Modelling real interest rate differentials:

Evidence from the newly-accession EU countries

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Abstract

We model the behaviour of real interest rate differentials against the EMU average in the ten countries that joined the EU in 2004. We do so by carrying out a number of unit root tests including the block bootstrap, endogenous structural breaks and non-linear unit-root tests on the monthly real interest rate differentials series for the period 1996-2005. Our main finding is the detection of mean reversion in all ten countries examined accompanied, however, by strong elements of idiosyncratic individual behaviour. We also find evidence of structural breaks and/or non-linear mean reversion, typically involving size-, and in some cases sign-, differential adjustment effects. Overall, our findings suggest increasing yet varied integration between the money markets of the newly-accession EU countries and the EMU. Our findings have implications both for the purposes of international portfolio investment as well as for the contact of national monetary policies towards adopting the Euro.

Keywords: Real interest rate parity; newly-accession EU countries; unit root; structural breaks; block-bootstrap; non-linear adjustment

JEL classification: F21, F32, C15, C22

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1. INTRODUCTION

Real interest rate parity (RIRP) states that under market rationality, risk neutrality and full mobility in the capital and goods’ markets, real interest rates are equalised across countries. RIRP assumes Uncovered Interest Parity (UIP) and \textit{ex ante} Purchasing Power Parity (PPP), thereby implying full integration between the national markets of capital, goods, services and factors of production. By covering financial and non-financial markets RIRP is a comprehensive measure of economic integration relevant to international investors and policy makers alike. With regards to the former, RIRP eliminates the basis for international portfolio diversification, while deviations from it leave room for investment strategies involving assets from different national markets. With regards to the latter, the long-run validity of RIRP and persistence of short-run deviations from it determine the degree to which national authorities can influence economic activity through the real interest rate channel in different time horizons. If RIRP holds, this is limited to the influence national authorities can exert on the world interest rate (Mark, 1985, Feldstein 1991); if it does not, authorities can use this channel effectively for the purpose of macroeconomic management.¹

The literature reviewed in section 2 below has mainly concentrated on the validity of RIRP against the USA has so far highlighted the importance of RIRP in the areas of international investment and macroeconomic management, having produced mixed results with regards to the latter’s validity. At least recently, however, it has overlooked the latter’s relevance in the context of the European Economic and Monetary Union (EMU). If RIRP does not hold within the EMU, or the persistence of deviations from it is heterogeneous across member states, then shifts in the single

¹ RIRP is also directly relevant to our understanding of exchange rate movements and the authorities’ ability to manage the latter, as it underpins a number of mainstream monetary models of exchange determination (see Frenkel 1976, Bilson 1978, Frankel 1979, Mussa 1982) and is widely thought to be directly related to the movements of real exchange rates (see Minford and Peel, 2006)
monetary policy can result in asymmetric economic effects among the members of the Eurozone (see Holmes, 2005). In a similar fashion, failure or highly persistent deviations from RIRP between the newly-accession countries (NACs) of the European Union (EU) and the EMU average would indicate that the financial and economic convergence process taking place in the former has not yet been fully completed. In that case, joining the single currency may come at a cost for the NACs, as it may result in a sub-optimal, for their own domestic economic requirements, single monetary policy.

In this paper we test for RIRP against the EMU average in the ten countries that joined the EU in May 2004 and use the results to draw inference relating to the portfolio diversification opportunities these countries offer to international investors, as well as the degree to which they have achieved financial and general economic progress towards joining the euro. Our analysis covers the period 1996-2005, a period during which the NACs had abolished almost all capital controls and overcome the initial phase of transition following the market reforms of the early 1990s. Unlike other studies, we test for RIRP working with real interest rate differentials calculated using data sets of identical definition for nominal interest rates and CPI inflation, which allows us compare real returns of identically measured across countries. We use a battery of econometric tests, including traditional unit root tests as well some more recently developed/highly innovative approaches, such as unit root tests allowing for endogenously determined structural breaks, non-linear real interest rate adjustment and an innovative procedure based on the block bootstrap. We find mean reversion in the real interest rates of all ten countries accompanied, however, by strong elements of idiosyncratic individual behaviour. We also find evidence of structural breaks and/or non-linear mean reversion, typically involving size-,
adjustment effects. Overall, our findings suggest increasing yet varied integration between the money markets of the NACs and the EMU.

The remainder of the paper is structured as follows: Section 2 reviews the empirical literature on real interest rate parity. Section 3 discusses our methodology while section 4 presents our data. Sections 5.1 and 5.2 respectively present the results of the standard unit root tests and those allowing for endogenously determined structural breaks on real interest rate differentials. Section 6 presents the results of our non-linear analysis and section 7 those of our block bootstrap analysis. Finally, section 8 summarises the paper and offers some concluding remarks.

2. RELATED LITERATURE

Early studies testing for RIRP include those by Mishkin (1984a, b), Cumby and Obstfeld (1984) and Mark (1985). These estimated regressions of the domestic real interest rate on the foreign and tested RIRP by respectively imposing zero and unity restrictions on the intercept and slope coefficients. These studies typically rejected but were criticised on the grounds that the real interest rates calculated for the periods covered by their analysis include unit roots, in which case testing restrictions on coefficients’ values is problematic as the estimated standard errors are inconsistent.

The literature subsequently moved towards two directions, namely applying cointegration tests on real interest rate regressions and examining the time series properties of real interest rate differentials (RIRDs).2 The first approach was adopted by

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2 A number of other approaches have also been employed to test the validity of the RIRP. For example, MacDonald and Taylor (1989) and Fraser and Taylor (1990) test and reject the RIRP-consistent hypothesis according to which nominal interest rate differentials predict future inflation differentials. Marston (1995) finds that movements of the RIRDs can be explained using variables included in the current information set, leading to rejection of the RIRP hypothesis. Kugler and Neusser (1993) adopt a stationary multivariate time-series framework using ex-post real interest rates; their findings provide evidence in favour of the RIRP hypothesis for a number of countries. Finally, Dutton (1993) tested for RIRP using real interest rates calculated using prices of traded goods, producing results favourable towards RIRP.
Goodwin and Grennes (1994), Chinn and Frankel (1995), Moosa and Bhatti (1996) and Awad and Goodwin (1998) who produced evidence of mean reversion for the RIRDs, although not necessarily around the zero mean postulated by RIRP. The second approach was originally applied by Meese and Rogoff (1988) and Edison and Pauls (1993) who applied Augmented Dickey-Fuller (ADF) tests (Dickey and Fuller, 1979) on the RIRDs of the major industrialised countries against the US Dollar. These rejected RIRP, as they found the RIRDs series to include a unit root. However, doubt on these negative findings can be cast on a number of grounds. First, there exists the well-known low-power problem of the ADF test; second, the standard ADF test does not account for structural breaks in the series examined, which increases the bias in favour of the null of unit-root (see Perron, 1989); finally, the ADF test does not capture potential non-linearities in the series’ adjustment process, thus increasing the risk of bias further.

A number of studies have sought to address these pitfalls. For example, Obstfeld and Taylor (2002) seek to increase the power of the ADF test by extending the sample period and using a generalised least square version of the ADF test. Their findings reject the unit root hypothesis, which they interpret as evidence in favour of RIRP. Wu and Chen (1998) also employ panel-based unit root tests providing evidence of mean-reverting behaviour for the RIRDs of a number of countries against the US Dollar and the German mark. Panel-data techniques are also used by Gagnon and Unferth (1995) and Ong et al (1999), both of which reach findings suggesting the existence of convergence to a world real interest rate when the latter are constructed using price indexes for traded goods.

