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Testing for Real Interest Rate Convergence in European Countries

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4 Empirical tests and results

4.1 Data

We use both short-term and long-term interest rates for eight European countries: Belgium, Denmark, France, Germany, Ireland, Italy, Netherlands and UK. Our measure of short-term rates is the 3-month Eurocurrency rate. Our measure of the long-term rate is the government bond yield. The inflation rates are constructed using the Consumer Price Index. The data are quarterly and cover the 1979–1993 period. The beginning of the period is chosen to coincide with the launch of the EMS and our set of countries includes all ERM members for this period. We have also included the UK as it is highly likely that it will participate in a future monetary union. Data on Eurocurrency rates are taken from the WEFA Group Database. Eurocurrency rates on Ireland were not available. All other data are taken from the International Financial Statistics of the IMF. We constructed both ex post and ex ante real interest rates. The ex post real interest rate series were created using the Fisher equation as follows:

$$r_t = i_t - (P_{t+4} - P_t)/P_t$$

where r_t is the real interest rate at time t earned from holding the investment for four quarters, i_t is the nominal interest rate and P_t is the price index. $(P_{t+4} - P_t)/P_t$ is the inflation rate from time t to $t+4$. In constructing the ex ante rate we created an expected inflation series using a 4-period moving average of actual inflation rates.

4.2 Unit root tests

Since a necessary condition for performing cointegration tests is that individual time series are non stationary, we first perform unit root tests. Table 1 lists the τ_μ and τ_τ statistics for the real short-term interest rates. As the results for ex ante rates were similar to those for ex post rates, we only report those for ex post rates. The order of the ADF test, k, is set at four following the approach suggested by Schwert (1987), i.e., $k = \text{int}[4(n/100)^{.25}]$. All real short-term rates are I(1) except for those for Belgium, Denmark and Italy. Therefore, these three countries are excluded from our cointegration tests for short-term real rates. Table 2 reports similar statistics for real long-term interest rates. Even though Table 2(b) indicates that the first difference of Italian rates is not stationary, we assume this to be the case given the low power of unit root tests. We, therefore, conclude that all real long-term rates are difference stationary, i.e., I(1).

Our finding that most real interest rate series are I(1) deserves some discussion. Assuming that nominal interest rates and inflation rates are I(1), the Fisher equation would imply that the real interest rate is I(0) or that cointegration exists between nominal interest rates and inflation. If, however, inflation is I(2), as much evidence has suggested, then, the finding of I(1) real interest rates could be explained[6]. A large part of the literature has concluded that real interest rates follow a random walk (e.g., Walsh, 1987, Rose, 1988). Garcia and Perron (1996) find a shift in the mean of US real interest rates that could account for the earlier findings of a unit root. The authors using a three-state Markov switching model for the inflation rate find a regime shift in the mean of the series. This mean shift in the inflation process could be what our tests are picking up when we test for cointegration of real interest rate pairs allowing for regime shifts. Malliaropulos (1996) confirms the Garcia and Perron finding of U.S. real interest rate stationarity but finds that this stationarity obtains around a negative linear trend that shifted upwards in the third quarter of 1980.

4.3 Engle-Granger cointegration tests

Tables 3(a) and 3(b) report the ADF(4) statistics on the residuals of the cointegration regression for ex post short-term and long-term real interest rates, respectively. Again, $k = 4$ following Schwert (1987). The null hypothesis is no bilateral cointegration between Germany and the rest of the countries. As the results for short-term and long-term rates are identical for both ex ante and ex post rates, we only report results for the latter[7]. In all cases, the null of no cointegration cannot be rejected. It is possible that this result is due to a structural break in the cointegration vector which could be seen at the plot of residuals of the Engle and Granger cointegration regression. As an example, Figures 1 and 2 show these residuals for France (short-term rates) and Ireland (long-term rates). The two figures show a shift in the series taking place around the date of the structural break. By incorporating an intercept dummy, this change in the interest rate gap is captured by the deterministic component of the model and does not enter

[6] A finding of I(1) real interest rates has important implications for some asset-pricing models. For example, the Consumption CAPM (CCAPM) predicts similar time series properties for the growth rate of consumption and real interest rates, a theoretical implication not validated by U.S. data due to the nonstationarity of real interest rates (Rose, 1988).

[7] Other studies (e.g., Cumby and Mishkin, 1986, Goodwin and Grennes, 1994) found similar results using both ex ante and ex post real interest rates.

the residual. Hence, to examine this possibility, we need to apply cointegration tests that allow for a structural break in the cointegration relation.

4.4 Cointegration tests with structural breaks

Table 4 reports the values of Gregory-Hansen (1996) statistics for the three models for ex post short-term real rates. The results using ex ante rates, being similar to those for ex post rates, are not reported. These results imply no evidence for cointegration between German and Dutch rates and strong evidence (at 1%) for cointegration between German rates and rates in France and the UK. The break point is reported in parentheses as a percentage of the sample size. The break points are in 1981:3 and 1982:1 for France and the UK, respectively. The break date for France corresponds to high French real interest rates that can be attributed to expansionary fiscal policy followed by the newly elected French socialist government in 1981.

