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REAL CONVERGENCE AND REGIME-SWITCHING AMONG EU ACCESSION COUNTRIES

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Abstract
Real convergence among the ten EU 2004 accession economies is investigated with respect to long-run real interest parity. We employ a novel approach where unit-root tests for real interest differentials are embedded within a Markov regime-switching framework. Whereas standard univariate unit-root tests provide mixed support for parity, we find parity is present in all cases where differentials either switch between regimes of stationary and non-stationarity behavior, or between alternative regimes of stationarity characterized by differing degrees of persistence. Further insights are obtained from the inferred probabilities of being in each regime, and the regime-switching nature of the differential variances.

JEL Classification: C330, E430, F300, G150
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1. Introduction

The extent to which real interest rates are equalized across countries has occupied researchers for a number of reasons. In an open economy, real interest parity (RIP) provides an indication of whether countries are economically and financially integrated or autonomous. RIP is also important as a key working assumption in various models of exchange rate determination and the focus of many studies as early as Frenkel (1976), Mussa (1976) and Frankel (1979). The purpose of this paper is to examine how plausible RIP is for the ten European countries that secured entry to the European Union (EU) in May 2004. Full membership of the European Economic and Monetary Union (EMU) requires that various criteria for nominal and real convergence are met. Evidence against RIP, or real convergence, would suggest that the accession countries may not be suitable candidates for participation in a common EU monetary policy. Evidence on the extent of real convergence for the EU accession countries is mixed. For example, Kozluk (2005) finds that the accession countries are better prepared for the single currency membership than some of the more established members were at the introduction of the Euro. However, Babetskii *et al*. (2004) find evidence in favor of nominal rather than real convergence.

The key contribution of our paper is with respect to the econometric approach we employ. We consider whether or not adjustments towards RIP are nonlinear. For this purpose, our investigation of RIP is based on Augmented Dickey Fuller (ADF) unit root testing of real interest rate differentials defined against the US and Germany within a Markov regime-switching framework. This MS-ADF approach is in sharp contrast to methodologies pursued in existing studies of RIP and offers three valuable new insights into the RIP debate where the time-series properties of interest differentials are modeled as regime-dependent. First, existing studies of RIP compute a single test statistic for testing non-stationarity making no distinction between regime-dependent behavioral characteristics. This approach can lead to a bias towards accepting the non-stationary null thereby rejecting RIP, or give a false impression of the speed of adjustment towards long-run equilibrium, because there is no distinction between alternative regimes. Second, we are able to estimate a regime-specific variance for the real interest differential. The general econometric approach taken in the literature on RIP assesses the dynamics of integration on the first moment of real interest rate distributions. While the contribution of these tests to the literature is clearly significant, it is important to consider the informational content of higher moments as well. In terms of second moments, increased financial integration might also be consistent with a reduced variance attached to real interest differentials. Third, the MS-ADF methodology allows the researcher to estimate the inferred probability of the real interest differential being subject to a particular regime in any time period. We can, for example, consider at what time the progress of the accession countries towards EU membership is most likely to have led to a shift towards regimes of stationary behavior from regimes of non-stationary behavior.

The paper is organized as follows. The following section discusses the rationale and literature on RIP. Generally speaking, evidence in favor of RIP is patchy. The third section sets out the econometric model and explains how the notion of regime-switching and RIP is addressed. The fourth section discusses the data and results. In contrast to single-regime ADF testing, our results indicate that the most of the sample are characterized by just one stationary regime, or two distinct stationary regimes. The final section offers our conclusions.

2. Real interest rate parity

Several studies have demonstrated the proposition that RIP can be derived from the interaction of uncovered interest parity (UIP), PPP and the Fisher condition.[1] Of these relationships, Chung and Crowder (2004) argue that UIP is the most commonly violated on account of non-stationary risk premia in the foreign exchange markets. The validity of the RIP relationship has been tested using various econometric approaches and provides conflicting empirical evidence. Early attempts using ordinary least squares regressions find evidence against long-term RIP [for example, Mishkin (1984) and Cumby and Mishkin (1986)]. Tests of a long-run cointegrating relationship between domestic and foreign real rates of interest provide evidence of weak-form RIP among OECD countries (for example, Goodwin and Grennes (1994) and Moosa and Bhatti (1996), among others). Studies such as Wu and Chen (1998) examine real interest differentials and find that a series of alternative unit root tests that are more powerful than the conventional ADF tests leads to the rejection of non-stationarity. Fountas and Wu (1999), on the other hand, find evidence in favor of strong RIP in the EU member countries using unit root tests that allow for structural breaks in the series.

More recent literature has also considered the possible nonlinear dynamics of realignment towards RIP, thereby building on the work of Cavaglia (1992) who examines the changing patterns in the behavior of real interest differentials over time. Mancuso *et al*. (2003) and Holmes and Maghrebi (2004) report evidence from smooth transition autoregressive models that indicates the presence of a series of thresholds straddling the equilibrium conditions so that arbitrage opportunities increase with larger deviations from parity against a background of transaction costs. There is also a growing strand of literature based on Markov-regime switching models. Dahlquist and Gray (2000) find that the speed of adjustment of nominal interest rates in the European monetary system is stronger during periods of high interest rates and high volatility. Kanas (2003), on the other hand, uses a bivariate Markov-switching vector error-correction model which accounts for both the regime switches in the UK and US real interest rates and their long-run cointegration properties over the postwar period.

