Testing real interest parity in the European Monetary System

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and

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Abstract

Current discussions on Economic and Monetary Union (EMU) in Europe have stressed the need for enhanced integration of goods and factor markets as a pre-condition of moving as costlessly as possible to a single currency system. The real interest differential—and hence tests of real interest parity—provide a summary measure of the degree of residual integration necessary such that these conditions are met. Empirical tests suggest a rejection of real interest parity among European Monetary System (EMS) member countries, at least during recent years. Further, a decomposition of the differential suggests that imperfect integration of goods markets, as reflected in a failure of *ex-ante* PPP, is largely responsible for this finding.

Testing Real Interest Parity in the European Monetary System

It is widely documented that full European Monetary System (EMS) members have continued to support at times large *nominal* interest rate differentials, despite the progress that has been made towards integration of foreign exchange markets since the advent of the Exchange Rate Mechanism (ERM). Much less well documented, however, is the extent of differentials in *real* rates of return across the ERM bloc. This omission is all the more surprising given the importance which attaches to the real interest rate as a monetary transmission mechanism in most textbook macroeconomic models.

More importantly in the context of current discussions on Economic and Monetary Union (EMU), the question of whether countries should be subject to budgetary rules which govern the size of their fiscal deficits(1) depends on the extent to which real interest rates are interdependent; or more generally on the extent to which different countries' debt issues are substitutable. If markets are perfectly integrated, then one country's fiscal deficit, to the extent that it affects domestic real interest rates, will affect Community-wide real interest rates. This then has the potential to impose output externalities, both between countries in the Union who are at different stages in the cycle, and for the Union as a whole *vis-à-vis* non-Union members.

By definition, real return differentials can have one of two sources: imperfect financial (capital and foreign exchange) market integration, as reflected in the *nominal* differential; and/or imperfect structural (goods market) integration, as reflected in price level divergences. Current EMU discussions have stressed the need for enhanced integration and convergence of goods and factor markets as a pre-condition of moving as costlessly as possible to a single currency system. The real interest rate differential provides an indication of the residual integration that is necessary between ERM countries such that these conditions are met. Correspondingly, decomposing the real interest differential identifies the markets in which this additional convergence must come if two countries are to converge on a steady-state of a common real rate of return. It is for these reasons—and given the on-going debate regarding the future of monetary arrangements in Europe and the fiscal implications these may have—that a study of real interest parity is relevant in an ERM context.

The paper is broadly planned as follows. Section I describes our real interest rate measures, derived from survey data on inflation expectations. Section II provides some formal empirical tests of the real interest rate parity proposition and of potential interdependencies between real interest rates across the ERM countries, in an attempt to provide a lead into the ERM and EMU issues raised above. Section III considers

(1) As has been recommended in the Report on Economic and Monetary Union in the European Community (1989).

the source of these real interest differentials, whether in financial or goods markets. Section IV briefly summarises and concludes.

I Real interest rates: some empirical evidence

In this section we aim to generate a number of real interest rate measures (conditioned on different expectational assumptions), compare them qualitatively over time and then draw some preliminary conclusions regarding their temporal pattern. This in turn helps motivate our later discussion.

For each country we define the *ex-ante* real interest rate as:(1)

$$E_t({}_t r_{t+j}) = {}_t i_{t+j} - E_t({}_t \pi_{t+j}) \tag{1}$$

where: ${}_t r_{t+j}$ = real return on a j-period bond held between time t and $t + j$,
${}_t i_{t+j}$ = nominal return on the above j-period bond,
${}_t \pi_{t+j}$ = inflation rate between t and $t + j$,
E_t = expectations operator

From (1) above, it is clear that any real interest rate measure will be conditioned by our expectational priors. Below we attempt to determine how sensitive our real interest rate projections are to these expectational assumptions by comparing two measures of real interest rates, one *ex-post*, the other *ex-ante*. We define the *ex-post* measure as:

$${}_t epr_{t+j} = {}_t i_{t+j} - {}_t \pi_{t+j}$$

That is, actual replaces expected inflation in (1). Under Rational Expectations (RE):

$${}_t epr_{t+j} = E_t({}_t r_{t+j}) + {}_t \varepsilon_{t+j} \tag{2}$$

where $E_t({}_t \varepsilon_{t+j} / I_t) = 0$

I_t = information set at time t

(1) The strict form of the Fisher hypothesis would include an additional term, $\pi_t i_t$, which we assume here to be second order. This does not appear an unreasonable assumption, since the countries studied are not hyperinflationary. In addition, it is worth noting that real interest rate measures should ideally be defined net of tax, ie

$$E_t({}_t r_{t+j}) = {}_t i_{t+j}(1 - E_t(\tau_{t+j})) - E_t({}_t \pi_{t+j})$$

where τ_{t+j} is the expected marginal rate of taxation at time $t + j$. Mark (1985) finds that real interest rates are relatively insensitive to the inclusion of tax rate terms, possibly reflecting the difficulties involved in constructing an appropriate expected marginal tax rate measure.

Since the inflation forecast error, ${}_t\varepsilon_{t+j}$, is unforecastable given I_t, *ex-post* and *ex-ante* real interest rates should coincide *on average*. While early studies made explicit use of *ex-post* real interest rate measures as a convienient means of obviating the unobserved expectations problem (see, for example, Cumby and Obstfeld (1981), Mishkin (1981)), here we have chosen to use *ex-post* real rates largely as our benchmark. The reasons for doing this are that the implicit restrictions which are imposed when using *ex-post* measures seem likely to be overly restrictive. Specifically, *ex-post* measures are equivalent to requiring that agents have perfect foresight; that is, the strictest (Muthian) form of (deterministic) RE in which no error in forecasting inflation is made in any period.

The dominant approach in the academic literature when generating *ex-ante* real interest rate measures has been to use a RE Instrumental Variables (IV) methodology (see, for example, Mark (1985), Cumby and Mishkin (1986), Frankel and MacArthur (1988)). In conditioning expectations upon an instrument subset of I_t, rather than the complete information set, such *ex-ante* IV measures are less restrictive in the information requirements they ascribe to agents than are *ex-post* measures. IV techniques are themselves not without problems, however. First, the information we obtain from *ex-ante* IV measures hinges critically upon our choice of an appropriate instrument subset. There is no academic consensus about what are the appropriate instruments to include within such a subset when proxying real interest rate behaviour (see Barro and Martin (1990) for some recently suggested alternatives). Second, the information provided by *ex-ante* measures is largely qualitative in nature, ie the expectations series remains unobservable. Third, the IV methodology imposes exact unbiasedness upon the expectations data.