The issue of structural breaks on RIRDs has not been explicitly addressed, although Fountas and Wu (2000), working within the cointegrating approach discussed

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above allow for structural breaks in the cointegrating relationship between domestic and foreign real interest rates.\textsuperscript{4} On the other hand, non-linear adjustment of RIRDs toward a RIRP-consistent long-run equilibrium is examined in three recent studies by Obstfeld and Taylor (2002), Mancuso et al (2003) and Ferreira and Leon-Ledesma (2006).\textsuperscript{5} Such non-linearities may take two forms. First size-adjustment effects where the speed of adjustment to RIRP is a function of the size of the deviations from the latter. Such effects are theoretically analysed by Dumas (1992) who justifies them on the grounds of rigidities in the process of restoring market equilibrium such as transaction costs in trading aiming to exploit arbitrage opportunities; contractual commitments imposing holding assets for minimum time periods and implying costs for early funds’ release; and trading rules requiring that rate differences exceed certain thresholds before arbitrage trading is initiated. Market imperfections of this kind imply that small values of RIRDs are not corrected due to the costs involved in arbitrage trading, while large deviations trigger market forces to restore RIRP.

The second type of non-linearities relates to sign-adjustment effects where the speed of adjustment towards RIRP is a function of the sign (positive versus negative) of deviations from it. Justification for such effects can be provided on the same grounds invoked to justify sign effects in the movements of variables closely associated with real interest rates such as inflation nominal interest rates and exchange rates.\textsuperscript{6} More specifically, a positive RIRD is associated with conditions of monetary tightening and real currency overvaluation, while a negative RIRD is associated with the opposite effects. Assume now a government whose loss function attributes a higher weight to

\textsuperscript{4} A similar idea is explored in the study by Cavaglia (1992) who tests for real interest rate convergence among OECD countries using a model allowing for time-varying model parameters.

\textsuperscript{5} Nakagawa (2002) also finds that a non-linear adjustment process due to transaction costs and sticky prices leads to nonlinear mean reverting behaviour for real exchange rates and real interest rate differentials.

\textsuperscript{6} Among others, see respectively Arghyrou, Martin and Milas 2005, Martin and Milas 2004, and Sarno 2005 and the references therein.
output- rather than inflation-stabilization around a target value. Under price rigidities allowing shifts in nominal interest rates to result in temporary shifts in real interest rates, such a government is likely to tolerate a negative RIRD more than a positive one, as the short-run expansionary output effects of higher liquidity and currency undervaluation are preferable to the contractionary demand effects of a monetary tightening and currency overvaluation. Obstfeld and Taylor (2002), Mancuso et al (2003) and Ferreira and Leon-Ledesma (2006) in combination provide evidence supporting the existence of non-linear effects in RIRD behaviour of both types, but mixed evidence on real interest rate equalisation. Hence, on the balance of the evidence presented, the question of RIRP validity still remains open.

3. METHODOLOGY

Our benchmark testing framework is standard and has been adopted by a number of researchers, including Ferreira and Leon-Ledesma (2006). Assuming full capital mobility and rational, risk-neutral market participants, equilibrium in the financial and goods’ markets is determined by UIP and relative PPP, respectively described by equations (1) and (2) below:

\[ i_t = i^*_t + \Delta s^e_t \]  \hspace{1cm} (1)

\[ \Delta s_t = \pi_t - \pi^*_t \]  \hspace{1cm} (2)

In (1) and (2), \( i \) denotes the nominal interest rate, \( s \) the nominal exchange rate \( \pi \) the rate of inflation; the asterisk symbol a variable referring to the foreign country; \( \Delta \) the first difference operator; and the superscript \( e \) the expectations’ operator. Under rational expectations,
\[ \Delta s_t = \Delta s^e_t + u_t \]  

where \( u_t \) is iid so that \( u_t \sim N(0, \sigma^2) \).

Rearranging (1) as \( \Delta s^e_t = i_t - i^*_t \), using (3) in (1) and combining with (2), we obtain

\[ \pi_t - \pi^*_t = i_t - i^*_t + u_t \]  

(4)

Defining the ex-post real interest rate \( r_t \) as the difference between nominal interest rate and inflation \( (r_t = i_t - \pi_t) \), equation (4) implies

\[ (r - r^*)_t = u_t \]  

(5)

Equation (5) describes RIRP, under which the real interest rate differential is a random error process with zero mean and constant variance. This is a special case of the stochastic model given by equation (6) below:

\[ (r - r^*)_t = \alpha + \rho (r - r^*)_{t-1} + u_t \]  

(6)

Given an initial condition for time 0, defined as \( (r - r^*)_0 \), the solution to this first-order difference equation in is given by

\[ (r - r^*)_t = \frac{\alpha}{1 - \rho} (1 - \rho^t) + \rho^t (r - r^*)_0 + \sum_{i=0}^{t-1} \rho^i u_{t-i} \]  

(7)
If \(|\rho| < 1\), equation (7) implies

\[
\lim_{t \to \infty} (r - r^*)_t = \frac{\alpha}{1 - \rho}
\]

(8)

Given (8), if \(\alpha = 0\), the expected value of \((r - r^*)_t\) equals to zero

\[
E (r - r^*)_t = \frac{\alpha}{1 - \rho} = 0
\]

(9)

Equation (9) is consistent with (5) and describes RIRP. More generally, (6) can be reformulated as an autoregressive model of order \(j\) given by equation (10):

\[
\Delta (r - r^*)_t = \alpha + \phi (r - r^*)_{t-1} + \sum_{i=2}^{j} \beta_i \Delta (r - r^*)_{t-i+1} + u_t
\]

(10)

where

\[
\phi = \sum_{i=1}^{q} \rho_i - 1
\]

(11)

Equation (10) is an Augmented Dickey-Fuller (ADF) regression allowing for four types of real interest rate differential behaviour described by (12) to (15) below:

\[
\phi > 0
\]

(12)

\[
\phi = 0
\]

(13)

\[
\phi < 0 \text{ and } \alpha = 0
\]

(14)
Condition (12) describes an explosive \((r - r^*)_t\) process; (13) is consistent with random walk behaviour, where any shocks to \((r - r^*)_t\) affects the series on a permanent basis. Both (12) and (13) are inconsistent with RIRP. Condition (14) implies a stationary path for \((r - r^*)_t\) with \(\phi\) describing the speed of adjustment to a zero equilibrium real interest rate differential. Like (5) and (9), (13) is consistent with long-run RIPR, with the higher (lower) the absolute value for \(\phi\) the lower (higher) the persistence of a real interest rate differential shock. Finally, condition (15) implies stationarity around a non-zero mean. This is consistent with a positive risk premium in one of the two real interest rates, hence despite the existence of stationarity, (15) is inconsistent with RIRP.

A direct way to test for RIRP is to carry out ADF unit root tests on the \((r - r^*)_t\) series to determine which of the conditions above holds. However, and as we discussed in section 2 above, this approach has a number of disadvantages. First, the ADF test is a low-power one, i.e. it is biased in favour of the null of unit root. Second, there is an issue regarding the selection of the appropriate lag structure of the ADF regression. This is typically determined using information criteria such as the Akaike (AIC). However, Ng and Perron (2001) have argued that the AIC and other information criteria tend to select too small lag structures, potentially resulting in misleading inference. On the other hand, Lopez (1997) has shown that increasing the number of lags in the autoregression of the ADF test results in further power reduction. Overall, rejecting stationarity using the ADF test may be a reflection of either erroneous specification of the ADF’s lag structure or additional reduction in power, rather than a genuine property of the data. Finally, the standard ADF test does not account for structural breaks in the
series examined. Finally, the ADF test does not capture potential structural breaks and non-linearities in the series’ adjustment process.