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1. See, for example, MacDonald and Taylor (1989).

3. The econometric model

Existing studies such as Gray (1996) and Ang and Bekaert (2002) consider and analyze regime-dependence in the mean-reversion of interest rate *levels* where the focus is on OECD countries. In this paper, we build on this literature by presenting a Markov regime-switching ADF unit root test for RIP which is to be applied to the ten accession countries. Suppose we denote the real interest differential between country *i* and a given base country as y_i. The usual test for linear adjustment towards long-run RIP is based assessing the unit root properties of y_i through the OLS estimation of ADF regressions such as

$$\Delta y_t = \varsigma + \lambda y_{t-1} + \sum_{i=1}^{k} \psi_i \Delta y_{t-i} + v_t \tag{1}$$

where v_t is a white noise residual. Here we find $-2 < \lambda < 0$ is indicative of stationarity of the real interest rate differential and supportive of long-run RIP. Now, suppose a discrete random variable S_t takes two possible values $S_t = (0,1)$ and serves as an indicator for the state of the world economy at time *t*. The expected component of Δy, conditional on the value of S_t, is given as follows.

$$\begin{aligned} E(\Delta y_t \mid S_t) = {} & [(1-S_t)\alpha_0 + S_t\alpha_1] + (1-S_t)\lambda_0 y_{t-1} + S_t\lambda_1 y_{t-1} \\ & + (1-S_t)\sum_{i=1}^{l} \xi_i \Delta y_{t-i} + S_t \sum_{i=1}^{l} \tau_i \Delta y_{t-i} + \varepsilon_t \end{aligned} \tag{2}$$

where $\varepsilon_t \sim i.i.d.N(0,\sigma^2(S_t))$ which allows for the variance to change across regimes and the unobserved indicator variable, S_t, evolves according to the first-order Markov-switching process described in Hamilton (1989):

$$\begin{aligned} & P[S_t = 0 \mid S_{t-1} = 0] = P_0, P[S_t = 1 \mid S_{t-1} = 0] = 1 - P_0 \\ & P[S_t = 1 \mid S_{t-1} = 1] = P_1, P[S_t = 0 \mid S_{t-1} = 1] = 1 - P_1 \end{aligned} \tag{3}$$

where P_0 and P_1 are the fixed transition probabilities of being in Regime 0 or 1 respectively where $0 < P_0, P_1 < 1$. Estimation of the model is carried out by maximum likelihood using the non-linear filter algorithm described in Hamilton's (1989). An important by-product of the estimation procedure is the computation of the filter probabilities - the probabilities of being in a particular regime at time *t*.

Stationary behavior in both regimes is confirmed when $-2 < \lambda_0, \lambda_1 < 0$. If $-1 < \lambda_0, \lambda_1 < 0$, the half-life associated with a deviation from long-run equilibrium may be approximated. In cases where $l = 0$, we may approximate the regime-specific half-lives as $HL_0 = (\ln 0.5)/(1+\lambda_0)$ and $HL_1 = (\ln 0.5)/(1+\lambda_1)$ for Regimes 0 and 1 respectively. However, where $l > 0$ we note the caution advocated by Seong *et al*.

(2006) and others that applying the same formula in the presence of higher order autoregressive processes above AR(1) may lead to bias in the half-life approximation. We therefore follow Seong *et al*. (2006) and modify our calculations accordingly. If $\lambda_0 \neq \lambda_1$, we may define the concept of *varied RIP* because stationary behavior is confirmed across the entire study period, but the autoregressive coefficients and speeds of adjustment towards long-run equilibrium are different. On the other hand, we may only be able to confirm that either λ_0 or λ_1 is significantly different from zero. In this case, we may define the concept of *partial RIP* because the real interest differential is switching between regimes of stationary and non-stationary behavior.

At this point it is helpful to reflect on the other issues associated with the MS-ADF test. The usual drawback with Markov regime-switching is that the interest rate differential must be within a single regime in each time period where there is a sharp switch between regimes. This may be contrasted with the smooth transition generalization of threshold class models that allow for the possibility that the interest rate differential might be in some intermediate state between regimes where the nature of adjustment varies with the extent of deviation from equilibrium. Nonetheless, the MS-ADF test has featured in a number of papers that includes Hall and Sola (1993), Gray (1996), van Norden and Vigfusson (1998), Hall *et al*. (1999), Kanas (2005) and Kanas and Genius (2005). With regard to these studies, Kanas and Genius (2005) report that the MS-ADF test has superior ability over the standard ADF test in terms of distinguishing a sometimes integrated model from an I(1) process.