One way around the above problems of IV techniques when arriving at an *ex-ante* real interest rate measure is to derive estimates of price expectations directly—thus dispensing with the problem of defining explicitly the information set of agents (the first problem above). This is the approach adopted in our study.[1] In generating our expectations series, we draw upon survey data of industrialists' selling price expectations published by the European Commission. The use of survey data appears to be attracting ever-greater academic approbation as an alternative to IV techniques for capturing otherwise unobservable movements in expectations series.[2] Our preference for using selling price expectations reflects the fact that most overseas investment is undertaken by larger firms, the aggregate consumption bundle for which is more likely to be intermediate or wholesale goods (to go towards the production of further output) than goods for final consumption. Suitably transformed, the qualitative responses from such survey data can be put into an explicitly quantitative

(1) See also, among others, Peek and Wilcox (1983) and Wilcox (1983), who study US real rates using Livingston's price expectations data.

(2) See, for example, the recent study by Frankel and Froot (1990).

form (thus obviating the second problem above). Specifically, we employ the method of Carlson and Parkin (1975) to transform our survey data into an inflation expectations series (see Appendix 1 for details of the method of transformation). While the Carlson and Parkin method, by construction, pushes our *ex-ante* measures in the direction of unbiasedness, it does not impose this restriction exactly, as is the case with the alternative methods outlined.[1] As such, our derived expectations series imposes a less severe unbiasedness restriction than is the case with the *ex-post* and *ex-ante* IV measures. It is for these three reasons that we believe our expectations data is to be preferred to that derived using alternative methodologies. As such, in what follows our *ex-ante* survey measure is used as the basis for our reported results; this being compared with a benchmark *ex-post* measure when testing real interest parity as a means of determining the robustness of our results. (We do not report results for an *ex-ante* IV measure, for the reasons outlined above.)

Our *ex-ante* real interest rate measure is then:

$$_t ear_{t+j} = {}_t i_{t+j} - {}_t\pi_{t+j}^{S}$$

where $_t\pi_{t+j}^{S}$ = expected inflation series derived from survey data

From (1) above, it is clear that, in addition to our assumptions regarding expectations formation, our real interest rate measures will depend additionally upon our choice of nominal interest rate (long or short rates; money market or eurocurrency rates) and our choice of price deflator (consumer or wholesale price indices). As regards nominal interest rates, we use short-term three month, end-month domestic money market rates (or their equivalent). This enables us to pick up the effect of capital controls between countries in a way potentially overlooked if eurocurrency rates are used. Consumer prices are taken as our deflator.[2] The data is monthly between

(1) Indeed, conducting unbiasedness tests on our survey data (using the techniques proposed in Brown and Maital (1981), Pesaran and Wright (1991) and Wallis (1989)) illustrated that we were able to reject unbiasedness of the data for more than half of the countries studied. The conclusions we draw from this are either that our expectations data is deficient, or that the informational assumptions implicit within *ex-post* inflation measures are overly restrictive. These two hypotheses are unfortunately difficult to decouple from a single time series. This is analogous to the familiar problem of distinguishing between the joint hypotheses of rational expectations and the robustness of the underlying structural model (see, for example, Begg (1982)). Encouragingly, however, correlations between ex-post and ex-ante inflation measures were, in the main, relatively consistent across countries. This is indicative of the rationality restrictions implicit within the ex-post real interest rate measures being overly burdensome, rather than the data being deficient. This in turn helps justify further our using *ex-ante* real interest rates as the basis for this study.

(2) For all countries (with the exception of the United Kingdom) consumer price measures were found to be more closely related to the survey data than were wholesale price measures. For the United Kingdom, a wholesale prices index was found to be more in line with our survey responses for inflation expectations and was used throughout. Using the UK retail price index excluding mortgage interest payments was found to make little difference to the ex-ante real interest rate measures generated; and it is *ex-ante* real rates which are our principal concern in this study. Monthly Irish consumer price data were arrived at by (non-linear) interpolation of the appropriate quarterly series.

January 1976 and July 1990, thus covering both pre- and post-EMS periods, and is expressed in annualised form. The data for nominal interest rates and prices are taken from *International Financial Statistics.* Bilateral exchange rate data, both spot and forward, are taken from *Financial Statistics.* Survey price expectations data, in balance form (ie the balance between those reporting price rises and those reporting price falls), is taken from *Supplement B* of *European Economy* and from European Commission sources. The countries covered are: West Germany, France, Italy, Belgium, Ireland and the United Kingdom (which acts as the ERM counterfactual over our sample).

Figures 1–6 show the *ex-post* and *ex-ante* real interest rate measures generated. Without exception, the pattern is of higher real interest rates during the 1980s than was the case during the 1970s, which is well documented in the literature. Moreover, even by 1980s standards real interest rates are high at the end of the sample, ranging from around 6% in Germany and Italy to around 9% in the United Kingdom and Ireland. The timing and degree of movement in real interest rates is markedly different across the countries studied. For example, real interest rates are positive in Germany and Belgium almost throughout the sample, while for France, the United Kingdom and Ireland this is not consistently the case until after 1979, and for Italy not until after 1980. This suggests some move towards more stringent monetary policies by non-German ERM members after the system's inception, though, as is evident from the UK experience, this shift towards a more restrictive stance was not confined exclusively to the ERM bloc.[1]

Figures 7–11 plot real interest rate differentials relative to Germany, the 'centre' EMS currency. For some ERM members, specifically Italy, Ireland and France, it is clear that, while adjustment towards the German policy standard started soon after 1979, this adjustment only became effective some time later as real interest rates in non-German ERM countries moved above those in Germany: for France this appears to have occurred after 1983; while for Italy, Ireland and Belgium the adjustment occurs somewhat later. Note in particular the extent to which real interest rates in non-German countries were forced to 'overshoot' German levels in order to bear down on inflationary pressures in these high inflation countries. After 1987 there are indications of convergence of ERM real interest rates on German levels, with French, Belgian and Italian real interest rates all little different from their German counterparts by 1989. Convergence of real interest rates is much less evident over this period for the non-ERM country, the United Kingdom.

(1) Note also that a shift in monetary policy will only affect real interest rates in the short run, and even then only when prices adjust sluggishly. In steady-state, real interest rates are invariant to a countries' monetary policy stance: see Shiller (1980).