Taking into account all the above, our testing approach consists of the following steps: First, we test for RIRP using the standard ADF regression in (10), choosing the latter’s lag structure through the AIC. If the standard ADF tests reject the unit root hypothesis, we accept that \((r - r^*)\) is stationary and use the ADF estimates in (10) to determine which of the two conditions, (14) or (15), is validated by the data. Second, if the standard ADF test returns a unit root verdict, we test for unit roots Perron’s (1997) unit-root test allowing for a structural break in the deterministic components of the series. Third, irrespective of the results of the first two tests, we test formally for the existence of non-linearities in the process of RIRD adjustment. For those countries for which the RIRD series is found by the standard ADF test to be stationary for the whole of the sample period, or the unit root hypothesis is maintained both by the standard ADF test as well as by Perron’s unit root test allowing for structural breaks, our non-linearity tests are carried out using the full sample period available. For those countries for which the ADF test maintains the null of unit root but Perron’s test suggests the existence of stationarity with a structural break, our non-linearity tests are carried out for two separate sub-periods defined according to the date of the identified structural break. In those cases for which linearity is rejected, we estimate a non-linear model allowing for size and/or sign effects in the process of real interest rate differential adjustment. Finally, we apply a new and innovative testing procedure based on the block bootstrap. This procedure, discussed in detail in section 7, simultaneously addressing the conflict between maintaining power in the ADF testing procedure while avoiding the biases associated with small lag structures.
4. DATA

Real interest rates in equation (5) can be calculated using nominal interest rates of various maturities and alternative price indexes. Securities of longer maturities involve higher interest-rate and foreign-exchange risk rendering them less attractive to investors seeking arbitrage opportunities in financial markets. To minimize the influence of such factors, we consider returns on short-maturity assets. In particular, we define $i_t$ to be the three-month money market rate. This is available, on an annualised basis, by the International Financial Statistics (IFS) and interest rate data bases provided by Datastream. With regards to price indexes, most existing studies calculate real interest rates using consumer price index (CPI) inflation. This, however, may result in biases due to different definitions of individual-national CPIs. To ensure that such biases do not arise in our analysis we use for all countries the harmonised consumer price index (HCPI) provided by the Eurostat Databank.

HCPI data is available for the post-1996 period. This defines our sample to cover the decade 1996-2005, a period prior to which the vast majority of capital controls had been abolished by the countries examined. We use observations of monthly frequency and, following the literature, work with ex-post real interest rates. The monthly HCPI series exhibit strong seasonality patterns for which we account by seasonally adjusting the data. As we are considering investments of 3-month maturity, we transform the annualised three-month nominal interest rates into a three-month continuously compounded nominal rate of return. Then, following Ferreira and Leon-Ledesma (2006), for every period $t$ we calculate the ex-post real rate of return as the difference between the three-month continuously compounded nominal rate of return in

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7 Working with ex-post rather than ex-ante real interest rates is typically justified on the grounds that the form bypasses the tricky subject of approximating inflation expectations necessary to calculate ex-ante real interest rates.
8 We adjust the series using the Census X11 multiplicative seasonal adjustment method, used by the US Bureau of Census to seasonally adjust publicly released data; the X11 routine is available in EViews.
period $t-3$ minus the quarter-to-quarter inflation rate, i.e. the percentage change of the HCPI between periods $t-3$ and $t$. We then use these calculated real rates of return to construct the $(r - r^*)_t$ series in equation (5).

Figure 1 plots the $(r - r^*)_t$ series for the ten NAC for the period 1996(4)-2005(12). Real interest rate differentials seem to follow different patterns in different countries. In Cyprus, Lithuania and Malta $(r - r^*)$, fluctuates around zero throughout our sample with a fairly even split between positive and negative values. For Hungary and Poland the series typically takes positive values while the opposite is true for Latvia. In the cases of the Czech Republic, Estonia and Slovakia, $(r - r^*)_t$ is on a volatile and increasing path during 1996-1998, followed by a significant reduction in mean and volatility during 1999-2005. Finally, in the case of Slovenia we observe a downward trend between 1996 and mid-2000, followed by a trend-reversal thereafter. Overall, a preliminary examination of the data does not offer a clear picture relating to the validity of RIRP in the NACs and suggests that the RIRD series in some of these countries may have been subject to structural breaks.

5. LINEAR ANALYSIS

5.1. ADF unit root tests

Table 1 presents the results of the benchmark ADF unit root tests described by equation (10). We present ADF tests including a constant, accounting for the possibility of constant risk premium, as well as a constant and trend aiming to capture the possibility of a time-varying risk premium, consistent with the hypothesis that the high growth rates, inflation reduction, and development of financial markets observed in the NAC over the past decade have gradually made them less risky for international investors. To determine which of the two ADF models is most appropriate, we choose
the one with the minimum Akaike score. The model including constant only is preferred for Cyprus, Estonia, Hungary, Lithuania and Poland. The model with constant and trend is preferred for the Czech Republic, Latvia, Malta, Slovakia and Slovenia. Note, however, that in the analysis that follows the two models typically return identical inference.

Using our preferred models, we reject the unit root hypothesis at the 5 per cent level or lower for Cyprus, Hungary, Lithuania and Malta. For the Czech Republic and Estonia the unit root hypothesis is rejected at 10 but not at 5 per cent. Finally, the unit root is maintained at the 5 per cent level for Latvia, Poland, Slovakia and Slovenia, implying that the RIRP does not hold in this group of countries.

For the countries for which the unit root hypothesis is rejected, the $\alpha$ coefficient is statistically equal to zero for Cyprus and Lithuania, implying that RIRP holds. Note that the $\rho$ coefficient suggests a low degree of persistence for Cyprus and a relatively moderate one for Lithuania. In the case of Hungary, our preferred model displays a statistically significant positive $\alpha$ value; this is consistent with a positive risk premium and results in rejection of the RIRP. Moreover, the degree of persistence in Hungary is much higher than in Cyprus and Lithuania. The picture is less clear in the case of Malta where, despite a low degree of persistence in our preferred model (constant and trend), both deterministic components are statistically significant at the 10 per cent level. The same holds true for the constant term of our preferred model for the Czech Republic, where the unit root hypothesis was rejected at the 10 but not 5 percent. By contrast, for Estonia, in line with the RIRP hypothesis, we obtain a non-significant $\alpha$ term. Note, however, that both for the Czech Republic and Estonia, our preferred models suggest a high level of persistence.
To sum-up, the evidence from the benchmark ADF tests is not very favourable towards RIRP: Out of the ten countries examined, the RIRP is maintained at the 5 per cent only in two cases; for six countries RIRP is rejected due to lack of stationarity in the \((r - r^*)_t\) series, for the remaining two, RIRP is rejected due to the presence of deterministic components in the preferred ADF equation.