Table 5 reports the Gregory and Hansen (1996) cointegration test results for long-term real interest rates. The null of a lack of cointegration is not rejected for Belgium, Italy and Netherlands[8] but is rejected (at 5% or less) for Denmark, France, Ireland and the UK. When ex ante rates are used, similar results are obtained (and hence are not reported) except in the case of Belgium where the null hypothesis is rejected. The break dates are as follows: 1983:1 for Belgium, 1983:4 for Denmark, 1982:1 (or 1985:3, or 1987:1 depending on the chosen model) for France, 1981:4 for Ireland and 1989:1 (or 1990:3, or 1991:1) for UK. The break date for Belgium takes place during the period of an ERM realignment where the Belgian franc depreciated by 4% against the DM. The break date for Denmark corresponds to a period of high real interest rates that can be attributed to the expansionary fiscal contraction taking place in the country following the initiation of a fiscal tightening programme in 1983 (Bertola and Drazen, 1993). 1982:1 represents a period of high French real interest rates accompanying the expansionary fiscal policy launched in the second half of 1981. 1987:1 concides with a 3% depreciation of the French franc against the DM in an ERM realignment and the start of a period of low German long-term real interest rates (and hence a rise in the gap between French and German real interest rates) due to lax

[8]The lack of a finding of cointegration for Netherlands seems surprising given that Dutch monetary policy is known to follow German monetary policy very closely. A look at the residual plot of the Engle-Granger cointegration test result for ex post long-term rates reveals the existence of a large spike in 1979:3 that cannot be picked up by the Gregory and Hansen tests as it happens too close to the beginning of our sample. Using the 1980:1–1993:4 period and the Engle-Granger cointegration test we reject the no cointegration null at 5%. A similar finding applies for ex ante long-term rates as the ADF(4) statistic for the 1980:3-1993:4 period is significant at 5%.

German monetary policy (Smaghi and Micossi, 1990). 1981:4 represents a period of sharply rising Irish real interest rates as fiscal policy turned very expansionary in the early 1980s. Also, during the fourth quarter of 1981, the Irish pound depreciated by 5.5% against the DM in an ERM realignment. Finally, 1990:3 is associated with a period of increasing British real interest rates as the UK applied contractionary monetary policy in preparation for joining the ERM.

4.5 Cointegration regressions

Tables 6 and 7 include the estimated cointegration regressions for real short-term and long-term rates, respectively, allowing for a structural break as determined in the previous subsection using the Gregory and Hansen tests. We use the dynamic OLS estimator (DOLS) suggested by Stock and Watson (1993) that provides more efficient estimates than alternative procedures (e.g., West, 1988). According to their approach, we regress one of the variables onto contemporaneous levels of the remaining variables, leads and lags of their first differences, and a constant, using ordinary least squares. The covariance matrix is estimated by averaging the first four autocovariances with the Bartlett kernel. As the results obtained for different models where cointegration exists in each country were qualitatively similar, we only report results for model 1. The strong form of RIP with structural change would imply that $a_1 = 0, b = 1$ and $a_2 \neq 0$. Table 6 shows that this result applies for France. However, cointegration with structural change does not hold for the UK since $a_2 = 0$. Table 7 shows that for all countries $a_2 \neq 0$ indicating a change in the long-run cointegration vector. In addition, b is not statistically different from one (except in the case of France) implying a perfect link between real long-term rates in each of these countries and Germany. This provides evidence for the weak but not the strong form of RIP as $a_1 \neq 0$.

In summary, our study, allowing for structural changes in the cointegrating vector, has obtained two major results: first, we find strong evidence for real short-term interest rate convergence between France and Germany. Equivalently, the strong form of RIP holds for France against Germany. Second, we have also provided significant evidence for convergence between long-term real rates in Germany and most ERM-member countries. This finding signifies that increasing convergence of German and other European real long-term rates has taken place gradually over the last 15 years and implies gradual loss of the effectiveness of stabilization policy by these European countries. This is because the effectiveness of monetary policy in affecting real long-term domestic interest rates, and hence private investment and output, is reduced.

5 Conclusions

We have tested for the weak and strong forms of the RIP in European countries assuming the German Dominance Hypothesis holds. Statistical evidence in favour of the RIP would have important implications for policymaking as it would be consistent with real interest rate convergence between German and other European rates and imply loss of the effectiveness of long-run stabilization policy[9] by these countries. We find that the results differ depending on the type of tests used: traditional cointegration tests do not support the hypothesis of real interest rate convergence whereas recently-developed tests that determine endogenously potential structural breaks imply that real interest rate convergence has taken place in several European countries, particularly for long-term real rates. In addition, the significant evidence we have provided in favour of real long-term interest rate convergence could possibly reflect the fact that the markets anticipate additional convergence of real short-term interest rates in these countries vis-a-vis Germany as we are approaching the launch date of the European monetary union.

Our results have important implications for the effectiveness of domestic stabilisation policies. In particular, for several countries where real long-term interest rate convergence applies, domestic monetary policy would be expected to have lost some of its effectiveness as a long-run stabilisation policy tool. In addition, as RIP represents a building block in some monetary models of exchange rate determination, the evidence we have supplied in favour of the weak form of RIP would hint to the need for the consideration of regime shifts when testing for the validity of these models. It is possible that the inability of these models to explain exchange rate behaviour, as the majority of the empirical evidence to date has suggested (Taylor, 1995), is due to a shift in the long-run relationship between the exchange rate and its determinants.