Ex-post and ex-ante real interest rates

- - - - Ex-post
——— Ex-ante

Figure 1: Germany

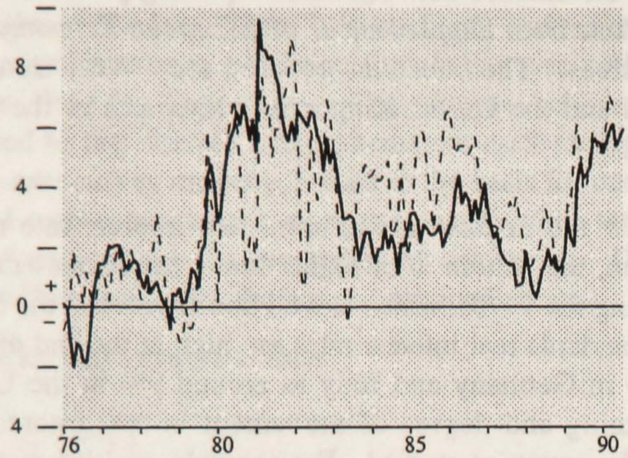


Figure 2: France

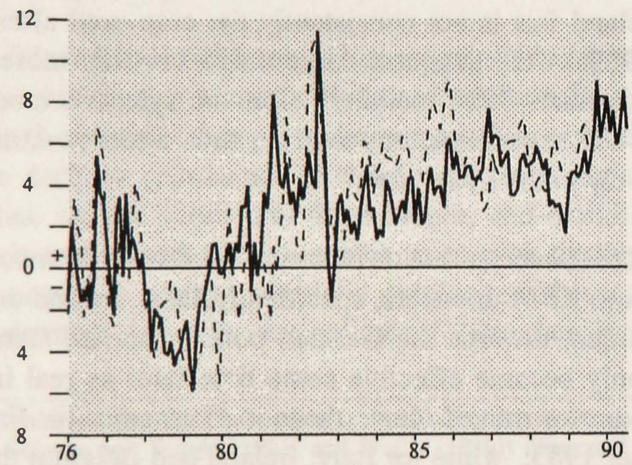


Figure 3: United Kingdom

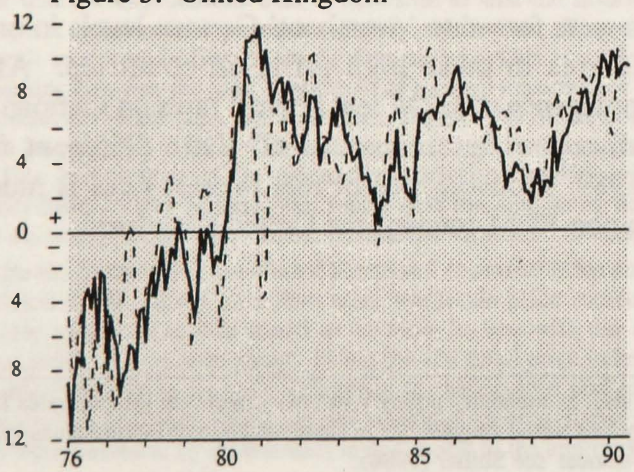


Figure 4: Belgium

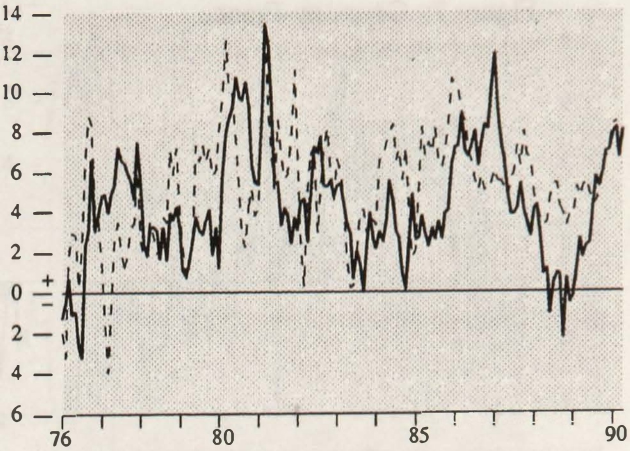


Figure 5: Italy

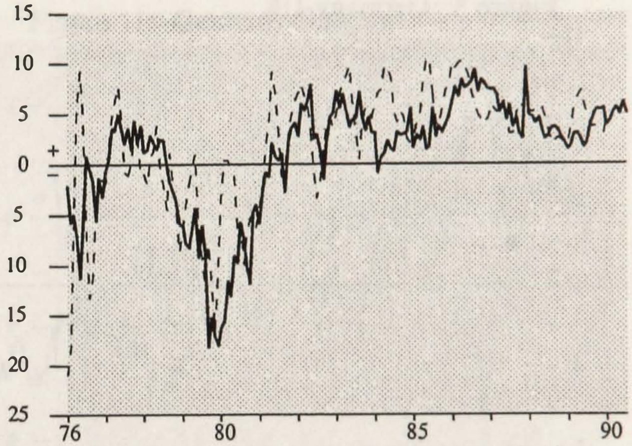
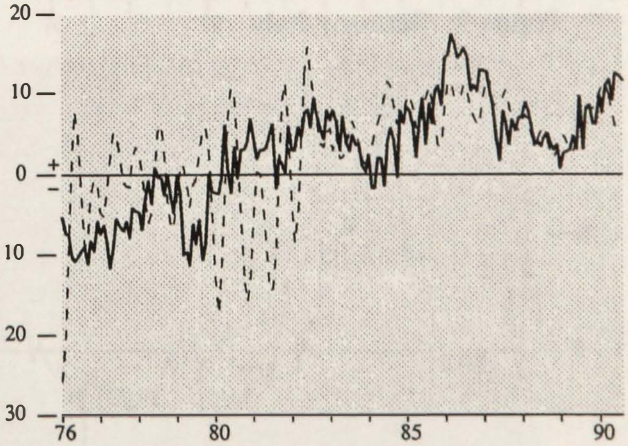