### 5.2. Unit root tests with structural break

In this section we further examine the cases where the null of non-stationarity was not rejected in the ADF test above by applying the Perron (1997) unit root test allowing for endogenously determined structural breaks. The Perron framework allows testing for the null hypothesis of unit root with no-break against the alternative that real interest rate differentials are stationary around a broken trend. Two alternative models are considered:

\[
(r - r^*)_t = \gamma_0 + \gamma_1 DU_t + \gamma_2 t + \gamma_3 D(T_b)_t + \gamma_4 (r - r^*)_{t-1} + \sum_{i=1}^{k} \delta_i \Delta (r - r^*)_{t-i} + \varepsilon_t
\]  

(16)

\[
(r - r^*)_t = \gamma_0 + \gamma_1 DU_t + \gamma_2 t + \gamma_3 D(T_b)_t + \gamma_4 (r - r^*)_{t-1} + \gamma_5 DT_t \sum_{i=1}^{k} \delta_i \Delta (r - r^*)_{t-i} + \varepsilon_t
\]  

(17)

where \(T_b\) denotes the time at which the change in the trend function occurs, \(DU_t = 1 (t > T_b)\), \(D(T_b)_t = 1 (t = 1 + T_b)\), and \(DT_t = 1 (t > T_b)t\), with \(1(.)\) being the indicator function.

Model (16), the innovative outlier model, allows only for a change in the intercept. It is assumed that this change takes place gradually and depends on the correlation structure of the noise function. Model (17) allows both the intercept and the slope to change at the break date. The unit root test is performed using the \(t\)-statistic for
the null hypothesis of no-break unit root in the regression models (16) and (17) with $t_a(i, T_b, k) (i = 2, 3)$. The break date, $T_b$, and the truncation lag parameter, $k$, are both treated as unknown. $T_b$ is determined endogenously using the following sequential method. Equations (16) and (17) are estimated using the full sample for each possible break date. Then, we select the $T_b$ which minimizes the $t$-statistic for testing $\gamma_4 = 1$:  

$$
t^*_a(i) = \min_{T_a(k+1, T)} t_a(i, T_b, k) (i = 2, 3) \tag{18}
$$

The null hypothesis is rejected if $t^*_a$ exceeds (in absolute value) the corresponding critical value. Perron (1997) shows that the results of Zivot and Andrews (1992) regarding the asymptotic distribution of $t^*_a(2)$ and $t^*_a(3)$ remain valid even without applying arbitrary trimming at each end of the sample. Following Perron (1997), the lag length is chosen endogenously using the ‘t-sig’ approach. That is, we set an upper bound of $k = 12$ and test down until a significant (at the 10% level of significance) lag is found. If all lags are insignificant, then we set $k$ equal to zero.

The Perron unit root test results can be seen in Table 2.1. For the countries for which the null of unit root was maintained in Table 1, the null of no-break unit root is rejected using either model (16) or (17) in the cases of Slovakia and Slovenia. In the case of Estonia, where the null is rejected only with model (17), we employ the residual sum of squares (RSS) and the Akaike information criterion (AIC) to choose between the models. We choose the model that minimizes RSS and AIC. Thus, Estonia is added to the list of countries where the null of non-stationarity is rejected, since Model (17) is chosen.

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9 Two alternative methods to identify break dates were also employed with the results remaining qualitatively quite similar: select $T_b$ so that the absolute value of the $t$-statistic associated with the change in the intercept in model (16), or the slope in model (17) is maximised; select $T_b$ that minimises the $t$-statistic on the parameter associated with the change in the intercept in model (16), or the slope in model (17). The results are not reported here but are available upon request.
Summarizing, once we allow for an endogenously determined structural shift in the series, the evidence of non-stationarity in real interest rate differentials disappears in three out of six transition countries for which the stationarity hypothesis was rejected at the 5 per cent level in section 5.1 (Estonia, Slovakia and Slovenia). Using the AIC and RSS criteria to select between models, the points at which breaks are found to occur are 1999(07) for Estonia (Model 17), 1999(05) for Slovakia (Model 17) and 2003(5) for Slovenia (Model 16).

6. NON-LINEAR ANALYSIS

6.1. Non-linearity tests

In this section we test for non-linear effects in the process of real interest rate differential adjustment using the testing procedure proposed by Saikonnen and Luukkonen (1988), Luukkonen et al (1988), Granger and Teräsvirta (1993) and Teräsvirta (1994). This involves estimating equation (4) below:

\[
(r - r^*)_t = \gamma_{00} + \sum_{j=1}^{\phi} [\gamma_{0j} (r - r^*)_{t-j} + \gamma_{1j} (r - r^*)_{t-j} (r - r^*)_{t-d} + \gamma_{2j} (r - r^*)_{t-j} (r - r^*)^{2}_{t-d} + \\
+ \gamma_{3j} (r - r^*)_{t-j} (r - r^*)^{3}_{t-d} ] + \gamma_4 (r - r^*)^{2}_{t-d} + \gamma_5 (r - r^*)^{3}_{t-d} + \nu_t
\]

In (19), \( \phi \) is the order of the autoregressive parameter \( \gamma \), \( d \) the delay parameter of the transition function; and \( \nu_t \) a random error term. The order of \( \phi \) is determined through the partial autocorrelation function of \( (r - r^*) \).\(^{10} \) Equation (19) is estimated for all

\(^{10}\) Granger and Teräsvirta (1993) and Teräsvirta (1994) advise against choosing \( \phi \) using information criteria, which may induce a downward bias.
plausible values of $d$. Given the monthly frequency of our data, we consider values of $d$ up to 12. For each value of $d$, the null of linearitry described by $H_0: \gamma_{1j} = \gamma_{2j} = \gamma_{3j} = \gamma_{4j} = \gamma_{5j} = 0, j = (1,2...\phi)$, is tested against the alternative of general non-linear adjustment using an LM-type test denoted by $LM^G$. A statistically significant $LM^G$ implies the rejection of the null of linearity with the optimum value of $d$ determined by the highest obtained $LM^G$ score.

If $LM^G$ is significant, another two LM-type tests can be undertaken to determine the exact form of non-linearity, i.e. logistic (consistent with sign adjustment effects) versus quadratic (consistent with size adjustment effects). The first test, denoted by $LM^L$, tests the null of linear or non-linear quadratic adjustment, defined as $H_0: \gamma_{3j} = \gamma_{5j} = 0, j \in (1,2...\phi)$, against the alternative of logistic non-linear adjustment. The second test, denoted by $LM^Q$, tests the null of linearity $H_0: \gamma_{1j} = \gamma_{2j} = \gamma_{4j} = 0 | \gamma_{3j} = \gamma_{5j} = 0, j \in (1,2...\phi)$ against the alternative of quadratic non-linear adjustment. Given a non-insignificant $LM^L$ score, a significant $LM^Q$ score implies quadratic non-linearity. On the other hand, a significant $LM^L$ score followed by a non-significant $LM^Q$ statistic implies logistic non-linearity.

An issue arises, when both the $LM^L$ and $LM^Q$ statistics are significant. In many empirical studies, the testing procedure is terminated when given a significant $LM^G$ score, a statistically significant $LM^L$ score is obtained; in that case, non-linearity of the logistic type is accepted without considering the $LM^Q$ statistic. However, Granger and Terasvirta (1993) and Terasvirta (1994) argue that such a testing sequence may result in misleading conclusions, as the higher order terms of the Taylor expansion used to derive these tests are disregarded (see Terasvirta, 1994 pp. 211-212). In view of this critique, our inference approach has as follows: If the $LM^G$ statistic is statistically insignificant, the testing procedure is terminated and the linearity hypothesis is
maintained. If, on the other hand, the LM$^G$ statistic is statistically significant, we calculate both the LM$L$ and LM$^G$ statistics. In that case, if we obtain a statistically significant LM$L$ score accompanied by an insignificant LM$^Q$ score, we accept the existence of sign effects only. If we obtain a statistically significant LM$^Q$ score accompanied by an insignificant LM$^G$ score, we accept the existence of size effects only. Finally, if both the LM$L$ and LM$^Q$ score are statistically significant, we investigate the existence of both sign and size effects.