4. Data and results

The data used in this study consist of monthly observations on three month deposit rates and annualized inflation rates estimated from the consumer price indices for the ten European countries that joined the EU on the 1st May 2004. The countries are Cyprus (CP), Czech Republic (CZ), Estonia (EO), Hungary (HN), Latvia (LV), Lithuania (LN), Malta (MA), Poland (PO), Slovakia (SX) and Slovenia (SJ). The real interest rate differentials are defined with respect to the United States (US) and Germany (GM). All the data are retrieved from the International Financial Statistics database where data availability dictates a start date of July 1993. The study period finishes in December 2005.[2] Frankel (1992) and others have argued that RIP and the degree of capital mobility are closely connected. The accession countries included in the sample differ in their liberalization status. According to Ems (2000) and Nerlich (2002), one can distinguish three groups. These are Estonia, Latvia, Lithuania and the Czech

2. Line I64 for the consumer price index and line I60L for deposit rate (the exception is the US where the corresponding line is I60LC). The use of the bank deposit interest rate data is dictated by the data availability across our large sample of countries. Treasury bill and interbank rates are not available on a consistent basis across the sample period.

Republic who have removed practically all exchange controls; Poland and Hungary who maintain restrictions on only short-term capital transactions; and the remaining countries who maintain a more comprehensive system of exchange control. On this basis, one would expect Estonia, Latvia, Lithuania and the Czech Republic to exhibit the greatest likelihood of RIP.

In common with many existing studies, we use data on *ex post* real interest rates. Moreover, rational expectations are assumed where inflation expected one year ahead is measured by actual inflation one year ahead plus a random forecast error. Use of the term structure of interest rates to extract information on expected inflation is ruled out because this study is concerned with onshore rates where data availability rules out consistent runs of monthly data for the desired maturity range over the entire study period. This same comment also applies to data for index-linked securities. Also, evidence on the information content of the term structure on inflation is mixed. For example, Mishkin (1991) analyses offshore rates for a sample of ten OECD economies and finds that the shorter maturity term structure does not contain a great deal of information about inflation. Koedijk and Kool (1995) examine German data and find evidence against the expectations hypothesis, concluding that there is no evidence that a steeper yield curve predicts higher inflation in the future. Conversely, Jorion and Mishkin (1991) and Gerlach (1997) use German data and find that the expectations hypothesis cannot be rejected. On the other hand, Schich (1999) finds that the information content is sensitive to maturity combination and sample period.

The first stage of the empirical investigation is to test whether the real interest rate differentials with respect to the US and Germany contain a unit root. This is done by implementing various unit root tests including the ADF, Ng-Perron (NP), DF-GLS unit root tests as well as the KSS test advocated by Kapetanios *et al*. (2003). The ADF test results reported in Table 1 are based on the exclusion of a deterministic trend. At the 5% significance level, the non-stationary null is accepted in the cases of the Czech Republic, Hungary and Poland differentials with respect to both the US and Germany, Malta with respect to US and Slovakia with respect to Germany. In all other cases, the non-stationary null is rejected leading us to initially conclude that evidence in favor of long-run RIP is limited. At the 5% significance level, the Ng-Perron (NP) and DF-GLS unit root tests are unable to provide any additional evidence in favor of stationarity. To rule out the possibility that the real interest rate may follow a nonlinear stationary process, we also conduct the KSS (2003) test on each of the differentials. Our results indicate that the KSS test rejects a unit root and in favor of nonlinear stationary processes in only seven out of twenty cases. It is only in the case of the Poland-US differential where the KSS indicates stationarity over and above the evidence provided by the earlier unit root tests.

Table1. Unit root tests on real interest rate differentials

Panel A: Real interest rate differentials vis-à-vis US				
	ADF	NP	DF-GLS	KSS
Cyprus	-4.005***	-0.400	-0.527	-2.045
Czech Republic	-2.280	-1.713*	-1.585	-2.808*
Estonia	-2.913**	-0.459	-0.525	-2.834*
Hungary	-2.048	-0.211	-0.296	-2.202
Latvia	-4.123***	1.132	0.465	-4.626***
Lithuania	-4.185***	0.935	-0.174	-2.311
Malta	-2.000	-1.769*	-1.569	-1.834
Poland	-2.560	-1.773*	-1.365	-3.493***
Slovakia	-4.464**	-0.720	-0.530	-2.705
Slovenia	-3.286**	-1.135	-1.723*	-5.451***
Panel B: Real interest rate differentials vis-à-vis Germany				
Cyprus	-3.187**	-0.487	-0.606	-2.123
Czech Republic	-2.351	-0.842	-0.736	-1.487
Estonia	-3.100**	-0.300	-0.360	-3.050**
Hungary	-2.594	-0.205	-0.176	-2.200
Latvia	-8.391***	1.448	0.880	-4.836***
Lithuania	-3.709***	0.379	0.093	-2.425
Malta	-3.735**	-0.569	-0.543	-3.422**
Poland	-2.504	-1.670*	-1.472	-1.749
Slovakia	-2.818	-0.852	-0.888	-2.107
Slovenia	-4.768**	-1.600	-2.507**	-5.638***

Notes: These are ADF, Ng-Perron (NP), DF-GLS and KSS (Kapetanios *et al*., 2003) unit root tests, respectively, conducted on interest rate differentials (excluding a deterministic trend). Lag length selection is based on BIC critical values. ***, ** and * denote rejection of the non-stationary null hypothesis at the 1, 5 and 10% significant levels respectively. The critical values for the ADF test at 1%, 5%, 10% levels without a deterministic trend are -3.46, -2.88 and -2.57 respectively.. The critical values for the DF-GLS test at 1%, 5%, 10% levels are -2.582, -1.943 and -1.615 respectively. The critical values for the Ng-Perron (NP) test at 1%, 5%, 10% levels are -2.58, -1.98 and -1.62 respectively. The critical values for KSS tests at 1%, 5% and 10% levels, for the de-meaned series, are -3.48, -2.93 and -2.66, respectively (Kapetanios *et al*., 2003).