Figure 6: Ireland



Ex-ante real interest rate differentials

Figure 7: Germany-France

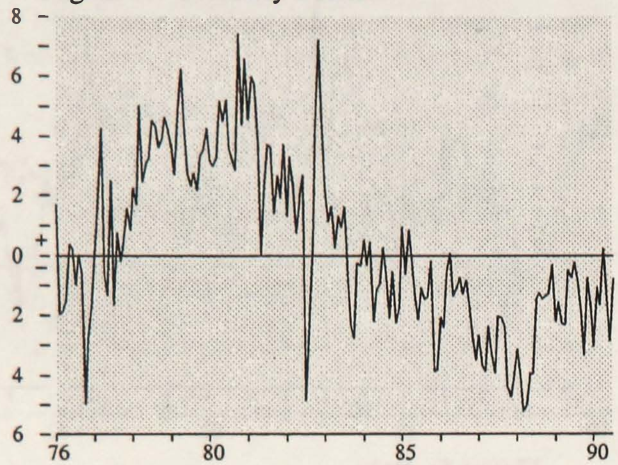


Figure 8: Germany-UK

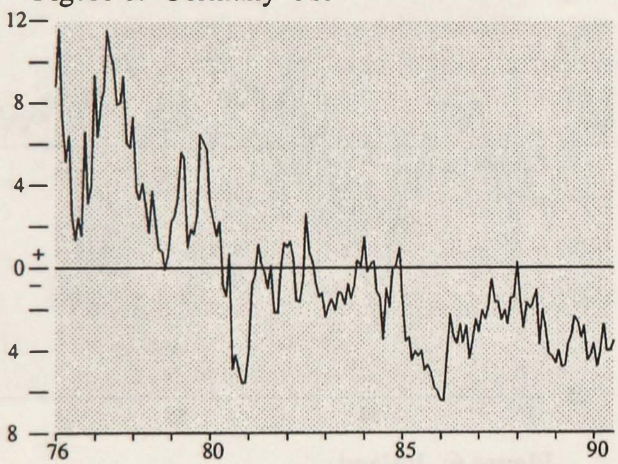


Figure 9: Germany-Italy

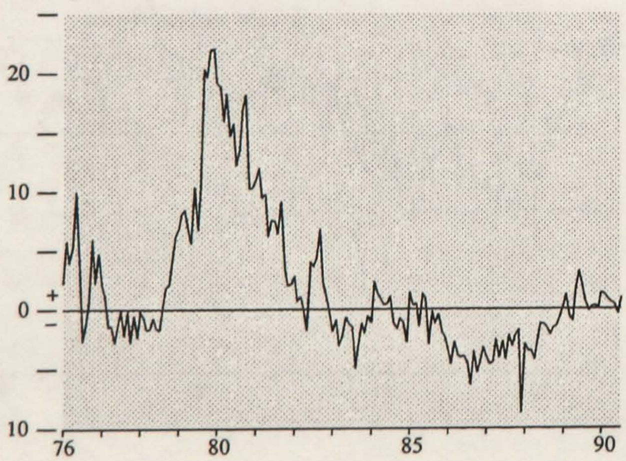


Figure 10: Germany-Ireland

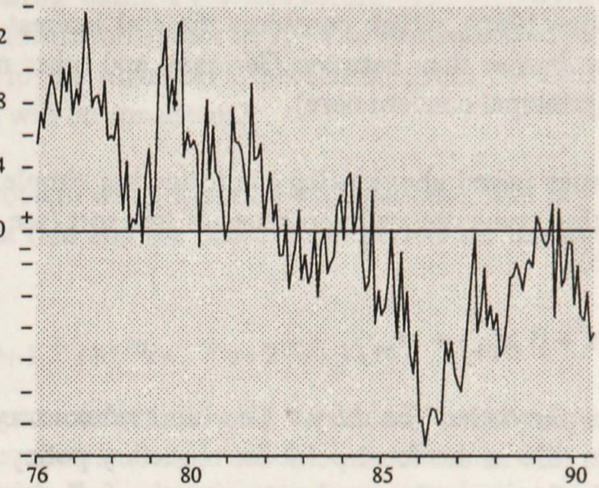
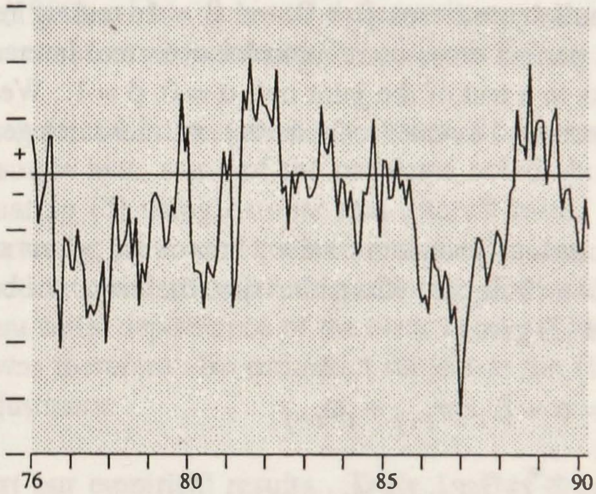


Figure 11: Germany-Belgium



II Testing the real interest rate parity proposition

In this section we formally test the extent to which real returns have become equal over time across the ERM and, on a related point, the degree of covariation between real interest rate measures. As such, this section resembles closely the study by Cumby and Mishkin (1986), which considers the real interest rate linkage between the US and Europe (rather than between Germany and other members of the ERM bloc, which is our principal concern here).

We address the issues raised above using the following simple framework, which is standard when testing convergence hypotheses of this sort (see, for example, Cumby and Mishkin (1986)):

$$ {}_t r_{t+j}^{*} = \alpha + \beta \, {}_t r_{t+j} + {}_t e_{t+j} \tag{3} $$

where a * denotes the 'other' country. The 'centre' country here is taken to be Germany, given its role as the focal point for monetary policy operations within the ERM. Using this framework we aim to test the following hypotheses. The independence, or otherwise, of real interest rates is tested using the β coefficient; specifically the null hypotheses $\beta = 0$ and $\beta = 1$, testing respectively complete independence and perfect cross-correlation between real interest rates. Real interest rate parity amounts to a test of the joint null $\alpha = 0$, $\beta = 1$. We test these hypotheses using both the *ex-post* and *ex-ante* real interest rate measures generated earlier.

A number of technical problems arise, however, when estimating (3) using conventional OLS techniques. Consider first the *ex-post* observable form of (3). Combining (2) and (3) gives:

$$ {}_t epr_{t+j}^{*} = \alpha + \beta \, {}_t epr_{t+j} + {}_t \omega_{t+j} \tag{4} $$

where $\quad {}_t \omega_{t+j} = {}_t \varepsilon_{t+j}^{*} - \beta \, {}_t \varepsilon_{t+j} + {}_t e_{t+j}$

Three econometric problems, familiar when estimating in a rational expectations framework, are evident from (4). First, the assumption typically made is that the structural error term in (3), ${}_t e_{t+j}$, is white noise and hence that this term by itself will not generate serial correlation problems in (4). Since our model is non-structural this assumption is unlikely to be satisfied, with omitted structural variables potentially complicating the error process ${}_t e_{t+j}$, causing inconsistencies in our estimates. We return to this problem again below.[1]

(1) If *a priori* we had knowledge of the order of serial correlation in the structural error, one solution would have been to transform our reduced-form equation (3) to ensure serial independence of the structural error.

Second, orthogonality between the composite error term ${}_t\omega_{t+j}$ and regressor ${}_tepr_{t+j}$ is clearly not satisfied in (4), with the realised *ex-post* real interest rate correlated with the inflation forecast error. OLS estimation of (4) will therefore yield inconsistent (downward biased) coefficient estimates. Consistent coefficient estimates can be obtained by estimating using IV (McCallum (1976)), with an instrument set conditioned on an information subset of I_t. The (instrumented) regressors in (4) will then be orthogonal to the inflation expectation errors by construction and hence coefficient estimates will be consistent.(1)

Finally, there is a problem of multiperiod expectations; that is $j = 3$, so the holding period on the interest rate and the observation period are unequal. Specifically, we have from (4):

$${}_t\omega_{t+3} = {}_t\omega_{t+1} + {}_{t+1}\omega_{t+2} + {}_{t+2}\omega_{t+3}$$

with the composite error term following an MA(2) process by construction. As a result, the covariance matrix in (4) will be inconsistent, with coefficient standard errors biased downwards, hence causing too frequent rejection of the null hypotheses. To overcome this problem, we use the generalised method of moments adjustment of Hansen and Hodrick (1980), which 'corrects' the covariance matrix in the presence of a moving average error to ensure consistency.(2)

To the extent that *ex-ante* real interest rates are observed with measurement error or that there is small sample bias, some of the problems noted above will be equally apparent when estimating (3) using *ex-ante* real interest rates. Additionally, and probably more importantly, however, the non-structural nature of (3) is likely to generate by construction serial correlation problems. Estimation confirmed the existence of significant serial dependence of the error in an *ex-ante* form of (3). The *ex-ante* form of (3) was therefore also estimated subject to the Hansen and Hodrick covariance matrix adjustment.