Table 3 presents the results of our non-linearity tests. For those countries for which a structural break has been found to occur in the ($r - r^*$) process, we test for non-linearity in each of the two sub-periods defined by the break dates suggested by our preferred model in Table 2. The reported LM$^G$ statistics suggest that the hypothesis of linearity is maintained for Cyprus, Hungary and Malta, as well as for Slovenia during the post-break period. In all other cases, linearity is rejected. The LM$L$ and LM$^Q$ tests suggest that there is no country displaying sign effects only. At the 5 per cent level or lower, size effects only are found for Estonia for both periods (pre- and post-break); Latvia, Lithuania, Poland, Slovakia during the post-break period; and Slovenia during the pre-break period. At the 5 per cent level, sign and size effects are potentially present for the Czech Republic and Slovakia during the pre-break period. Finally, at the 10% level, sign and size effects may also exist in Lithuania as well as Estonia during the pre-break period.

6.2. Models of non-linear adjustment

In this section we model the size and possibly sign effects in the process of RIRD adjustment identified in Table 3. We do so by estimating the Quadratic Logistic
Smooth Threshold Error Correction Model (QL-STECM) given by equations (20) to (24) below (see Van Dijk, 2002):

\[
\Delta(r - r^*)_t = \theta_t M_{1t} + (1 - \theta_t) M_{2t}
\]

(20)

\[
M_{1t} = \alpha + \phi_1 (r - r^*)_{t-1} + \sum_{i=2}^{j} \beta_i \Delta(r - r^*)_{t-i+1} + u_{1t}
\]

(21)

\[
M_{2t} = \alpha + \phi_2 (r - r^*)_{t-1} + \sum_{i=2}^{j} \beta_i \Delta(r - r^*)_{t-i+1} + u_{2t}
\]

(22)

\[
\theta_t = pr \{ \tau^L \leq (r - r^*)_{t-d} \leq \tau^U \} = 1 - \frac{1}{1 + e^{-\sigma[(r-r^*)_{t-d} -\tau^L]\|(r-r^*)_{t-d} -\tau^U\]}}
\]

(24)

The QL-STECM distinguishes between two RIRD-adjustment regimes, an inner \((M_{1t})\) and an outer \((M_{2t})\), respectively given by equations (21) and (22). These are RIRD-adjustment equations similar to the linear ADF equation given by (10). At any point in time \(t\), the actual adjustment of \((r - r^*)_t\) is given by equation (20) as a weighted average of \(M_{1t}\) and \(M_{2t}\), where \(\theta\) is the probability that the transition variable takes values in the inner regime modelled using the quadratic function as per equation (24). In (24), \(\tau^U\) and \(\tau^L\) respectively denote the upper and lower threshold defining the inner adjustment regime and \(d\) is the delay parameter of the transition function determined by the highest value of the \(LM^G\) statistic reported in Table 3. The QL-STECM captures size effects by allowing the speed of adjustment in the inner regime to be slower than in the outer i.e. \(|\phi_1| < |\phi_2|\). Random walk behaviour in the inner regime is a special case in which \(\phi_1 = 0\) and \(\phi_2 < 0\). On the other hand, the QL-STECM captures sign effects by

---

11 The QL-STECM model is almost identical to the E-STAR model, the only difference being that unlike the latter, it allows for asymmetric regime threshold values.
allowing the transition function to enter one of the outer regimes faster than the other, that is by allowing for asymmetric regime thresholds \((\tau^U + \tau^L \neq 0)\).\(^{12,13}\)

Table 4 presents the results of the QL-STECh models for the countries and sub-periods for which the linearity hypothesis was rejected in Table 3.\(^{14}\) For those countries for which the linearity hypothesis was maintained, Table 4 reports the results of the estimation of the linear model (10), in which \(\phi_1 = \phi_2 = \phi\). In those cases for which linearity was rejected in Table 3 we obtain in all but two exceptions, namely Poland and Slovenia for the pre-2003(5) period, statistically significant mean reversion in the outer regime towards a non-significant constant, with zero included in the inner-regime band defined by the estimated upper and lower inner-regime thresholds. With the exception of Latvia, within the inner regime the \((r - r^*)\) series follows a random walk as \(\phi_1\) is statistically equal to zero. For Latvia, we obtain mean reversion in both regimes but the speed of adjustment towards the statistically zero mean prevailing in the inner regime is only half compared to the speed prevailing in the outer. On the other hand, for Poland and Slovenia during the pre-2003(5) period, we find non-linear mean reversion towards a slightly positive and slightly negative mean value respectively, with the estimated inner-regime bands respectively consisting of positive and negative values only. Finally, for the countries for which linearity was maintained, we obtain stationarity around a

\(^{12}\) In \(\phi_i = \phi\) and \(\beta_i = \beta\) \(\forall i \in (1...j)\) the model simplifies to the linear ADF equation in (10), as behaviour in both regimes is the same. If \(\tau^U + \tau^L = 0\) the model simplifies to an E-STAR type model, as the thresholds defining the inner regime are symmetric. Finally, if \(\tau^U - \tau^L = 0\) the two adjustment regimes are distinguished between values of the transition variable above and below a unique threshold real interest rate value rather than between values falling without or outside an inner-adjustment band.

\(^{13}\) Size- and sign-adjustment effects can be simultaneously modelled using a 3-regime STECM, involving three regimes, an inner regime, an outer-upper regime and an outer-lower regime (see van Dijk, 2002). For the countries for which size and sign-effects were detected, we modelled RIRD adjustment using such a model but the results were unsuccessful.

\(^{14}\) Although linearity was rejected for Estonia and Slovakia during the periods 1996(4)-1999(6) and 1996(4)-1999(4) respectively, we do not present estimates for these periods as the small number of available observations did not allow us to obtain convergence for our non-linear models.
zero mean in the cases of Cyprus and Malta, and stationarity around a slightly positive mean in the cases of Hungary and Slovenia during the post-2005 period.

In addition to the above, Table 4 offers a number of additional interesting insights. First, for the countries for which linearity was rejected in Table 3, in all countries except from the Czech Republic and Estonia we obtain evidence of sign effects in RIRD adjustment, as the point estimates of the regime thresholds appear significantly asymmetric (i.e. \( \tau_U + \tau_L \neq 0 \)) in the cases of Latvia, Lithuania, Slovakia, Poland and Slovenia for the pre-2003(5) period. These asymmetries are not uniform across countries, as in the cases of Latvia, Slovakia and Slovenia \( \tau_U + \tau_L < 0 \), whereas in the case of Lithuania and Poland \( \tau_U + \tau_L > 0 \). Second, the width of inner regime band is not uniform across countries, as the estimated band is relatively narrow in the cases of the Czech Republic, Estonia and Slovenia, and more extended for Latvia, Lithuania, Poland and Slovakia.