A potential reason for finding limited evidence in favor of long-run RIP is that these initial ADF unit root tests do not incorporate regime switching between non-stationary and stationary behavior, or even between two alternative regimes of stationary behavior given the specific feature of these transition economies. Indeed, it might be the case that this latter type of regime-switching behavior is present despite the identification of stationarity in the Table 1 results. We therefore move on to consider results provided by the MS-ADF unit root test. Table 2 reports results from formally testing the null hypothesis of no regime switching against the alternative of regime switching. The null hypothesis is equivalent to the standard ADF regression, whilst the alternative hypothesis corresponds to the MS-ADF regression as shown in equations (2) and (3). Testing the null against the alternative hypothesis relies on a non-standard LR test. We therefore follow studies such as Garcia and Perron (1996), Kanas (2005), Kanas and Genius (2005) by utilizing the upper bound approach proposed by Davies (1987). In Markov models, testing for the number of regimes using the usual likelihood ratio test is problematic because the likelihood ratio test does not have the standard asymptotic distribution. The Davies test is based on the idea of giving a range of values to the parameters under the alternative hypothesis. Table 2 indicates that the log-likelihood value of the MS-ADF model is significantly higher than the value of the standard ADF model in eighteen out of twenty cases where the MS-ADF model is therefore preferred over the standard ADF model. It is only in the cases of the Cyprus-US and Hungary-US differentials where the null cannot be rejected at the 5% significance level. Regime-switching behavior is not detected in these two latter cases and they are excluded from further analysis.

Table 3 reports estimates obtained from the MS-ADF unit root testing model. With regard to the volatilities attached to each regime, we find the standard deviations σ_0 and σ_1 are statistically significant in all cases at the 5% level apart from the Czech Republic-US and Poland-Germany differentials where σ_1 is significant at the 10% level. The null $\sigma_0 = \sigma_1$ is rejected at the 5% significance level in fifteen out of eighteen cases. In each of these rejections, σ_0 is less than σ_1 suggesting that Regime 0 (Regime 1) can be identified as the low (high) volatility regime. In many cases the estimated probabilities are high, indicating that both regimes are characterized by a high expected duration. There are a number of noticeable cases where P_0 (the probability of remaining in Regime 0) has reached unity. This suggests that once the differential moves to the low-volatility Regime 0, it is unlikely to shift to the high-volatility Regime 1 over the sample period.

Table 2. Test for regime switching

	vis-à-vis US			vis-à-vis Germany		
	LL (Standard ADF)	LL (MS –ADF)	LR	LL (Standard ADF)	LL (MS –ADF)	LR
Cyprus	-155.61	-150.92	9.38 (0.095)	-180.13	-168.79	22.68 (0.000)
Czech Republic	-154.22	-131.20	46.05 (0.000)	-176.00	-148.30	55.4 (0.000)
Estonia	-201.71	-175.61	52.20 (0.000)	-224.07	-199.66	48.82 (0.000)
Hungary	-162.81	-158.32	8.980 (0.112)	-171.70	-154.44	34.52 (0.000)
Latvia	-311.41	-230.11	162.6 (0.000)	-312.67	-224.88	175.58 (0.000)
Lithuania	-278.19	-210.23	135.92 (0.000)	-287.24	-204.49	165.5 (0.000)
Malta	-151.31	-144.78	13.06 (0.000)	-155.59	-143.77	23.64 (0.000)
Poland	-167.52	-141.70	51.64 (0.000)	-197.20	-170.37	53.66 (0.000)
Slovenia	-321.11	-245.72	150.78 (0.000)	-321.83	-245.25	153.17 (0.000)
Slovakia	-206.62	-182.48	48.28 (0.000)	-216.49	-193.63	45.71 (0.000)

Notes: These results are based on testing the null hypothesis of no regime switching against the alternative of regime switching. LL denotes log likelihood and LR denotes likelihood ratio. The figures reported in the brackets are Davies (1987) upper bound p-values.