Tables 1 and 2 report our empirical results. Table 1 gives the test statistics from 2SLS estimation of equation (3) using *ex-post* interest rate measures and allowing for the Hansen and Hodrick covariance matrix adjustment.(3) The instruments used for German *ex-post* real interest rates were a constant, a linear time trend, the nominal German interest rate and (lagged) actual inflation. The *ex-ante* regression is reported

(1) Consistency also requires, however, that the instruments chosen be uncorrelated with ${}_te_{t+j}$, which may be a more doubtful assumption given the problems noted above.

(2) An alternative approach would have been to use the two-step 2SLS technique of Cumby, Huizinga and Obstfeld (1983).

(3) In practice, the order of serial correlation when estimating (4) was found to be greater than two, and an MA(4) correction was applied. This may point towards some correlation between the instrument set and the structural error ${}_te_{t+j}$. Some residual seasonality was also detectable at higher-order lag lengths.

Table 1: Ex-post regressions; 2SLS with adjusted covariance matrix

$$_t epr_{t+j}^* = \alpha + \beta \ _t epr_{t+j} + {}_t\omega_{t+j}$$

	Full Sample	Sub-samples (1) 76m1–79m3	(2) 79m4–83m3	(3) 83m4–87m10	(4) 87m11–90m6
United Kingdom					
α	-2.86	-15.62	1.32	4.29	2.69
	(1.58)	(32.24)	(3.29)	(3.49)	(1.66)
β	1.84^{**}	15.56	0.60	0.35	1.08
	(0.43)	(42.68)	(0.73)	(0.88)	(0.42)
$\alpha = 0; \beta = 1$	3.85	1.20	0.44	8.14^{**}	17.30^{*}
France					
α	1.35	0.31	-1.30	8.22	3.30
	(1.40)	(7.02)	(3.41)	(3.15)	(0.54)
β	0.47	-0.76	0.54	-0.76^{**}	0.67^{**}
	(0.38)	(9.30)	(0.76)	(0.80)	(0.14)
$\alpha = 0; \beta = 1$	2.25	1.11	8.13^{**}	10.40^{*}	102.23^{*}
Belgium					
α	2.23	3.42	3.38	2.55	3.14
	(0.68)	(5.10)	(1.67)	(2.52)	(0.79)
β	0.91	-0.89	0.69	0.84	0.60^{**}
	(0.19)	(6.77)	(0.37)	(0.64)	(0.20)
$\alpha = 0; \beta = 1$	40.13^{*}	10.35^{*}	14.90^{*}	15.45^{*}	30.75^{*}
Italy					
α	-0.15	2.87	-2.10	3.17	4.68
	(1.99)	(21.29)	(5.57)	(3.04)	(1.31)
β	0.69	-6.22	0.32	0.74	-0.13^{*}
	(0.55)	(28.25)	(1.24)	(0.77)	(0.35)
$\alpha = 0; \beta = 1$	1.85	0.89	6.88^{**}	23.45^{*}	12.70^{*}
Ireland					
α	1.12	7.64	6.13	2.82	2.76
	(2.27)	(32.35	(6.99)	(4.69)	(1.71)
β	0.54	-12.82	-1.48	1.17	0.89
	(0.62)	(42.92)	(1.53)	(1.18)	(0.43)
$\alpha = 0; \beta = 1$	0.68	0.51	7.97^{**}	19.56^{*}	10.55^{*}

Standard errors are in parentheses. A * (**) alongside the coefficient β indicates rejection of the null $\beta = 1$ at 1% (5%). A * (**) alongside the $\alpha = 0; \beta = 1$ test statistic indicates rejection of the null at 1% (5%). The test statistics for the β coefficient are distributed as a 't' and are two-sided, with critical values 2.58 at 1% and 1.96 at 5%. The Wald test statistic for the joint hypothesis $\alpha = 0; \beta = 1$ is distributed as a $x^2(2)$, with critical values 9.21 at 1% and 5.99 at 5%.

Table 2: Ex-ante regressions; adjusted covariance matrix

$$_{t}ear_{t+j}^{*} = \alpha + \beta\ _{t}ear_{t+j} + {}_{t}\omega_{t+j}$$

	Full Sample	Sub-samples (1) 76m1–79m3	(2) 79m4–83m3	(3) 83m4–87m10	(4) 87m11–90m6
United Kingdom					
α	-1.62	-4.47	-3.44	0.58	2.21
	(1.18)	(1.11)	(1.78)	(1.62)	(0.64)
β	1.60**	-0.42**	1.62**	1.72	1.28**
	(0.31)	(0.83)	(0.31)	(0.60)	(0.17)
$\alpha = 0; \beta = 1$	3.77	27.19*	4.04	19.06*	81.08*
France					
α	0.77	-1.12	-3.23	2.84	3.39
	(0.83)	(0.93)	(1.31)	(1.12)	(0.78)
β	0.64*	0.73	1.02	0.43	0.60**
	(0.22)	(0.69)	(0.23)	(0.42)	(0.22)
$\alpha = 0; \beta = 1$	2.95	2.18	40.37*	15.14*	27.79*
Belgium					
α	2.01	2.23	1.38	-0.28	-0.23
	(0.66)	(0.34)	(1.76)	(1.99)	(1.23)
β	0.78	1.76*	0.81	1.89	1.17
	(0.17)	(0.28)	(0.31)	(0.74)	(0.34)
$\alpha = 0; \beta = 1$	12.25*	79.08*	0.68	9.01**	0.40
Italy					
α	0.66	-2.08	-11.16	0.34	3.25
	(1.92)	(0.94)	(6.30)	(1.68)	(0.87)
β	0.13**	2.20	1.47	1.66	0.21*
	(0.50)	(0.74)	(1.10)	(0.62)	(0.23)
$\alpha = 0; \beta = 1$	5.54	5.58	11.43*	12.14*	14.33*
Ireland					
α	-1.40	-6.21	-7.19	-3.45	3.43
	(1.97)	(1.41)	(3.06)	(3.15)	(1.64)
β	1.33	0.57	1.66	4.14*	0.89
	(0.51)	(1.03)	(0.53)	(1.17)	(0.43)
$\alpha = 0; \beta = 1$	0.53	23.51*	10.63*	23.16*	11.54*

Standard errors are in parentheses. A * (**) alongside the coefficient β indicates rejection of the null $\beta = 1$ at 1% (5%). A * (**) alongside the $\alpha = 0$; $\beta = 1$ test statistic indicates rejection of the null at 1% (5%). The test statistics for the β coefficient are distributed as a 't' and are two-sided, with critical values 2.58 at 1% and 1.96 at 5%. The Wald test statistic for the joint hypothesis $\alpha = 0$; $\beta = 1$ is distributed as a $x^2(2)$, with critical values 9.21 at 1% and 5.99 at 5%.

in Table 2. The regressions were run over both the full sample (1976–90) and over four mutually-exclusive sub-samples. The aim of sub-sample estimation is to assess the extent to which real interest rate equality and covariance may have altered over time. The sub-samples chosen (following Ungerer et al (1990)) were: the pre-ERM period (January 1976–March 1979); the period from the inception of the EMS to the ERM realignment of March 1983, the latter which is widely believed to have marked the turning point in the fortunes of the system; the period between the March 1983 realignment and the Basle-Nyborg Accord of October 1987; thereafter until the end of the sample. In the tables below these are listed as sub-samples 1–4 respectively.