Second, Table 4 suggests significant differences relating the persistence of deviations of the RIRD from its long-run equilibrium in the outer regime for the countries for which non-linearities were found, and the single adjustment regime for which linearity was maintained in Table 3. More specifically, adjustment to equilibrium varies greatly, with the estimated speed of adjustment taking values ranging from -0.12 in the case of Poland to -0.84 in the case of Slovenia during the post-2003(5) period. If the degree of persistence of deviations from RIRP is a measure of the financial and general economic integration these countries have achieved against the EMU average, thereby providing indications relating to the degree to which the NACs are ready to adopt the single currency, the differences reported in Table 4 suggest that some of these countries are more ready than others to enter the Eurozone, with Slovenia and Cyprus being at the forefront and Hungary and Poland having more ground to cover.
Overall, the results reported in Table 4 are generally speaking favourable towards RIRP, as for seven out of ten countries we obtain mean reversion towards a statistically zero mean value (in five out of these seven cases, this reversion is subject to non-linear effects), while for the remaining three (Hungary, Poland and Slovenia) we obtain stationarity around non-zero values which, however, are not too far away from zero. Hence, the results of our non-linear analysis overturn the negative findings of our benchmark linear ADF analysis in section 5.1, where the majority of the \((r-r^*)\) series were found to include unit roots. At the same time, the results reported in Table 4 establish the existence of significant idiosyncratic behaviour in the RIRD of the NACs. This suggests that these countries still provide opportunities for portfolio diversification to international investors on the one hand, and have not achieved uniform convergence progress in their effort to join the single currency on the other.

### 7. BLOCK BOOTSTRAP ANALYSIS

We end our econometric investigation by testing for RIRP using an innovative procedure based on the block bootstrap. This aims to address simultaneously the downwards bias involved in determining the lag-length of the standard ADF test using information criteria such as the AIC on the one hand and the loss of power involved in increasing the lag structure of the ADF equation on the other.

The standard ADF equation described by (10) specifies the null hypothesis in terms of the series being generated by a non-stationary process. Rejection of the null of non-stationarity requires \(\phi\) in (10) to be negative and significantly different from zero. The critical values of the test statistic \(t_c\) used to determine the significance of the \(\phi\) parameter are assumed to follow a student t-distribution. The lagged terms
\[ \sum_{j=2}^{l} \beta_j \Delta(r - r^*)_{t-j+1} \] are included to remove any serial correlation in \( u_t \). For any given number of \( n \) lagged values of \( \Delta(r - r^*) \) the asymptotic distribution of \( t_c \) will depend on the pattern of serial correlation in \( u_t \). If \( n \) is allowed to grow with the sample size at a suitable rate, the dependence will disappear and the asymptotic distribution will be the same as if there were no serial correlation; (see Galbraith and Zinde-Walsh, 1999). In practice, however, information criteria such as the AIC may result in a downward bias in determining \( n \) (see Ng and Perron (2001); on the other hand as \( n \to \infty \), the power of the ADF test decreases (see Lopez, 1997).

Therefore, given that the critical values of \( t_c \) may be inaccurate, we construct our own critical values using a bootstrap simulation. Given that we have to take into account of serial correlation in the model, we construct a block bootstrap (see among others Davison and Hinkley 1997 and MacKinnon, 2002). Blocks of length \( l > 1 \) are constructed as follows. The first block is \( [\Delta y_1, \ldots, \Delta y_l] \), the next \( [\Delta y_{l+1}, \ldots, \Delta y_{2l}] \) and so on, where we define \( y_t = (r - r^*)_t \). Assuming \( L = T/l \) is an integer, there are \( L \) such blocks. The bootstrap then resamples each block with probability \( 1/L \) and concatenates them into a bootstrap sample of length \( T \).

If the length of the block is sufficiently large for blocks to be pair-wise independent, the bootstrap will replicate most of the time-series properties of the original data. However, the problem with a large block is that it reduces the variation in re-sampled values, resulting in a poor description of the distribution of the test statistic used to determine statistical inferences. The problem with reduced variation within replications may be reduced by using overlapping blocks combined with a wrapping of the sample near the beginning and end to account for the problem that overlapping
blocks cause sample observations at the beginning and end being sampled less often than observations near the centre (Davison and Hinkley, 1997).

A problem with the moving blocks bootstrap is that the re-sampled blocks are not stationary. To address this we use the stationary bootstrap (Politis and Romano, 1994). The idea behind the stationary bootstrap is that $l$ is generated from a geometric distribution with probability $P(l = j) = q(1 - q)^{j-1}$, where $0 < q < 1$ is a parameter. A new block length is drawn from each block and results in blocks having average length $q^{-1}$.

Thus, our procedure amounts to resampling blocks of the $y_i$'s using stationary blocks bootstrap. Like Davison and Hinkley (1997) we use a pragmatic procedure to determine $q$. We compare for several values of $q$ the estimated serial correlations of $u_t$ in (25) from the original data as well as from the bootstrap simulations. The experiment reveals that the stationary overlapping blocks bootstrap generates regression coefficients very similar to the coefficients estimated on the original sample for both high and low block lengths. However, the serial correlation properties of the bootstrapped data only match those of the original data if the block length $l$ is sufficiently large (around 5). Thus, for our empirical analysis we choose $q^{-1} = 6$.

We generate 10,000 sets of stationary overlapping blocks bootstrap of average block length equal to 5. Consequently, for each bootstrap iteration a series of artificially constructed DF tests are constructed under the null hypothesis $\phi = 0$ so that

$$\Delta y_{t}^* = \alpha^* + \sum_{i=1}^{n} \beta_i \Delta y_{t-i}^* \quad \text{for } t = 1 \ldots 10,000 \quad (25)$$

---

15 The empirical results however, are not sensitive to the choice of $q^{-1}$. Results (available upon request) for $q^{-1} = 5$ and $q^{-1} = 7$ are similar to the results for $q^{-1} = 6$. 

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25
As can be seen from (25), the generated sequence of artificial data has a true $\phi$ coefficient of zero. However, when we regress the artificial DF test for a given bootstrap sample $0_t$ estimated values of $\phi$ that differ from zero will result. This procedure provides an empirical distribution for $\phi$ and their associated standard errors based exclusively on the resampling of the residuals from the original regression displayed in equation (1). The idea in 10,000 replications is to determine appropriate critical values for the t-test so we do not reject the null of $\hat{\phi} = 0$. These critical values can be used to determine whether the estimates of $\phi$ obtained in (10) reject the null hypothesis of non-stationarity.

Table 5 presents the 90, 95 and 99% confidence intervals obtained from the stationary overlapping block bootstraps for the estimated t-statistics of the DF test including a constant as well as a constant and a trend term using the whole of the sample period available, i.e. 1996(1)-2005(12). These are compared with the values of the reported DF-tests; if the latter are lower than the lower limit of the reported block bootstrap confidence intervals, the unit root hypothesis of no-cointegration is rejected. Using the block bootstrap we reject the unit root hypothesis in all cases except from Poland. Overall, in contrast to the findings of the standard ADF tests presented in section 5.1, the results of our block bootstrap analysis provide significant evidence in favour stationary RIRD behaviour. The results of our block bootstrap analysis are consistent with those of our non-linear analysis presented in section 6 and provide additional evidence in favour of the hypothesis that RIRD against the EMU average in the NACs are mean-reverting processes.

8. SUMMARY AND CONCLUDING REMARKS

16 We report the confidence intervals rather than the p-values, to allow a more powerful test in the presence of extreme outliers in the data.
This paper has tested for real interest rate parity (RIRP) in the ten countries that joined the EU in May 2004 using a number of econometric approaches based on the time series properties of their ex-post real interest rate differentials (RIRD) against the EMU average, calculated using three-month nominal interest rates and consumer price indexes of identical definition. Our analysis is based on observations of monthly frequency and covers the period 1996-2005, a period during which the ten newly-accession EU countries (NACs) had abolished the majority of capital controls and had overcome the initial shock caused by the introduction of market reforms in the early 1990s. We test for RIRP using a number of econometric approaches, including standard unit root test on RIRD, unit root tests allowing for structural break, unit root tests capturing size- and sign non-linear RIRD adjustment effects and, finally, an innovative testing procedure based on the block-bootstrap.