Table 3. Estimates of MS –ADF unit root test

Panel A: Real interest rate differentials vis-à-vis US

	CZ	EO	LV	LN	MA	PO	SJ	SX
α_0	-0.056 (-0.826)	-0.012 (-0.147)	-0.121 (-1.054	0.135 (1.105)	-0.034 (-0.473)	0.146 (3.679)	0.061 (0.540)	-0.021 (-0.240)
α_1	0.084 (0.195)	-0.225 (-1.190)	0.283 (0.322	-0.435 (-0.578)	0.162 (0.743)	1.180 (1.284)	2.077 (2.025)	0.729 (1.591)
λ_0	-0.031* (-2.092)	-0.042* (-9.555)	-0.139* (-2.742)	-0.179* (-16.364)	-0.080* (-3.614)	-0.017* (-2.801)	-0.075* (-5.129)	-0.057* (-3.257)
λ_1	-0.036 (-0.771)	-0.046* (-2.551)	-0.214* (-4.422)	-0.117* (-4.628)	0.105 (1.581)	-0.502* (-3.636)	-0.285* (-5.402)	-0.471* (-3.151)
σ_0	0.289 (7.453)	0.292 (5.383)	0.808 (5.142)	0.693 (6.057)	0.331 (3.743)	0.343 (6.569)	1.122 (5.402)	0.471 (4.850)
σ_1	2.478 (1.808)	2.573 (5.844)	28.372 (5.111)	20.113 (3.371)	0.958 (5.303)	0.831 (2.550)	31.714 (3.940)	4.453 (2.278)
P_0	0.992	0.999	0.999	1.000	0.999	0.994	1.000	0.919
P_1	0.000	0.999	0.999	0.988	0.973	0.095	0.999	0.000
HL_0	23.500	17.487	4.545	3.501	9.478	N/A	6.622	12.506
HL_1	N/A	16.489	2.503	6.477	N/A	1.481	1.496	1.471
$\alpha_0 = \alpha_1$	0.111	1.198	0.217	0.540	0.888	1.298	3.705	2.844
$\lambda_0 = \lambda_1$	0.011	0.031	1.663	5.201	7.893	12.599	15.280	8.297
$\sigma_0 = \sigma_1$	2.490	25.746	24.630	10.51	8.510	2.315	14.346	4.294

Panel B: Real interest rate differentials vis-à-vis Germany

	CP	CZ	EO	HN	LV	LN	MA	PO	SJ	SX
α_0	-0.881 (-24.904)	-0.920 (-22.814)	-0.484 (-8.224)	-0.414 (-6.214)	-0.447 (-4.457)	-0.384 (-3.430)	-0.851 (-12.937)	-0.082 (-0.847)	-0.620 (-5.151)	-0.348 (-10.782)
α_1	-0.755 (-8.189)	0.292 (1.931)	-0.707 (-2.997)	0.043 (0.168)	-1.241 (-1.757)	-1.204 (-1.408)	-1.804 (-7.782)	0.283 (0.804)	0.607 (0.630)	-1.088 (-4.845)
λ_0	-0.107* (-22.071)	-0.105* (-49.673)	-0.061* (-13.375)	-0.098* (-9.425)	-0.052* (-3.573)	-0.053* (-7.002)	-0.127* (-14.762)	-0.021 (-1.523)	-0.102* (-6.778)	-0.045* (-5.726)
λ_1	-0.174* (-13.787)	-0.017 (-1.492)	-0.072* (-4.005)	-0.009 (-0.329)	-0.286* (-3.057)	-0.146* (-4.272)	-0.297* (-11.830)	-0.354* (-5.906)	-0.305* (-3.856)	-0.237* (-5.629)
σ_0	0.085 (2.380)	0.261 (6.978)	0.201 (5.855)	0.313 (4.194)	0.714 (4.215)	0.557 (5.386)	0.337 (5.099)	0.517 (5.082)	1.097 (5.122)	0.439 (5.487)
σ_1	1.185 (4.601)	2.802 (2.682)	3.002 (3.984)	2.411 (4.098)	27.299 (6.825)	20.241 (2.376)	1.583 (2.537)	2.639 (2.016)	30.126 (3.276)	4.508 (3.621)
P_0	0.142	0.998	0.986	0.996	1.000	0.999	0.999	0.999	0.999	0.845
P_1	0.649	0.895	0.992	0.836	0.999	1.000	0.939	0.985	0.999	0.023
HL_0	6.293	6.510	11.432	7.479	13.013	11.140	5.180	N/A	6.505	15.512
HL_1	2.529	N/A	10.481	N/A	1.581	4.526	0.681	0.696	0.716	2.438
$\alpha_0 = \alpha_1$	1.513	68.526	1.088	3.453	1.178	0.986	17.579	1.027	1.645	10.396
$\lambda_0 = \lambda_1$	21.861	50.348	0.420	9.144	5.086	7.461	48.469	28.520	6.369	20.561
$\sigma_0 = \sigma_1$	19.277	5.704	13.980	12.506	43.799	5.315	4.012	2.597	9.985	10.111

Notes: Cyprus (CP), Czech Republic (CZ), Estonia (EO), Hungary (HN), Latvia (LV), Lithuania (LN), Malta (MA), Poland (PO), Slovakia (SX) and Slovenia (SJ). These results are for the estimation of the MS-ADF model given in equations (2) and (3). Figures in brackets are t-statistics based on Newey-West standard errors. * denotes significance of the autoregressive parameter at the 5% significance level or better based on a critical value of -1.65. HL_0 and HL_1 denote the half lives (months) of Regime 0 and Regime 1 respectively. The tests for $\alpha_0 = \alpha_1$, $\lambda_0 = \lambda_1$ and $\sigma_0 = \sigma_1$ are each distributed as $\chi^2(1)$ on the null with a 5% critical value of 3.84.