[9] As cointegration is a long-run concept, a finding of cointegration would mean that there is bilateral real interest rate convergence in the long run and therefore, stabilisation policy can still be effective in the short run.

Table 1: Dickey-Fuller tests (ex post short-term real rates)

(a) Level data

	ADF(4)	
	τ_μ	τ_τ
Belgium	-5.92*	-6.04*
Denmark	-3.84*	-4.29*
France	-2.90	-2.64
Germany	-2.83	-2.78
Ireland	-	-
Italy	-4.46*	-4.43*
Netherlands	-2.98*	-3.02
U.K.	-2.29	-1.90

(b) Differenced data

	ADF(4)	
	τ_μ	τ_τ
France	-4.05*	-4.21*
Germany	-3.53*	-3.51*
Netherlands	-3.55*	-3.60*
U.K.	-4.24*	-4.52*

Note: A * implies significance at 5%. The critical values, given by MICROFIT, for the no trend and trend models are -2.91 and -3.49, respectively.

Table 2: Dickey-Fuller tests (ex post long-term real rates)

(a) Level data

	ADF(4)	
	τ_μ	τ_τ
Belgium	-2.13	-1.82
Denmark	-2.38	-2.71
France	-3.19*	-2.18
Germany	-1.89	-2.11
Ireland	-2.38	-1.58
Italy	-3.33*	-2.79
Netherlands	-2.13	-2.06
U.K.	-2.48	-2.39

(b) Differenced data

	ADF(4)	
	τ_μ	τ_τ
Belgium	-4.01*	-4.17*
Denmark	-4.00*	-4.23*
France	-2.71	-3.56*
Germany	-4.11*	-4.46*
Ireland	-3.95*	-4.34*
Italy	-2.21	-2.69
Netherlands	-3.67*	-4.12*
U.K.	-2.94*	-3.06

Note: * implies significance at 5%. The critical values, given by MICROFIT, for the no trend and trend models are -2.91 and -3.49 respectively.

Table 3: Engle-Granger Cointegration Tests

(a) Ex post short-term real interest rates

	ADF(4)
France	-2.43
Netherlands	-2.63
U.K.	-2.81

(b) Ex post long-term real interest rates

	ADF(4)
Belgium	-2.25
Denmark	-2.36
France	-2.04
Ireland	-1.55
Italy	-2.06
Netherlands	-2.66
U.K.	-2.02

Note: The critical value (5% level) is -3.45 (see Table 1 in MacKinnon (1991)).

Table 4: Gregory-Hansen Cointegration tests (ex post short-term rates)

	ADF*	Z_t^*	Z_α^*
France			
Model (1)	-5.06**(0.18)	-7.27***(0.18)	-56.68***(0.18)
Model (2)	-5.59**(0.18)	-7.82***(0.18)	-60.94***(0.18)
Model (3)	-5.31**(0.18)	-7.42***(0.18)	-57.89***(0.18)
Netherlands			
Model (1)	-3.62(0.37)	-3.84(0.38)	-26.93(0.38)
Model (2)	-4.21(0.43)	-4.42(0.42)	-31.95(0.42)
Model (3)	-3.61(0.37)	-3.85(0.38)	-27.00(0.38)
UK			
Model (1)	-6.57***(0.22)	-4.13(0.20)	-28.29(0.20)
Model (2)	-6.63***(0.22)	-4.19(0.20)	-28.75(0.20)
Model (3)	-6.44***(0.22)	-4.12(0.20)	-28.81(0.20)

Note: *, **, and *** denote significance at 10%, 5% and 1%, respectively. The numbers in parentheses are the break points reported as a percentage of the sample size.

Table 5: Gregory-Hansen Cointegration tests (ex post long-term rates)

	ADF*	Z_t^*	Z_α^*
Belgium			
Model (1)	-4.30(0.30)	-3.25(0.35)	-19.50(0.35)
Model (2)	-4.32(0.30)	-3.33(0.35)	-20.00(0.35)
Model (3)	-4.35(0.30)	-3.41(0.35)	-20.66(0.35)
Belgium (ex ante)			
Model (1)	-4.39*(0.28)	-2.99(0.30)	-16.94(0.30)
Model (2)	-4.37(0.28)	-2.99(0.30)	-16.94(0.30)
Model (3)	-4.56(0.28)	-3.19(0.30)	-18.59(0.30)
Denmark			
Model (1)	-4.46*(0.33)	-3.11(0.32)	-17.02(0.32)
Model (2)	-5.15**(0.33)	-3.45(0.32)	-21.96(0.32)
Model (3)	-5.01**(0.33)	-3.59(0.32)	-23.59(0.32)
France			
Model (1)	-5.10**(0.55)	-4.11(0.55)	-26.52(0.55)
Model (2)	-5.14**(0.22)	-3.82(0.20)	-24.18(0.53)
Model (3)	-6.75***(0.45)	-4.70(0.45)	-35.27(0.45)
Ireland			
Model (1)	-4.56*(0.22)	-5.02**(0.20)	-35.55(0.20)
Model (2)	-4.45(0.22)	-5.00**(0.20)	-35.39(0.20)
Model (3)	-4.41(0.22)	-5.02**(0.20)	-34.87(0.20)
Italy			
Model (1)	-3.67(0.83)	-2.64(0.78)	-21.38(0.78)
Model (2)	-3.71(0.37)	-3.62(0.38)	-23.12(0.37)
Model (3)	-3.82(0.63)	-3.66(0.62)	-22.51(0.62)