Our econometric investigation showed that the RIRD of all ten NACs against the EMU average are characterised by mean-reverting behaviour. More specifically, we find that in seven out of 10 cases examined, RIRD converge towards a zero equilibrium value and in the three remaining cases towards levels that do not deviate from zero substantially. In addition, for seven countries we find strong evidence of non-linear adjustment towards the long-run RIRD equilibrium value, involving size- and sign-RIRD adjustment effects. Furthermore, in three countries the RIRD process is found to have been subject to a structural break. Finally, and very importantly, we find that the persistence of short-run deviations of the RIRD from its long-run equilibrium value differs significantly across countries, which suggests that existence of strong idiosyncratic elements in the movements of national RIRD.

Overall, our findings are generally supportive towards RIRP and suggest increasing yet varied integration between the money markets of the newly-accession EU
countries and the EMU. This has important implications both for the purposes of international portfolio investment as well as for the contact of national monetary policies towards adopting the Euro. More specifically, our findings indicate that the newly-accession EU countries provide opportunities for portfolio diversification to international investors on the one hand, and have not achieved uniform convergence progress in their effort to join the single currency on the other.

REFERENCES


### Table 1 – ADF unit root tests

<table>
<thead>
<tr>
<th>Country</th>
<th>ADF-test</th>
<th>Akaike score</th>
<th>$\alpha$</th>
<th>$\rho$</th>
<th>ADF-test</th>
<th>Akaike score</th>
<th>$\alpha$</th>
<th>$t$</th>
<th>$\rho$</th>
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<td>-4.288 (10)**</td>
<td>-0.892</td>
<td>0.052</td>
<td>-0.690**</td>
<td>-4.477 (10)**</td>
<td>-0.890</td>
<td>-0.107</td>
<td>0.0026</td>
<td>-0.754**</td>
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<td>-2.864 (9)+</td>
<td>-1.437</td>
<td>0.067</td>
<td>-0.195</td>
<td>-3.267 (9)+</td>
<td>-1.443</td>
<td>0.262+</td>
<td>-0.0026</td>
<td>-0.255</td>
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<td>-1.467</td>
<td>0.010</td>
<td>-0.128</td>
<td>-2.868 (3)</td>
<td>-1.452</td>
<td>0.052</td>
<td>-0.0007</td>
<td>-0.128</td>
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<td>-1.813</td>
<td>0.158*</td>
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<td>-0.178</td>
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<td>-1.706</td>
<td>-0.306**</td>
<td>0.0035*</td>
<td>-0.243</td>
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</table>

NOTES: Parentheses denote the lag orders of the ADF test, determined by the Akaike information criterion; +, *, and ** respectively denote statistical significance at the 10, 5 and 11 per cent level. The statistical significance of the $\alpha$ and $t$ is coefficient is determined using the $t$-statistic critical values; the significance of the $\rho$ coefficient is determined using the ADF critical values.
Table 2.1: Perron endogenous break unit root test results, 1996(04)-2005(12)

<table>
<thead>
<tr>
<th>Countries</th>
<th>Model (16)</th>
<th></th>
<th></th>
<th></th>
<th>t-ratio</th>
<th>RSS</th>
<th>AIC</th>
<th>Break Date</th>
<th>Model (17)</th>
<th></th>
<th></th>
<th></th>
<th></th>
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<th>RSS</th>
<th>AIC</th>
<th>Break Date</th>
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<td></td>
<td>t-ratio</td>
<td>RSS</td>
<td>AIC</td>
<td>Break Date</td>
<td>t-ratio</td>
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</table>

NOTES: The numbers in square brackets shows the number of lagged difference terms (k) in Models (16) and (17). It was chosen by the ‘t-sig’ approach. The break date was chosen by minimizing the t-statistic for testing $\gamma_4 = 1$. The reported t-ratio tests the null of unit root-no break (see Perron, 1997, for the critical values). RSS and AIC denote the residual sum of squares and Akaike information criterion. **, *, + indicate rejection of the null-unit root with no-break at the 1%, 5%, 10% level of significance, respectively.

Table 2.2: Estimated parameters for preferred models from Perron endogenous break unit root tests.

<table>
<thead>
<tr>
<th>Countries</th>
<th>Preferred Model</th>
<th>$\gamma_0$</th>
<th>$\gamma_1$</th>
<th>$\gamma_2$</th>
<th>$\gamma_3$</th>
<th>$\gamma_4$</th>
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<tr>
<td>Estonia</td>
<td>Model (17)</td>
<td>-1.93 **</td>
<td>2.00 **</td>
<td>0.08 **</td>
<td>0.35</td>
<td>0.58 **</td>
<td>-0.08 **</td>
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<td>Model (17)</td>
<td>1.07</td>
<td>-0.89</td>
<td>0.06 +</td>
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<td>-0.00004</td>
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NOTES: The preferred model minimizes the RSS and AIC in Table 2.1. **, *, + indicate statistical significance at the 1%, 5%, 10% level, respectively.
Table 3
Tests for non-linear interest rate differential adjustment

<table>
<thead>
<tr>
<th></th>
<th>φ</th>
<th>d</th>
<th>LM^t</th>
<th>LM^l</th>
<th>LM^b</th>
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<td>Cyprus</td>
<td>11</td>
<td>10</td>
<td>0.808 [0.743]</td>
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<td>N/A</td>
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<td>3</td>
<td>4.003 [0.000]**</td>
<td>6.161 [0.000]**</td>
<td>5.295 [0.00]**</td>
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<td>1</td>
<td>12</td>
<td>3.179 [0.028]*</td>
<td>0.391 [0.681]</td>
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<td></td>
<td>1999 (7) – 2005 (12)</td>
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<td>11</td>
<td>3.225 [0.001]**</td>
<td>2.203 [0.069]+</td>
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<td>1</td>
<td>1.945 [0.127]</td>
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<td>N/A</td>
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<tr>
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<td>2</td>
<td>2.623 [0.028]*</td>
<td>0.007 [0.993]</td>
<td>4.488 [0.000]**</td>
</tr>
<tr>
<td>Lithuania</td>
<td>11</td>
<td>9</td>
<td>1.791 [0.024]*</td>
<td>1.764 [0.075]+</td>
<td>2.210 [0.008]**</td>
</tr>
<tr>
<td>Malta</td>
<td>4</td>
<td>3</td>
<td>1.557 [0.118]</td>
<td>N/A</td>
<td>N/A</td>
</tr>
<tr>
<td>Poland</td>
<td>4</td>
<td>9</td>
<td>1.877 [0.040]*</td>
<td>1.643 [0.157]</td>
<td>2.567 [0.011]**</td>
</tr>
<tr>
<td>Slovakia</td>
<td>1996(4) – 1999(4)</td>
<td>1</td>
<td>3</td>
<td>3.776 [0.010]*</td>
<td>8.337 [0.002]**</td>
</tr>
<tr>
<td></td>
<td>1999 (5) – 2005 (12)</td>
<td>11</td>
<td>12</td>
<td>2.725 [0.007]**</td>
<td>1.539 [0.185]</td>
</tr>
<tr>
<td>Slovenia</td>
<td>1996(4) - 2003(4)</td>
<td>2</td>
<td>1</td>
<td>3.484 [0.004]**</td>
<td>0.263 [0.852]</td>
</tr>
<tr>
<td></td>
<td>2003(5) – 2005 (12)</td>
<td>2</td>
<td>3</td>
<td>1.771 [0.150]</td>
<td>N/A</td>
</tr>
</tbody>
</table>