An examination of the regime-switching autoregressive parameters λ_0 and λ_1 indicates that the non-stationary null can be rejected in both regimes for thirteen out of eighteen differentials.[3] For the five remaining cases involving the Czech Republic-US, Malta-US, Czech Republic-Germany, Hungary-Germany and Poland-Germany differentials, either λ_0 or λ_1 is statistically insignificant. While these pairs of countries are engaged in regime-switching activity, only one regime is itself characterized by the presence of long-run RIP. We therefore argue that partial RIP is applicable here.

At the 5% significance level, the null $\lambda_0 = \lambda_1$ is rejected in fourteen out of eighteen cases. For most cases where there is evidence in favor of long-run RIP, the low-volatility Regime 0 is associated with a negative and significant autoregressive coefficient (λ_0). Moreover, in the cases of the Czech Republic and Malta differentials with respect to the US, and the Hungary differential with respect to Germany, regime-switching of the real interest rate differential towards a low-volatility regime also involved switching from a regime of non-stationary behavior to a regime of stationary behavior. Figure 1 offers a visual inspection of the inferred probability of being in the low-volatility Regime 0. Subject to a few spikes, when λ_0 is statistically significant there is a clear switch towards Regime 0 during the study period. Generally speaking, the inferred probabilities show a noticeable move towards regimes of low-volatility during the mid-1990s. Before 1996, many of these series were in a high-volatility regime, reflecting turbulence with unstable fundamentals associated with the transition from a command economy to a market economy. However, in the case of the Poland-Germany differential it appears that Regime 1 is the stationary regime with only λ_1 being negative and significant. In this case, there is no such clear switching point towards stationarity and RIP during the study period.

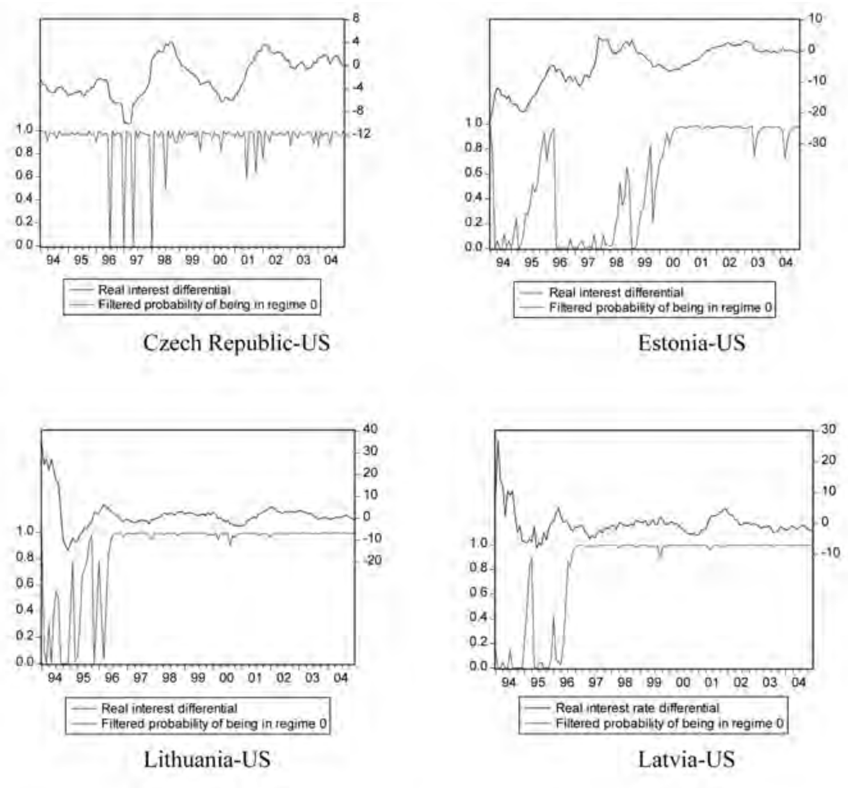
Most differentials are characterized by the presence of two significant autoregressive coefficients thereby suggesting varied RIP is present. These are the cases of the Estonia, Latvia, Lithuania, Poland, Slovenia and Slovakia differentials with respect to US; and the Cyprus, Estonia, Latvia, Lithuania, Malta, Slovenia and Slovakia differentials with respect to Germany. Varied RIP signifies stronger evidence of RIP than partial RIP. In these cases, real interest differentials always exhibit stationary behavior, but the differentials are subject to regime shifting with different speeds of adjustment towards RIP. In terms of differentials from Germany, the relevance of varied RIP to the cases of Estonia, Latvia and Lithuania is consistent with their status of high liberalization of capital movements. These countries have a longer track record of removed exchange controls where a currency board or fixed peg regime has been maintained. However, varied RIP is also present in cases where capital mobility

