>
<th>Czech Republic</th>
<th>Hungary</th>
<th>Poland</th>
<th>Slovakia</th>
<th>Slovenia</th>
</tr>
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<tbody>
<tr>
<td>constant</td>
<td>1.507</td>
<td>-1.708</td>
<td>0.406</td>
<td>1.573</td>
<td>2.208</td>
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<tr>
<td></td>
<td>(0.004)</td>
<td>(0.004)</td>
<td>(0.006)</td>
<td>(0.002)</td>
<td>(0.002)</td>
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<tr>
<td>( \bar{y} - \bar{y}_{EMU} )</td>
<td>-0.873</td>
<td>-0.684</td>
<td>-0.650</td>
<td>-1.123</td>
<td>0.135</td>
</tr>
<tr>
<td></td>
<td>(0.136)</td>
<td>(0.032)</td>
<td>(0.041)</td>
<td>(0.041)</td>
<td>(0.149)</td>
</tr>
<tr>
<td>( y^r - y^r_{EMU} )</td>
<td>0.025</td>
<td>0.157</td>
<td>0.271</td>
<td>0.010</td>
<td>0.031</td>
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<tr>
<td></td>
<td>(0.097)</td>
<td>(0.083)</td>
<td>(0.085)</td>
<td>(0.046)</td>
<td>(0.057)</td>
</tr>
<tr>
<td>Cointegration ADF</td>
<td>-2.09</td>
<td>-4.38**</td>
<td>-3.55**</td>
<td>-4.00**</td>
<td>-3.81**</td>
</tr>
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Notes: Standard errors in parentheses; Critical values for cointegration ADF test: 95 per cent: -2.88; 99 per cent: -3.48.
Table 2

Linear Error Correction Models

<table>
<thead>
<tr>
<th></th>
<th>Czech Republic</th>
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<th>Poland</th>
<th>Slovakia</th>
<th>Slovenia</th>
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<tr>
<td>constant</td>
<td>-0.001 (0.001)</td>
<td>-0.002 (0.001)</td>
<td>-0.001 (0.001)</td>
<td>0.000 (0.001)</td>
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<tr>
<td>(\Delta q_{t-1})</td>
<td>-0.236 (0.077)</td>
<td></td>
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<tr>
<td>(\Delta q_{t-6})</td>
<td>0.163 (0.073)</td>
<td>0.199 (0.068)</td>
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<tr>
<td>(\Delta q_{t-7})</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td>0.261 (0.072)</td>
</tr>
<tr>
<td>(\Delta q_{t-8})</td>
<td></td>
<td></td>
<td>0.184 (0.071)</td>
<td>0.264 (0.074)</td>
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<tr>
<td>(\Delta q^*_1)</td>
<td>1.422 (0.586)</td>
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<tr>
<td>(\Delta q^*_{t-6})</td>
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<td>-0.186 (0.089)</td>
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<tr>
<td>(\Delta q^*_{t-9})</td>
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<td></td>
<td>-0.244 (0.086)</td>
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<tr>
<td>(\Delta q^*_{t-12})</td>
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<td></td>
<td></td>
<td>-0.470 (0.172)</td>
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<tr>
<td>(\Delta (q-q^*)_{t-1})</td>
<td>-0.073 (0.030)</td>
<td>-0.038 (0.019)</td>
<td>-0.115 (0.035)</td>
<td>-0.233 (0.053)</td>
<td>-0.227 (0.061)</td>
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<tr>
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<td>0.27</td>
<td>0.36</td>
<td>0.19</td>
<td>0.32</td>
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<td>Regression Std Error</td>
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<td>0.00975</td>
<td>0.01187</td>
<td>0.00959</td>
<td>0.00889</td>
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<tr>
<td>F-AR</td>
<td>0.56 [0.79]</td>
<td>0.81 [0.58]</td>
<td>0.42 [0.89]</td>
<td>0.33 [0.94]</td>
<td>1.49 [0.18]</td>
</tr>
<tr>
<td>F-ARCH</td>
<td>0.28 [0.96]</td>
<td>0.49 [0.84]</td>
<td>1.35 [0.23]</td>
<td>4.99 [0.00]</td>
<td>0.51 [0.83]</td>
</tr>
<tr>
<td>Chi sq. Normality</td>
<td>7.47 [0.02]</td>
<td>0.51 [0.77]</td>
<td>0.33 [0.85]</td>
<td>2.96 [0.23]</td>
<td>2.64 [0.27]</td>
</tr>
<tr>
<td>F-Het</td>
<td>0.99 [0.46]</td>
<td>0.42 [0.88]</td>
<td>0.65 [0.87]</td>
<td>0.57 [0.64]</td>
<td>0.60 [0.78]</td>
</tr>
<tr>
<td>RESET</td>
<td>4.81 [0.03]</td>
<td>2.92 [0.09]</td>
<td>0.51 [0.48]</td>
<td>0.85 [0.36]</td>
<td>1.34 [0.25]</td>
</tr>
</tbody>
</table>