Abstract

We use cointegration tests that determine endogenously the regime shift to test for bilateral short-term and long-term real interest rate convergence in the European Monetary System in the 1979–1993 period. The results of these tests provide strong evidence in favour of bilateral real interest rate convergence between Germany and several countries in our sample, particularly for long-term real interest rates. This result carries the important policy implication that in several European countries monetary policy has lost some of its effectiveness as a stabilisation policy tool.

Keywords: Real interest parity, cointegration with regime shifts

JEL Classification: F3, F4

Netherlands			
Model (1)	-3.87(0.52)	-3.52(0.50)	-22.42(0.45)
Model (2)	-3.89(0.43)	-3.57(0.45)	-22.77(0.45)
Model (3)	-4.10(0.27)	-3.67(0.25)	-24.48(0.25)
UK			
Model (1)	-5.19***(0.82)	-3.59(0.78)	-21.03(0.78)
Model (2)	-5.30**(0.78)	-3.91(0.78)	-23.65(0.78)
Model (3)	-5.36**(0.68)	-3.45(0.80)	-20.26(0.80)

Note: *, **, and *** denote significance at 10%, 5% and 1%, respectively. The numbers in parentheses are the break points reported as a percentage of the sample size.

Table 6: DOLS regressions (ex post short-term real rates)

	a_1	a_2	b
France (0.18)	-0.028	0.045	1.333
	(-1.58)	(3.84*)	(0.73)
UK (0.22)	-0.009	0.011	1.453
	(-0.73)	(1.59)	(1.11)

Note: The numbers in parentheses under the columns of a_1, a_2 are t-statistics for the null that the corresponding coefficients are zero. The number in parentheses under the column of b is the t-statistic for the null that b is equal to one. The t-statistics follow the Student's t distribution asymptotically. * indicates significance at 5%.

Table 7: DOLS regressions (ex post long-term real rates)

	a_1	a_2	b
Belgium (0.28)	0.012	0.011	0.879
(ex ante)	(2.30*)	(3.57*)	(-0.53)
Denmark (0.33)	0.050	-0.020	0.892
	(4.31*)	(-3.18*)	(0.21)
France (0.55)	-0.050	0.032	1.955
	(-6.28*)	(9.27*)	(3.32*)
Ireland (0.22)	-0.053	0.077	0.952
	(-4.55*)	(7.63*)	(-0.10)
UK (0.82)	-0.026	0.040	1.476
	(-2.05*)	(5.17*)	(0.85)

Note: The numbers in parentheses under the columns of a_1, a_2 are t-statistics for the null hypothesis that the corresponding coefficients are zero. The number in parentheses under the column of b is the t-statistic for the null that b is equal to one. The t-statistics follow the Student's t distribution asymptotically. * indicates significance at 5%.

IGURE 1

ENGLE-GRANGER REGRESSION RESIDUALS
(FRANCE, SHORT-TERM INTEREST RATES)