3. Based on a 5% critical value (no trend) of -1.65 (see van Norden and Vigfusson (1998)).

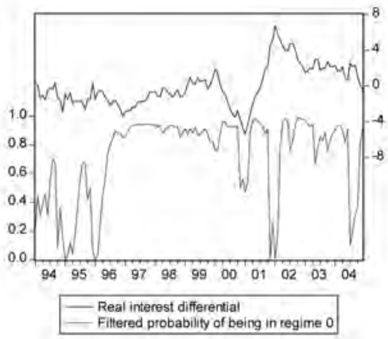
is less fluid. In this sense, it would appear that limited capital mobility has not been sufficient to impede long-run RIP.[4] Regime 0 is generally characterized as the regime characterized by a relatively slower speed of adjustment. Figure 1 reveals that most differentials have a tendency to having a high probability of being in Regime 0 beyond the early stages of the study period, but with short periods of shifting for short periods towards Regime 1. Again, there is considerable volatility in the transition probabilities in the period up to the mid 1990s.

Figure 1. Real interest rate differential and the filtered probability of being in Regime 0

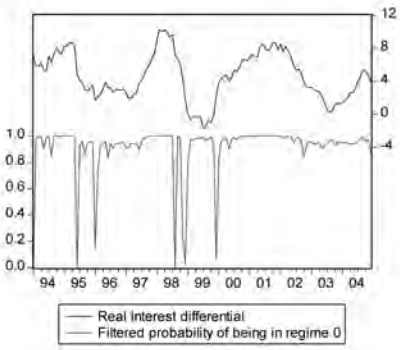
A. Vis-à-vis US



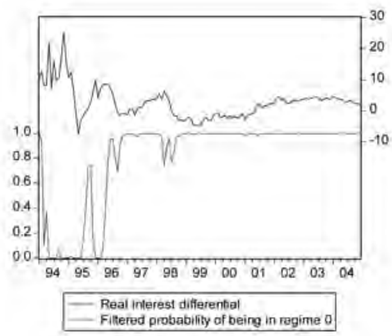
4. Cole and Obstfeld (1989) argue that under certain conditions, real interest rates can be equal and perfectly correlated even in the presence of complete barriers to capital mobility.