The results from Table 1 are broadly in line with our priors: over the full sample we are unable to reject real interest parity between Germany and all the other countries studied, with the exception of Belgium. Moreover the evidence of real interest parity is generally stronger for ERM than for non-ERM countries (ie the United Kingdom). This provides an indication that convergence of real interest rates is at least some of the way towards being achieved. More surprisingly, however, there are few indications that real interest rate equality is becoming noticeably more pronounced over time: there are rejections of this proposition in the last sub-sample for all the countries studied. Second, and probably more importantly, the estimates from the *ex-post* regressions appear to be inefficient, resulting in a failure to reject the null on too great a number of occasions. This may point towards some inadequacy in the instruments used and in turn supports the case for drawing more heavily upon our *ex-ante* estimates.

Table 2 reports the results from our *ex-ante* regressions. Again we find that over the full sample we are able to reject real interest rate parity only in the Belgium case and even then only marginally. This suggests a much higher degree of intra-EMS real interest rate linkage than was found between the US and Europe in the Cumby and Mishkin (1986) study. This generally accords with the qualitative evidence presented in Section 2, which suggested that there were strong real returns linkages across Europe. This must be considered encouraging from an EMU perspective, *vis.* convergence on a common real rate of return in steady-state.

There is, however, much less evidence of real interest rate linkages having increased monotonically over time: we are still able to reject real interest parity over the final sub-sample for most countries, although this rejection is much stronger for the non-ERM country, the United Kingdom, than for ERM member countries. The sub-sample estimates suggest that the pattern of real interest rate linkage is far from uniform, with coefficient instability clearly evident. Further, tests of the proposition that there is perfect cross-correlation between real rates can generally be rejected for most European countries.

The degree of coefficient instability evident in the estimates presented in Tables 1 and 2 may obviously be exaggerated by our choice of sub-samples and specifically the imposition of *a priori* structural breaks on the estimated relationships. This problem would seem especially likely given the sub-samples chosen: in practice, the temporal path of β is likely to be much less lumpy than implied by Tables 1 and 2. Additionally, as noted above, the absence of a well-defined structural relationship within our postulated model (3) is likely to hinder our ability to draw meaningful inferences, with omitted variable biases probable. Haldane and Hall (1991) have recently suggested a way of offsetting the two statistical problems highlighted above. This can be achieved by reformulating (3) as:

$$ {}_t r_{t+j}^{*} = \alpha_t + \beta_t \, {}_t r_{t+j} + {}_t e_{t+j} \qquad (5)$$

The time varying constant, α_t, partials out all systematic influences upon real interest rates other than those resulting from movements in German real interest rates. Omitted variable biases are thereby offset, with the error term in (5) hence stationary by construction. In addition, the presence of the time varying coefficient β_t allows for completely endogenous estimation of the temporal relationship between the real interest rate measures. Structural breaks in the linkages between real interest rates therefore need no longer be imposed *a priori*, as was true of the discrete period estimation reported in Tables 1 and 2. The technical details of estimating time varying parameter models of the form (5) are given in Haldane and Hall (1991) and therefore are not repeated here. Suffice to say, however, that we assume the stochastic parameters (state equations) behave as random walks, with the measurement equations taking the form (5), and estimate the whole system using Kalman filter techniques.

Figures 12–16 plot the time varying β coefficients for each of the countries. In very broad terms, there is some evidence of a higher degree of convergence of, and covariation between, real interest rates at the end than at the beginning of the ERM period. This enhanced convergence is evident both for the ERM and non-ERM countries. The notable exception is Italy,[1] whose real interest rate linkage with Germany is largely unaltered from the early 1980s. The large and growing divergences between the relative fiscal positions of Germany and Italy over the period may be partly responsible for this finding. In general terms our estimates indicate that, while some progress has been made towards an equilibration of real interest rates across some ERM countries (β_t is higher currently than was the case in 1979), this convergence is far from perfect even in the latest period (β_t is less than unity at the end of the sample).

As with our earlier discrete period analysis, a relatively high degree of coefficient evolution is evident in the estimates, with the pattern of convergence clearly non-monotonic. For example, in the UK case there is a clear structural break in the

(1) France is a further apparent anomaly.