Table 3

**Linearity tests**

<table>
<thead>
<tr>
<th>Country</th>
<th>φ</th>
<th>d</th>
<th>F-test [p-value]</th>
</tr>
</thead>
<tbody>
<tr>
<td>Czech Republic</td>
<td>2</td>
<td>1</td>
<td>4.54 [0.00]</td>
</tr>
<tr>
<td>Hungary</td>
<td>1</td>
<td>4</td>
<td>3.62 [0.00]</td>
</tr>
<tr>
<td>Poland</td>
<td>9</td>
<td>4</td>
<td>1.82 [0.01]</td>
</tr>
<tr>
<td>Slovakia</td>
<td>1</td>
<td>4</td>
<td>2.88 [0.02]</td>
</tr>
<tr>
<td>Slovenia</td>
<td>9</td>
<td>12</td>
<td>1.65 [0.04]</td>
</tr>
</tbody>
</table>

**NOTES:** φ is the order of the autoregressive component and d the order of the delay parameter in the artificial regression.

\[
(q - \hat{q}^*)_t = \gamma_{00} + \sum_{j=1}^{\phi} \gamma_{0j} (q - \hat{q}^*)_{t-j} + \gamma_{1j} (q - \hat{q}^*)_{t-j} (q - \hat{q}^*)_{t-d} + \gamma_{2j} (q - \hat{q}^*)_{t-j} (q - \hat{q}^*)_{t-d}^2 + \gamma_{3j} (q - \hat{q}^*)_{t-j} (q - \hat{q}^*)_{t-d}^3 + \gamma_{4j} (q - \hat{q}^*)_{t-d} + \gamma_{5j} (q - \hat{q}^*)_{t-d}^2 + v(t)
\]

The reported LM statistics are the estimated scores and the associated p-values are obtained from applying an LM F-test on equation (5) where the null is described by:

\[
H_0 = \{\gamma_{2j} = \gamma_{3j} = \gamma_{4j} = \gamma_{5j} = 0\}
\]
<table>
<thead>
<tr>
<th></th>
<th>Czech Republic</th>
<th>Hungary</th>
<th>Poland</th>
<th>Slovakia</th>
<th>Slovenia</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>M₁ (Inner regime)</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Constant</td>
<td>-0.003 (0.002)</td>
<td>-0.001 (0.133)</td>
<td>-0.001 (0.001)</td>
<td>-0.002 (0.001)</td>
<td>0.001 (0.001)</td>
</tr>
<tr>
<td>Δq₁</td>
<td>0.298 (0.101)</td>
<td>0.172 (0.086)</td>
<td>0.184 (0.089)</td>
<td>2.755 (1.063)</td>
<td></td>
</tr>
<tr>
<td>Δq₆</td>
<td>2.556 (1.042)</td>
<td>0.405 (0.133)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Δq₈</td>
<td>-2.134 (0.984)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Δq₉₆</td>
<td>0.280 (0.123)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(q-q*)₉₆</td>
<td>0.078 (0.129)</td>
<td>-0.066 (0.046)</td>
<td>-0.088 (0.044)</td>
<td>-0.193 (0.075)</td>
<td>-0.240 (0.085)</td>
</tr>
<tr>
<td><strong>M₂ (Outer regime)</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Constant</td>
<td>0.001 (0.001)</td>
<td>-0.005 (0.001)</td>
<td>-0.005 (0.002)</td>
<td>-0.003 (0.002)</td>
<td>-0.001 (0.001)</td>
</tr>
<tr>
<td>Δq₁</td>
<td>-0.379 (0.105)</td>
<td>0.325 (0.117)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Δq₂</td>
<td>0.298 (0.145)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Δq₇</td>
<td>0.260 (0.126)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Δq₉₇</td>
<td>0.255 (0.108)</td>
<td>-7.207 (2.236)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Δq₉₉₇</td>
<td>1.495 (0.700)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Δq₉₉₉₇</td>
<td>2.774 (1.347)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Δq₉₉₉₉₇</td>
<td>2.577 (1.140)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Δq₉₉₉₉₉₇</td>
<td>-0.346 (0.136)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(q-q*)₉₉₉₉₉₇</td>
<td>-0.081 (0.032)</td>
<td>-0.136 (0.043)</td>
<td>-0.163 (0.052)</td>
<td>-0.317 (0.089)</td>
<td>-0.311 (0.090)</td>
</tr>
<tr>
<td>σ</td>
<td>46.72 (106.8)</td>
<td>50.00 (167.3)</td>
<td>146.2 (526.8)</td>
<td>50.00 (198.8)</td>
<td>26.02 (93.9)</td>
</tr>
<tr>
<td>τₑ</td>
<td>0.020 (0.003)</td>
<td>0.035 (0.003)</td>
<td>0.033 (0.001)</td>
<td>0.014 (0.002)</td>
<td>0.014 (0.001)</td>
</tr>
<tr>
<td>τₑ</td>
<td>-0.030 (0.003)</td>
<td>-0.024 (0.004)</td>
<td>-0.035 (0.001)</td>
<td>-0.016 (0.001)</td>
<td>-0.012 (0.005)</td>
</tr>
<tr>
<td>R²</td>
<td>0.49</td>
<td>0.34</td>
<td>0.38</td>
<td>0.21</td>
<td>0.39</td>
</tr>
<tr>
<td>Regression Std Error</td>
<td>0.01033</td>
<td>0.00951</td>
<td>0.01161</td>
<td>0.00953</td>
<td>0.00852</td>
</tr>
<tr>
<td>F-AR</td>
<td>1.02 [0.42]</td>
<td>0.86 [0.54]</td>
<td>0.64 [0.72]</td>
<td>0.57 [0.77]</td>
<td>0.76 [0.62]</td>
</tr>
<tr>
<td>F-ARCH</td>
<td>0.40 [0.90]</td>
<td>0.75 [0.63]</td>
<td>0.95 [0.47]</td>
<td>2.36 [0.03]</td>
<td>0.37 [0.92]</td>
</tr>
<tr>
<td>Chi sq. Normality</td>
<td>1.22 [0.54]</td>
<td>0.31 [0.85]</td>
<td>0.22 [0.90]</td>
<td>1.96 [0.38]</td>
<td>5.58 [0.06]</td>
</tr>
<tr>
<td>F-Het</td>
<td>1.43 [0.11]</td>
<td>0.96 [0.51]</td>
<td>1.32 [0.17]</td>
<td>0.49 [0.95]</td>
<td>0.74 [0.80]</td>
</tr>
<tr>
<td>F-Test H₀: τₑ + τₑ = 0</td>
<td>3.82</td>
<td>5.05</td>
<td>1.10</td>
<td>2.85</td>
<td>2.51</td>
</tr>
</tbody>
</table>

**NOTES:** A. All models have been estimated including dummy variables for periods of particular exchange rate turbulence. These are defined as follows: For Czech Republic 1997-May, 1999-Jan, 2002-June, 2003-
Figure 2: Real exchange rate misalignment versus regime thresholds

Czech Republic

Poland

Hungary

Slovakia
Figure 2 (continued): Real exchange rate misalignment versus regime thresholds

Slovenia

![Graph showing real exchange rate misalignment versus regime thresholds for Slovenia. The graph plots the real exchange rate misalignment against time, with vertical markers indicating regime thresholds.]