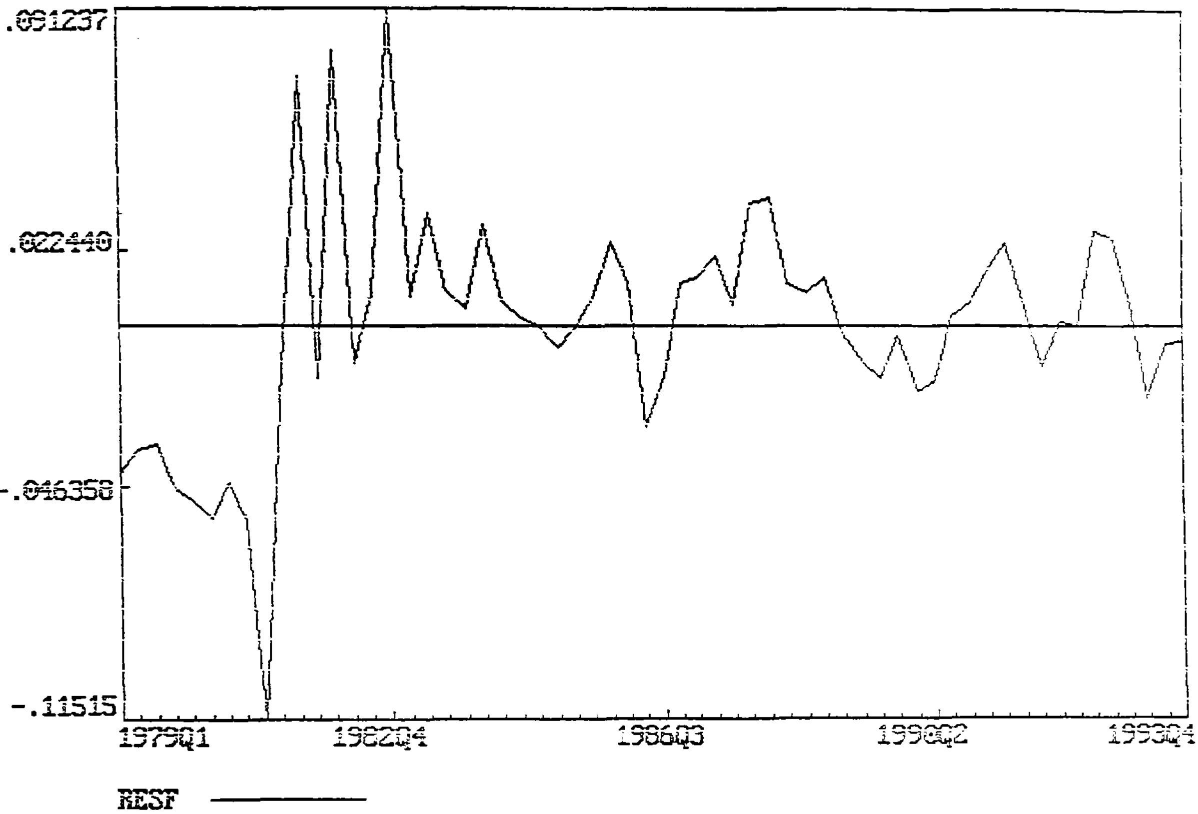
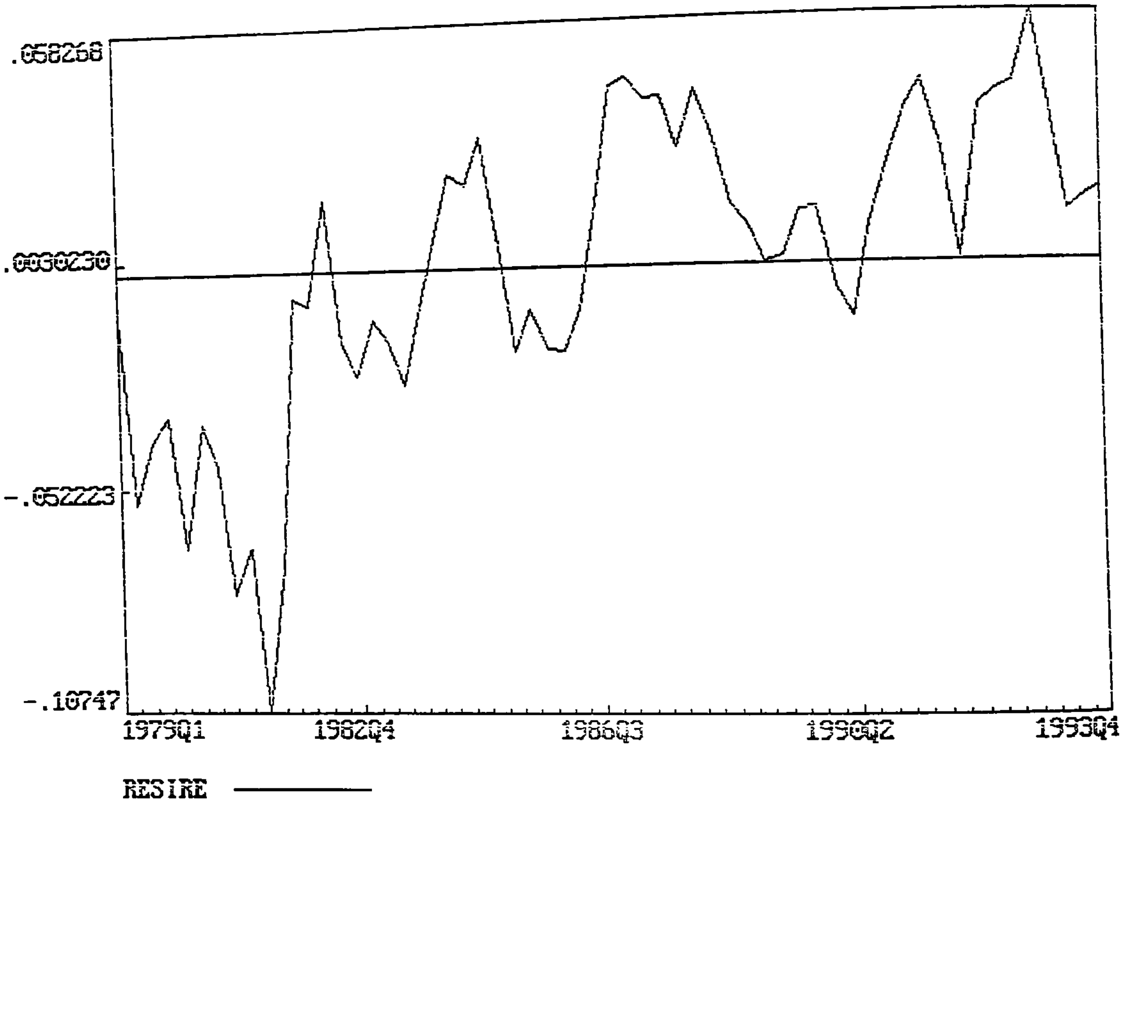