Time varying parameters, β_t

Figure 12: France

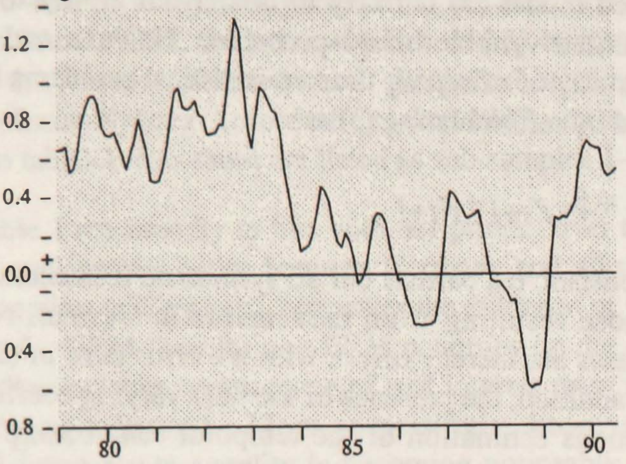


Figure 13: United Kingdom

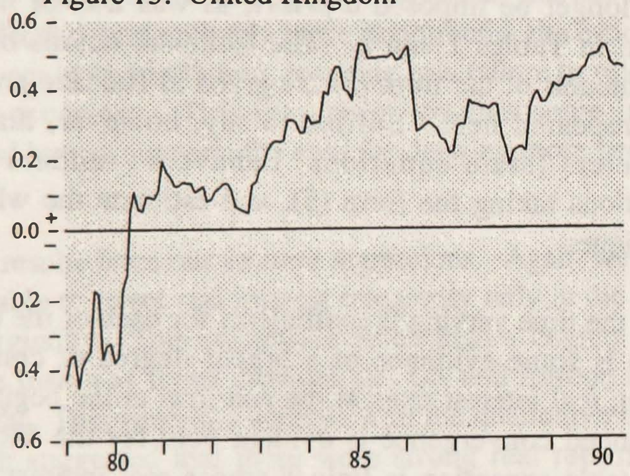


Figure 14: Belgium

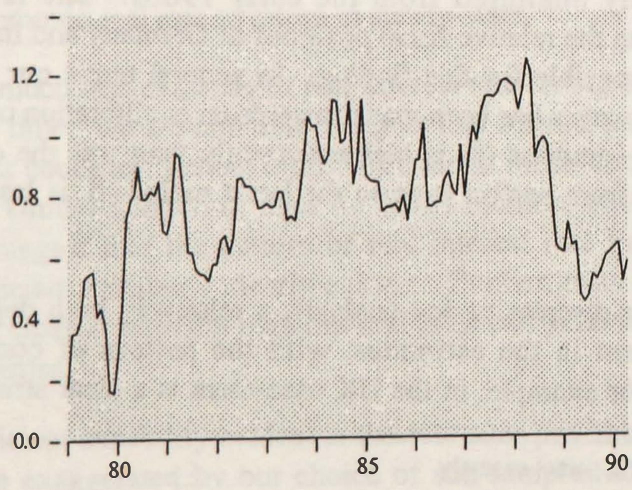


Figure 15: Italy

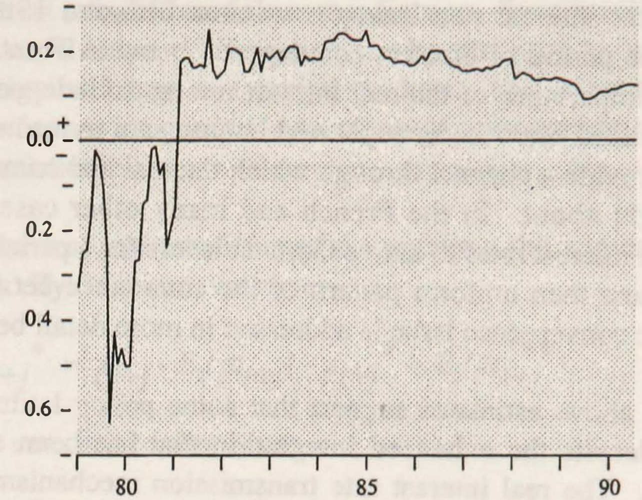
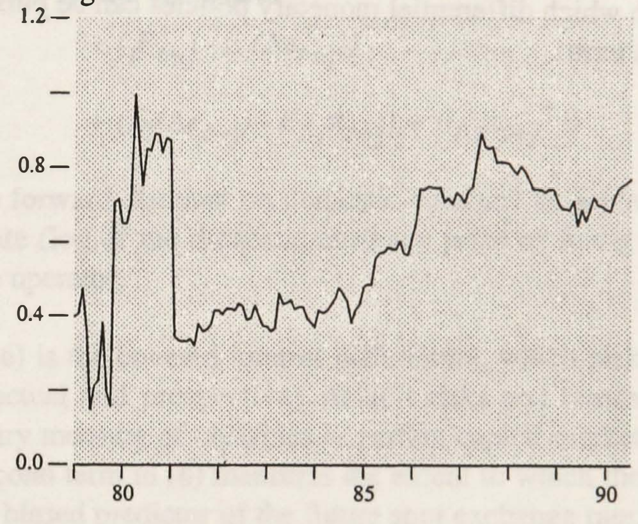


Figure 16: Ireland



relationship at the beginning of 1980 corresponding to the shift towards a more restrictive monetary stance. For Ireland, there are clear structural breaks at the beginning of the EMS period as policy adjustments were put into place, but thereafter there is a relatively smooth convergence of real interest rate linkages. Interestingly in the French case, real interest rate linkages weaken between 1983–88, generally acknowledged as the period of greatest convergence between France and Germany. This highlights the importance of the real interest rate as an independent policy tool, and specifically the need often to decouple real interest rate movements between two countries so as to provide a channel through which the residual convergence between them can be brought about. In the French and many other cases, this meant an overshooting of real interest rates to above German levels for a period during the mid-1980s—hence the less than uniform pattern of the covariance between real interest rates. This dynamic convergence issue is addressed in more detail below.

Further to this, the above estimates suggest that some policy latitude has remained within the ERM, despite the enhanced integration that has been evident since the system's inception. The real interest rate transmission mechanism, while far from being independent of real interest rates elsewhere in the ERM, therefore does offer one channel through which differential monetary policies can be pursued, at least over the short to medium term.

III Decomposing the real interest rate differential

Having established that the real interest rate linkage between Germany and other European countries is less than perfect, we now aim to consider in greater depth the *source* of real interest rate differentials across Europe. Specifically, we aim to assess the extent to which real interest differentials can be attributed to imperfect integration of the goods, foreign exchange and capital markets. This provides an indication of the areas in which further integration is necessary if convergence on a single real rate of return is to be satisfied.

In decomposing the real interest differential we follow the approach of Frankel and MacArthur (1988). By definition we have:

$$ {}_t r_{t+j} - {}_t r_{t+j}^{*} = {}_t i_{t+j} - E_t({}_t \pi_{t+j}) - {}_t i_{t+j}^{*} + E_t({}_t \pi_{t+j}^{*}) $$

which we can rewrite simply as:

$$ \begin{aligned} {}_t r_{t+j} - {}_t r_{t+j}^{*} = & \ ({}_t i_{t+j} - {}_t i_{t+j}^{*} - {}_t F_{t+j}) \\ & + ({}_t F_{t+j} - E_t(\Delta e_{t+j})) \\ & + (E_t(\Delta e_{t+j}) - E_t({}_t \pi_{t+j}) + E_t({}_t \pi_{t+j}^{*})) \end{aligned} \qquad (6) $$

where ${}_t F_{t+j}$ is the forward discount on domestic currency between time t and $t + j$, e is the exchange rate (log of the domestic currency price of foreign exchange) and Δ the first difference operator.

The first term in (6) is the covered interest differential, which picks up the impact of capital controls (actual and prospective), default risks and transactions costs. It is therefore a summary measure of obstacles to perfect capital market integration across countries. The second term in (6) measures the extent to which the (riskless) forward exchange rate is a biased predictor of the future spot exchange rate. In the absence of systematic expectational errors, it is therefore a summary measure of the extent to which foreign exchange markets are imperfectly integrated, ie the degree of (imperfect) substitutability between currencies—or so-called risk premium. Taken together, the first two terms in (6) gauge the extent to which uncovered interest parity is violated; that is, the importance of imperfectly integrated financial (foreign exchange and capital) markets.(1) The third term in (6) measures the expected real depreciation of domestic currency, ie the extent to which *ex-ante* purchasing power

(1) Our definition of imperfect integration is fairly restrictive here in that it includes the risk premium. Perfect integration would require that returns across financial markets be equalised after accounting for expected exchange rate changes, as indicated by the forward exchange rate. Any divergence of returns in excess of this we define as resulting from imperfect integration.

parity is violated. This can be thought to arise as a result of imperfect integration of goods (and labour) markets. In summary, any real interest rate differential can be thought of as measuring the degree to which the markets in goods, capital and foreign exchange between two countries are imperfectly integrated. As such, the real interest differential, when decomposed as in (6), is suggestive of the markets in which further integration will be necessary if static convergence on a single real rate of return is to be achieved.