NOTES: Numbers in square brackets denote p-values; +, *, ** denote significance at the 10%, 5% and 1% level respectively.
## Table 4

### Non-linear QL-STECP models

<table>
<thead>
<tr>
<th>Country</th>
<th>α</th>
<th>$\phi_1$ (inner regime)</th>
<th>$\phi_2$ (outer regime)</th>
<th>$\tau^L$</th>
<th>$\tau^U$</th>
</tr>
</thead>
<tbody>
<tr>
<td>Cyprus</td>
<td>0.052 (0.060)</td>
<td>N/A</td>
<td>-0.690 (0.161)**</td>
<td>N/A</td>
<td>N/A</td>
</tr>
<tr>
<td>Czech Republic</td>
<td>0.051 (0.053)</td>
<td>0.056 (0.158)</td>
<td>-0.207 (0.057)**</td>
<td>-0.326 (0.003)**</td>
<td>0.401 (0.048)**</td>
</tr>
<tr>
<td>Estonia</td>
<td>1999 (7) – 2005 (12)</td>
<td>-0.035 (0.049)</td>
<td>-0.075 (0.094)</td>
<td>-0.439 (0.178)**</td>
<td>-0.607 (0.048)**</td>
</tr>
<tr>
<td>Hungary</td>
<td>0.158 (0.066)*</td>
<td>N/A</td>
<td>-0.173 (0.059)**</td>
<td>N/A</td>
<td>N/A</td>
</tr>
<tr>
<td>Latvia</td>
<td>-0.066 (0.051)</td>
<td>-0.142 (0.073)*</td>
<td>-0.381 (0.121)**</td>
<td>-1.529 (0.045)**</td>
<td>0.440 (0.158)**</td>
</tr>
<tr>
<td>Lithuania</td>
<td>0.023 (0.043)</td>
<td>-0.032 (0.067)</td>
<td>-0.259 (0.105)*</td>
<td>-0.524 (0.081)**</td>
<td>1.222 (0.399)**</td>
</tr>
<tr>
<td>Malta</td>
<td>0.011 (0.034)</td>
<td>N/A</td>
<td>-0.391 (0.110)**</td>
<td>N/A</td>
<td>N/A</td>
</tr>
<tr>
<td>Poland</td>
<td>0.131 (0.070)+</td>
<td>-0.069 (0.038)+</td>
<td>-0.124 (0.061)*</td>
<td>1.175 (0.002)**</td>
<td>3.452 (0.004)**</td>
</tr>
<tr>
<td>Slovakia</td>
<td>1999 (5) – 2005 (12)</td>
<td>0.031 (0.058)</td>
<td>0.034 (0.823)</td>
<td>-0.451 (0.156)**</td>
<td>-3.864 (0.397)**</td>
</tr>
<tr>
<td>Slovenia</td>
<td>1996(4) - 2003(4)</td>
<td>-0.225 (0.076)**</td>
<td>-0.084 (0.217)</td>
<td>-0.509 (0.085)**</td>
<td>-0.352 (0.012)**</td>
</tr>
<tr>
<td></td>
<td>2003(5) – 2005 (12)</td>
<td>0.448 (0.176)*</td>
<td>N/A</td>
<td>-0.837 (0.175)**</td>
<td>N/A</td>
</tr>
</tbody>
</table>

NOTES: +, * and ** respectively denote statistical significance at the 10, 5 and 1 per cent level respectively; for Cyprus, Hungary, Malta and Slovenia 2003(5)-2005(12), for which the linearity hypothesis is maintained in Table 3, we report the estimates of the linear equation (10). Although linearity was rejected for Estonia and Slovakia during the periods 1996(4)-1999(6) and 1996(4)-1999(4) respectively, we do not present estimates for these periods as the small number of available observations did not allow us to obtain convergence for our non-linear models.
### Table 5 – Block bootstrap analysis

<table>
<thead>
<tr>
<th></th>
<th>Cyprus</th>
<th>Czech Rep.</th>
<th>Estonia</th>
<th>Hungary</th>
<th>Latvia</th>
<th>Lithuania</th>
<th>Malta</th>
<th>Poland</th>
<th>Slovakia</th>
<th>Slovenia</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>DF with constant</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Autocorrelation p-value</td>
<td>0.00</td>
<td>0.00</td>
<td>0.00</td>
<td>0.09</td>
<td>0.00</td>
<td>0.00</td>
<td>0.00</td>
<td>0.00</td>
<td>0.00</td>
<td>0.00</td>
</tr>
<tr>
<td>AIC</td>
<td>-0.592</td>
<td>-1.334</td>
<td>-1.175</td>
<td>-1.629</td>
<td>-1.202</td>
<td>-0.778</td>
<td>-1.795</td>
<td>-1.623</td>
<td>0.406</td>
<td>-1.297</td>
</tr>
<tr>
<td>90% BB CI</td>
<td>(-2.62, 2.70)</td>
<td>(-2.63, 2.71)</td>
<td>(-2.67, 2.71)</td>
<td>(-2.65, 2.70)</td>
<td>(-2.67, 2.71)</td>
<td>(-2.65, 2.70)</td>
<td>(-2.67, 2.70)</td>
<td>(-2.69, 2.68)</td>
<td>(-2.66, 2.70)</td>
<td>(-2.66, 2.70)</td>
</tr>
<tr>
<td>95% BB CI</td>
<td>(-2.81, 2.87)</td>
<td>(-2.83, 2.89)</td>
<td>(-2.82, 2.83)</td>
<td>(-2.85, 2.80)</td>
<td>(-2.84, 2.81)</td>
<td>(-2.83, 2.84)</td>
<td>(-2.87, 2.87)</td>
<td>(-2.85, 2.84)</td>
<td>(-2.85, 2.85)</td>
<td>(-2.85, 2.85)</td>
</tr>
<tr>
<td>99% BB CI</td>
<td>(-2.93, 2.97)</td>
<td>(-2.94, 2.98)</td>
<td>(-2.95, 2.94)</td>
<td>(-2.96, 2.95)</td>
<td>(-2.94, 2.92)</td>
<td>(-2.95, 2.93)</td>
<td>(-2.96, 2.95)</td>
<td>(-2.93, 2.95)</td>
<td>(-2.92, 2.95)</td>
<td>(-2.92, 2.94)</td>
</tr>
<tr>
<td><strong>DF with constant and trend</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
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</tr>
<tr>
<td>Autocorrelation p-value</td>
<td>0.00</td>
<td>0.01</td>
<td>0.00</td>
<td>0.04</td>
<td>0.00</td>
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<td>0.00</td>
<td>0.00</td>
<td>0.00</td>
<td>0.00</td>
</tr>
<tr>
<td>AIC</td>
<td>-0.580</td>
<td>-1.346</td>
<td>-1.168</td>
<td>-1.621</td>
<td>-1.194</td>
<td>-0.789</td>
<td>-1.788</td>
<td>-1.617</td>
<td>0.380</td>
<td>-1.280</td>
</tr>
<tr>
<td>90% BB CI</td>
<td>(-2.62, 2.70)</td>
<td>(-2.63, 2.71)</td>
<td>(-2.67, 2.71)</td>
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<td>(-2.84, 2.81)</td>
<td>(-2.83, 2.84)</td>
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<td>(-2.92, 2.94)</td>
</tr>
</tbody>
</table>