FIGURE 2

ENGLE-GRANGER REGRESSION RESIDUALS
(IRELAND, LONG-TERM INTEREST RATES)



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1 Introduction

One of the most noticeable changes in the world financial markets since the 1970s has been the growing degree of integration as constraints to the movement of financial capital have been gradually relaxed. In particular, for most of the member countries of the Exchange Rate Mechanism (ERM), the abolition of capital controls progressed during the 1980s and was completed by the 1st July of 1990. In an environment of fixed exchange rates, the increasing degree of international integration of financial markets would lead to a tendency for equalization of real interest rates across national boundaries (real interest rate parity). Under a system of flexible exchange rates (or a system of quasi-fixed exchange rates like the European Monetary System), real interest rate convergence may not obtain because of expectations about exchange rate changes and foreign exchange risk premia. The advent of the flexible exchange rate regime in 1973 and the relaxation of capital controls in some major industrial countries have opposite effects on the degree of real interest rate convergence in these countries. However, for countries that belong to the ERM, the relaxation of capital controls in the 1980s along with the lower variability of nominal and real exchange rates, as the member countries coordinated their monetary policies, should be expected to lead to increasing real interest rate convergence.

A finding of real interest rate convergence has important policy implications for the effectiveness of domestic stabilization policy. With real interest rates set in international markets, domestic monetary policy would have no influence on savings/consumption decisions (Feldstein, 1980). This applies, in particular, for small open economies or countries whose monetary policy is dependent on that of the dominant country of a monetary union. A finding of real interest rate convergence between Germany and other ERM countries, assuming the German Dominance Hypothesis (GDH) holds[1], would imply that these countries would have limited success in stabilizing their domestic economies since they would face difficulties in influencing their real interest rates.

An empirical analysis of real interest rate convergence or real interest rate parity[2] (RIP) is also useful because RIP represents a building block in ex-

[1] If Germany is the dominant country in the ERM, it should determine both interest rates and inflation rates, and hence real interest rates, in the system. This implies independent German monetary policy and the borrowing of anti-inflation reputation by other countries in the ERM. For evidence in favour of the GDH see Giavazzi and Giovannini (1989), Herz and Roger (1992), Karfakis and Moschos (1990) and Thom (1994).

[2] In this paper, the terms real interest rate convergence and real interest rate parity are used interchangeably. In the next section, we distinguish between weak and strong forms of the real interest parity.

change rate models. Early versions of the monetary model of exchange rate determination introduced by Frenkel (1976) and Mussa (1976), among others, assume both Purchasing Power Parity (PPP) and Uncovered Interest Parity (UIP) and hence RIP in the long run. The real interest rate differential model introduced by Frankel (1979) allows for sticky prices in the short run and implies that RIP holds in the long run when the exchange rate reaches its long-run equilibrium value. Moreover, real interest rate differentials are an important determinant of short-run exchange rates.

All previously published empirical attempts to test for real interest rate convergence have either employed classical regression analysis or conventional cointegration tests that do not consider regime shifts. We depart from this literature by making use of recently-developed cointegration techniques that allow for structural shifts in the cointegration vector. The rationale for using these tests is our intention to capture the changing degree of capital mobility, convergence of some macroeconomic fundamentals like inflation, and credibility of exchange rate policy in the ERM countries during the 1980s. Use of the conventional cointegration tests and, therefore, failure to allow for changes in regime might lead to what appears to be a conventional result based on the empirical evidence of the last fifteen years, i.e., rejection of the real interest rate parity condition. In contrast, as this paper shows, in several cases, the hypothesis of real interest rate convergence cannot be rejected if we allow for the possibility of structural changes in the cointegration vector.

The paper is organized as follows: section 2 provides a short overview of the concepts of weak and strong forms of real interest parity and the empirical literature to date, section 3 discusses our methodology and section 4 our results. Finally, section 5 summarizes our conclusions and draws some policy implications.

2 Background and Literature

Measuring perfect capital mobility in international financial markets has been one of the most extensively researched topics in the area of international finance. Alternative definitions of perfect capital mobility differ in terms of the assumptions imposed[3]. In this paper we focus on the RIP for the reasons mentioned in the introduction. RIP requires that the following three conditions hold: (a) UIP, (b) PPP and (c) the ex ante Fisher equation in

[3]Frankel (1992) classifies the four well known definitions in ascending order of specificity: (i) Feldstein-Horioka tests, (ii) RIP, (iii) UIP and (iv) Covered interest parity (CIP).

both the domestic and foreign country (Hallwood and MacDonald, 1994, p. 45).

Alternatively, one could determine what factors explain deviations from RIP. Using simple algebra and the ex post version of the Fisher equations for the domestic and foreign countries we derive:

$$r_t - r_t^* = (i - i^* - \Delta s) - (p - p^* - \Delta s) \tag{1}$$

where r, i, s, and p stand for the real interest rate, nominal interest rate, nominal exchange rate, and inflation rate, respectively, Δ is the difference operator and an asterisk denotes foreign variables. The first three terms of the right-hand side represent the deviation from the UIP and the last three terms the deviation from PPP. The deviation from UIP is due to the country premium (e.g., capital controls, differential tax systems, political risk) and the currency premium (i.e., exchange risk premium). The developments in the ERM in the 1980s would be expected to lead to a decline in these premia. For example, the increasing dismantlement of capital controls and the boost in credibility of national exchange rate policies in the 1980s would be a contributing factor to the reduction of the country and currency premium, respectively. In addition, the increasing convergence in inflation rates along with the smaller nominal exchange rate changes should lead to decreasing deviations from PPP. In summary, the ERM developments of the 1980s should be expected to contribute to declining deviations from UIP and PPP and hence, according to equation (1) above, increasing real interest rate convergence.