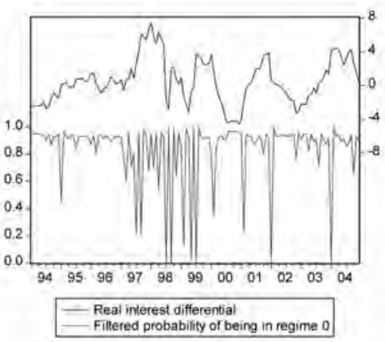
Malta-US



Poland-US

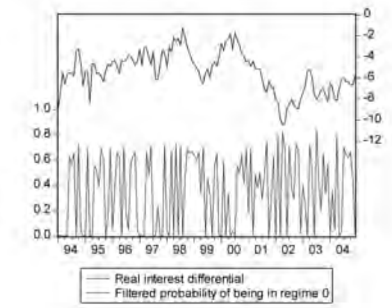


Slovenia-US

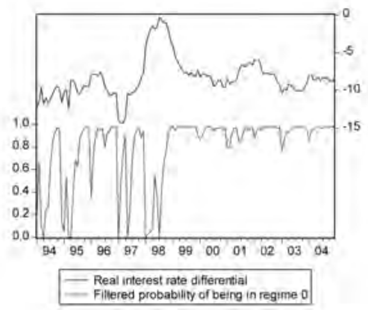


Slovakia-US

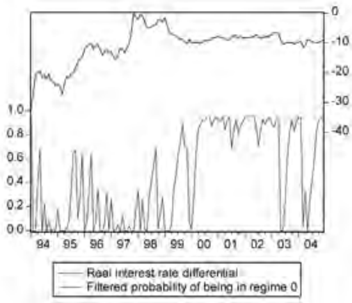
B. Vis-à-vis Germany



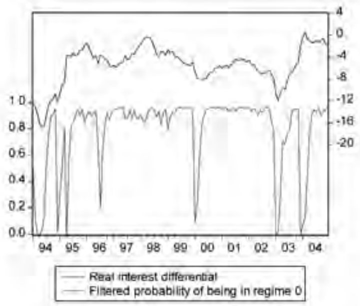
Cyprus-Germany



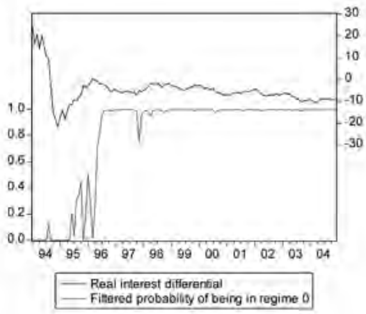
Czech Republic-Germany

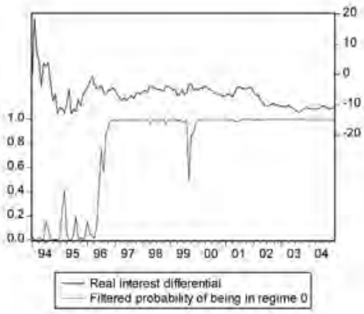


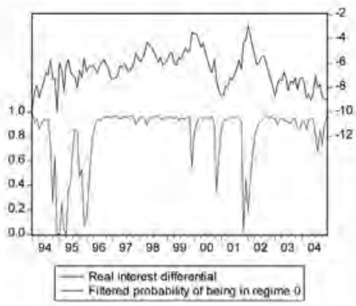
Estonia-Germany



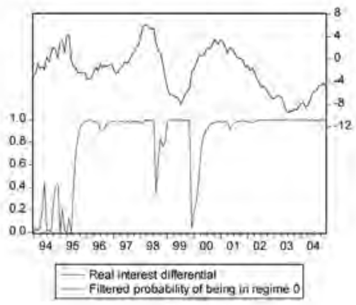
Hungary-Germany



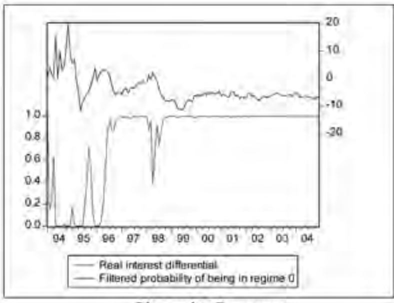




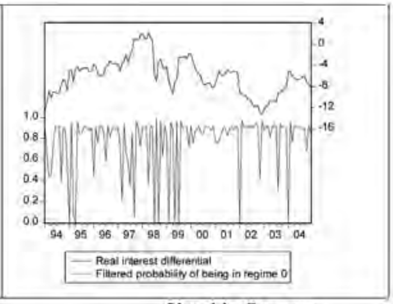
Malta-Germany



Poland-Germany



Slovenia-Germany



Slovakia-Germany

Note: In each case, the real interest rate differential (right axis) is plotted above the filtered probability (left axis).

The findings reported and discussed here might be compared with recent studies of interest rate convergence involving the accession countries. For example, Kutan and Yigit [(2004), (2005)] analyze both nominal and real interest rate differentials. They consider the post-1993 period and adopt a number of panel linear estimation approaches. Their results indicate limited degrees of nominal and real economic convergence. They find that the Baltic states exhibit the strongest monetary policy and price-level convergence, suggesting that they are ready to adopt the Euro. However, the Central and East European countries are characterized by a lack of convergence. Brada *et al*. (2005) use rolling cointegration to measure the convergence of base money, M2, the consumer price index (CPI) and industrial output between Germany and France and recent EU members and a sample of transition countries that are now joining the EU. They find that cointegration for the transition economies was comparable for M2 and prices, but not for monetary policy and industrial output. Holtemoller (2005) analyzes UIP with the Euro area of potential EMU accession countries. Using a methodology based on backward recursive statistical tests, error correction models and rolling regressions applied to study the co-movement of interest rates, Hungary, Poland and Slovenia are found to exhibit the least degree of convergence. With regard to Hungary and Poland, this finding is again consistent with our study, though in this case we find Slovenia is more converged than is suggested in the literature. This finding of stronger real convergence may be seen against Slovenia's acceptance into the single currency area at the beginning of 2007.

Finally, for the instances where at least one regime of stationary behavior is identified, we may consider the degree of persistence or the speeds of adjustment back towards long-run RIP following a random disturbance. The half-life of a random disturbance can be approximated using methods discussed in the previous section. We find that the approximated half lives range from 23.5 and 17.5 months in the cases of the respective Czech-US and Estonia-US differentials to 0.7 months in the cases of the Malta-Germany and Slovenia-Germany differentials. Where RIP holds, these values are considerably less than studies of relative PPP among OECD countries (see, for example, Wu (1996)) that find half lives in the region of 30 months to be appropriate. In the context of the accession members, Cihak and Holub (2005) analyze price convergence in new EU countries and find that it may take about 10-25 years for new EU countries to converge to that of the least developed, more-established EU members. However, it should be remembered that the speed of RIP adjustment considered in our study comprises both adjustment to UIP and PPP. Given the differences in trading in financial assets and goods and services, one might expect the individual speed of adjustment towards UIP to exceed that associated with PPP.

5. Conclusion

The extent to which real interest rates are involved in a long-run equilibrium relationship is a useful guideline for assessing the degree of real convergence. In this study, we have conducted a formal analysis of real interest parity involving ten EU accession countries along with the US and Germany. However, in contrast to the existing literature, this paper has brought into focus the possibility that evidence on long-run real interest parity is susceptible to variations in the persistence of regimes over time. It is important to account for states of the world where the economies fluctuate between different regimes of real interest rate relationships. It is possible indeed that the dynamics of real interest differentials are more accurately described by regimes where local adjustment towards RIP does or does not occur. The failure to account for such nonlinearities, in terms of switches between regimes of stationary and non-stationary behavior may increase the likelihood of evidence against long-run parity by introducing a bias towards longer half-lives of deviations from real interest parity. Based on Markov-regime switching ADF tests of stationarity, this study demonstrates that it is possible to identify real interest differentials across the EU accession countries that embody regimes of local stationary and non-stationary behavior.

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