Emerson et al. (1992, p. 160) use the results obtained by Frankel (1991) to derive estimates of the above-mentioned determinants of the deviation from real interest rate parity with Germany being the centre country. Using three-month money market interest rates and actual values as proxies for inflation and exchange rate expectations for the period September 1982–April 1988, the authors find that the smallest deviation from RIP applies for Netherlands, Belgium, France and UK from the countries included in our sample.

Two forms of RIP can be defined. Consider the following regression:

$$r_t = a + br_t^* + e_t \tag{2}$$

where r_t and r_t^* are the foreign and domestic (German) ex post real interest rates respectively, a and b are parameters and e_t is an error term. Provided that r_t and r_t^* have single unit roots, the following forms of the RIP can be considered:

1. The strong form holds if e_t is stationary (i.e., r_t and r_t^* are cointegrated) and $a = 0, b = 1$ or equivalently, if the real interest rate differential $r_t - r_t^*$ is stationary.

2. The weak form holds if e_t is stationary and $a \neq 0$ and/or $b \neq 1$.

a and b may differ from the values implied by strong RIP, even though financial markets are fully integrated, for several reasons:

(a) The presence of transaction costs that creates a neutral band with no profitable arbitrage around real interest parity.

(b) The existence of non-traded goods whose prices cannot be equalized internationally (in a common currency) causing price indexes and real interest rate differences across countries in the presence of fully integrated financial markets.

(c) A constant foreign exchange risk premium.

(d) Differential national tax rates.

To test for the strong form of RIP we can test for stationarity of the real interest rate differential. Equivalently, one can test for cointegration between the two real interest rates and once cointegration is established (i.e., the error term e_t in equation (2) is stationary) to test the joint null hypothesis $a = 0$ and $b = 1$.

However, even if the strong form of RIP does not hold, if domestic and foreign real interest rates are cointegrated (i.e., they do not tend to drift apart over time), real interest rate convergence would exist but would not be perfect. In such a case, policymakers would still have some, but not full, control over their domestic stabilization policies.

Previous empirical research on RIP can be divided into two groups[4]: early studies tested for the strong form of RIP (i.e., real interest rate equality) using classical regression analysis (i.e., ran simple ordinary least squares in equation (2)) and hence not allowing for potential nonstationarity of the real interest rate series. Examples include Mishkin (1984a, 1984b), Cumby and Mishkin (1986), Mark (1985), Gaab, Granziol and Horner (1986) and

[4]The study by Fraser and Taylor (1990) does not belong to any of these two groups. The authors test for RIP by testing whether the nominal interest rate differential is an optimal predictor of the relative future inflation rates subject to the maintained hypothesis of rational expectations. Using data on seven OECD countries, the authors reject RIP for all 18 country pairs investigated.

Cumby and Obstfeld (1984). In the majority of cases examined, these studies found evidence against the parity. These studies are subject to three criticisms: first, they cannot account for transaction/information costs that lead to deviations from perfect real interest rate equalization. Second, the conventional statistical tests employed in these studies are inappropriate if the individual real interest rate series are non stationary as the OLS estimators are not consistent and the standard t and F statistics do not follow the student's t and F distributions. Third, even if the non-stationary real interest rate series are pairwise cointegrated, classical statistical inference is invalid since the estimated standard errors are inconsistent (Stock, 1987).

More recent studies test for the strong and weak forms of RIP using cointegration techniques in both a bivariate and multivariate framework. Some evidence is provided for the weak RIP but no evidence for the strong RIP. Throop (1994) tests for bilateral RIP between the US and each of Japan, UK, and a foreign trade-weighted real interest rate using the Johansen approach. Some evidence for the weak form of RIP is provided. However, there is no evidence for a long-run, one-to-one association between US real interest rates and foreign real rates, i.e., strong RIP is rejected by the data. Goodwin and Grennes (1994) perform bilateral and multilateral tests using the US as a base country. Using both Eurocurrency and domestic money market interest rates for the period 1975–1987, the authors find that bilateral cointegration applies for the US against Canada, UK and Germany. This evidence supports the weak form of RIP.

3 Econometric Methodology

Conventional cointegration tests suffer from a major drawback when the time period under study includes changes in the *modus operandi* of the monetary system, fiscal policy changes, institutional changes, political upheavals, etc. These tests do not consider the possibility that what appears to be nonstationarity of a linear combination of variables (i.e., lack of cointegration), is in fact a deterministic break in the mean or trend of a linear combination of these variables (i.e., a shift in the cointegration vector over the sample period). In other words, the presence of breaks in the time series biases tests for the null of no cointegration in favour of acceptance.

We think it is appropriate to consider whether real interest rate convergence has actually taken place in the European Monetary System (EMS) by allowing for possible shifts in the long-run cointegration relationship. A break in the long-run (cointegration) relationship between pairs of real interest rates can happen for the following reasons: First, our sample period

1979–1993 includes a time span of significant dismantlement of restrictions on the free movement of capital controls across national boundaries in the EMS. Second, the pre-September 1992 EMS period can be divided into three subperiods (Gros and Thygesen, 1992) that correspond to different degrees of credibility of exchange rate policy and convergence of monetary policy and inflation rates among the member countries. Third, drastic changes in the stance of fiscal and monetary policy that sometimes are associated with a change in the political regime can also account for a change in the relationship between pairs of real interest rates. This section discusses our econometric methodology with emphasis placed on the Gregory and Hansen (1996) tests for cointegration that allow for the endogenous determination of the structural break in the cointegration vector.

We first test for cointegration between pairs of real interest rates with Germany being the reference country, as discussed earlier, using the Engle-Granger methodology. Engle and Granger (1987) suggest a two-step procedure where simple regressions are run for pairs of real interest rates (i.e., equation (2) above) and tests for the null of a unit root (lack of cointegration) in the estimated residuals are performed. Several test statistics for the above null hypothesis have been proposed. We have decided to use the Augmented Dickey Fuller (ADF) statistic for two reasons: first, Engle and Granger (1987) recommended that this test has the largest power. Second, the Gregory-Hansen (1996) tests employed in this study are a direct extension of the ADF test. More recently, Johansen (1988) and Johansen and Juselius (1990) have suggested maximum likelihood cointegration tests. However, as the Gregory and Hansen (1996) tests are residual-based cointegration tests like the Engle-Granger tests, we have chosen to use as a basis of comparison the Engle-Granger approach instead of the Johansen approach[5].

Gregory and Hansen (1996) develop cointegration tests under regime shifts where the timing of the regime shift is not known *a priori* but needs to be determined endogenously by appealing to the data. Gregory and Hansen (1996) consider three models of an endogenous one-time regime shift that reflect three different alternative hypotheses:

$$\begin{array}{llll} \text{Model 1:} & r_t = a_1 + a_2 D_t + b r_t^* + u_t & t = 1, \ldots, n & (3) \\ \text{Model 2:} & r_t = a_1 + a_2 D_t + b r_t^* + ct + u_t & t = 1, \ldots, n & (4) \\ \text{Model 3:} & r_t = a_1 + a_2 D_t + b_1 r_t^* + b_2 r_t^* D_t + u_t & t = 1, \ldots, n & (5) \end{array}$$

$$\text{where} \quad D_t = \begin{array}{ll} 0, \text{if} & t \leq [n\tau] \\ 1, \text{if} & t > [n\tau] \end{array}$$

and $\tau \in (0, 1)$ is an unknown parameter denoting the relative timing of the

[5]We also ran the Johansen cointegration tests for the purpose of comparison. However, it turned out that the results are very sensitive to the lag choice.

change point and [ ] denotes integer part. The use of the dummy variable D_t allows one to test for a structural change or regime shift. In model 1, there is a level shift in the cointegrating relationship which is modeled as a change in the intercept by the size of coefficient a_2. In model 2 a linear trend is added to model 1. Finally, model 3 extends model 1 in that it allows the structural change to affect both the intercept and the slope. b_1 represents the cointegration slope coefficient before the regime shift and b_2 the change in the slope coefficient following the regime shift. It is obvious that model 1 is nested within model 3. The null hypothesis in all three models is that u_t is nonstationary or, in other words, r_t and r_t^* are not cointegrated. Cointegration with structural change implies that u_t is an I(0) process and that a_2 (and b_2) are significantly different from zero.

To test for cointegration between r_t and r_t^* with structural change, i.e., stationarity of u_t in models 1 through 3, Gregory and Hansen (1996) suggest the use of three tests. These tests are modifications of the test statistics Z_α and Z_t (suggested by Phillips (1987)) and the ADF statistic. These statistics are defined as:

$$\begin{aligned} Z_\alpha^* &= \inf_{\tau \in T} Z_\alpha(\tau) \\ Z_t^* &= \inf_{\tau \in T} Z_t(\tau) \\ \mathrm{ADF}^* &= \inf_{\tau \in T} \mathrm{ADF}(\tau) \end{aligned}$$

where $Z_\alpha(\tau), Z_t(\tau)$ and $\mathrm{ADF}(\tau)$ correspond to the choice of change point τ. The set T can be any compact subset of $(0,1)$. Gregory and Hansen (1996) suggest that a reasonable choice is $T = (0.15, 0.85)$. Following Gregory and Hansen we compute the test statistic for each break point in the interval $([0.15n], [0.85n])$. According to the definition of Z_α^*, Z_t^* and ADF*, we are interested in the smallest values of $Z_\alpha(\tau), Z_t(\tau)$ and $\mathrm{ADF}(\tau)$ across all possible break points since small values of the statistics are required to reject the null hypothesis. Gregory and Hansen (1996) derive asymptotic critical values for alternative models. Their table 1 lists the critical values for our case. Based on Monte Carlo evidence for the model with structural break in the intercept and the slope, they also find that Z_t^* has the largest power and Z_α^* the lowest power (see Table 3 in Gregory and Hansen (1996)).