To derive an empirical decomposition of (6), we evidently need some measure of exchange rate expectations. In arriving at these expectations we drew upon an IV methodology, using as instruments a constant, a linear time trend, lagged rates of inflation (domestic and foreign), nominal interest rates (domestic and foreign) and forward exchange rates.(1) For inflation expectations we again drew upon our survey estimates. The empirical results are summarised in Table 3, which reports the sample means and coefficients of variation over the full sample and selected sub-samples (sub-samples 2–4 from Tables 1 and 2).

As regards capital controls,(2) the pattern from Table 3 is clearly one of diminishing importance over time, as might be expected, with the covered interest differential negligible for most countries over the last sub-sample: for France, the United Kingdom and Belgium the differential averages less than 0.1% between 1987–90, while for Ireland and Italy it is small and decreasing. Capital controls are clearly in evidence over earlier sub-samples, however, for example in France and Italy during the early EMS years of frequent realignment. The signs on the covered interest differential are also as expected during this period, implying restrictions on capital outflows from France and Italy as a means of reducing the probability of a successful speculative currency attack in anticipation of a realignment. The evidence is consistent with capital markets being the most highly integrated of the three markets considered, with the implication therefore being that little residual convergence of capital markets is necessary for convergence on a single real rate of return. The EC Capital Liberalisation Directive seems likely to have contributed greatly to such a dismantling of obstacles to free capital movement.

The foreign exchange risk premia reported in Table 3 are in almost all cases positive, indicating that investors, on average, were demanding a positive additional return for holding non-German currencies over the EMS period. This is again broadly in line with our expectations. The size of the premium is, however, subject to quite marked variation across countries, ranging from an average 4% in Italy to around ½% for the

(1) Alternatively, survey measures of exchange rate expectations could have been employed. Much of the survey data on exchange rate expectations proved, however, to be of relatively poor quality.

(2) Default risks and transactions costs represent further reasons why a covered interest differential may persist. In practice, these effects are likely to be small and fixed.

Table 3: Capital controls (CC), risk premia (RP) and real exchange rate changes (RX); %

	Full Sample		Sub-samples (1) 79m3-83m3		(2) 83m4-87m10		(3) 87m11-90m6	
	Mean	Co.Var.	Mean	Co.Var.	Mean	Co.Var.	Mean	Co.Var.
United Kingdom								
CC	0.34	1.64	0.77	0.71	0.15	2.63	-0.02	11.68
RP	0.55	16.53	2.19	5.77	-2.60	2.22	3.26	1.38
RX	-1.39	6.72	1.66	8.15	-5.16	0.86	0.18	21.56
France								
CC	1.62	1.62	3.16	1.06	1.17	1.49	-0.04	7.88
RP	2.00	1.66	0.28	13.86	3.42	0.78	2.32	0.85
RX	0.50	9.90	0.28	23.76	0.87	4.84	0.19	8.67
Belgium								
CC	0.46	2.80	1.11	1.69	0.17	3.04	-0.10	2.44
RP	1.81	1.83	0.33	13.32	3.05	0.70	2.04	0.71
RX	0.33	11.32	-1.21	4.10	0.90	2.99	1.89	0.68
Italy								
CC	1.79	1.68	3.58	1.12	0.34	4.27	1.38	0.45
RP	4.27	0.75	4.16	1.03	3.76	0.65	5.37	0.33
RX	4.69	1.30	9.17	0.79	1.40	2.29	3.13	0.67
Ireland								
CC	-1.40	2.14	-3.10	1.40	-0.52	2.28	-0.40	2.70
RP	1.57	3.28	-0.10	68.37	2.63	1.53	1.95	1.41
RX	1.66	4.12	6.83	0.96	-1.41	4.13	-0.75	4.95

Co.Var. denotes the coefficient of variation. A positive sign for capital controls (CC) suggests restrictions on capital inflows into Germany and/or on capital outflows from the other country, and vice-versa for CC negative. A positive sign for the risk premium (RP) indicates a positive premium for the other country (ie, a discount for Germany). A positive sign for the real exchange rate (RX) indicates a real depreciation (appreciation) of the deutschemark (other currency).

United Kingdom over the full sample. The large and persistent risk premium on the lira can perhaps be explained largely by the worsening budgetary position of Italy over the period, with the risk premium on the lira at its largest in the final sub-sample. In terms of static convergence, therefore, it is evident that a relatively high degree of residual integration will be necessary between currencies in Europe before ultimate real interest rate convergence can be achieved: there is imperfect substitutability between intra-ERM currencies, despite the progress made in this respect by the operation of the ERM.

Finally we turn to the real exchange rate. Though real exchange rate movements are evidently subject to quite marked variation between sub-samples, broadly speaking real exchange rate effects are larger than those arising as a result of capital controls or risk premia.[1] Imperfect integration of goods (and labour) markets therefore appears to be the greatest single contributor to the observed real interest differentials across the countries studied. It follows that it is enhanced integration of goods markets that is most needed if there is to be static convergence of real returns across Europe. This finding is generally in line with our *ex-ante* expectations regarding the relative degrees of integration of European financial and goods markets.

Tables 4-6 report the results of some formal econometric tests of the importance of capital controls, risk premia and real exchange rate movements. The methodology used in deriving these test statistics is reported in Appendix 2. Hansen and Hodrick covariance matrix adjustments were again applied to improve the efficiency of our estimates, given the underlying serial correlation problems likely within our regressions. In general the results from these tests accord with the informal inferences drawn from Table 3. For example regarding capital controls, we are unable to reject a semi-strong form of covered interest parity in the final subsample for all the countries studied, with the exception of Ireland (Table 4). There is, however, widespread evidence of sizeable risk premium and real exchange rate effects (Tables 5 and 6): significant risk premia (on the basis of both strong and semi-strong form tests) are present for all countries in the final sub-sample; while *ex-ante* purchasing power parity can similarly be (semi-strong form) rejected in all cases in the final sub-sample. The pattern of rejections over time also squares with our earlier analysis. For example, the importance of capital controls in France between 1979–83; the importance of real exchange rate movements for Italy and Ireland over this same period; and the significant risk premium effects apparent in Italy and the United Kingdom over the final sub-sample.

In summary, our estimates suggest that it is risk premia and real exchange rate movements which largely explain the prevalance and persistence of real interest rate differentials across Europe. This finding is broadly in line both with our priors and with the vast majority of academic evidence to date assessing these propositions (see, for example, Mishkin (1984), Gaab et al (1986)). Correspondingly, if further integration of markets across Europe is to be forthcoming then it is in goods (labour) and foreign exchange markets that this residual convergence is most needed.

There is, however, a less insidious explanation of these observed 'failures' of integration, as reflected in the real return differentials. This is that the elements in the real interest differential are themselves important endogenous transmission mechanisms, which may be important for the steady-state stability of the system.[2]

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(1) Although, of course, the well-rehearsed problems of finding an appropriate instrument set for expected exchange rate movements means that any decomposition between risk premia and expected real depreciation needs to be treated cautiously.

(2) See Haldane and Pradhan (1992).

Table 4: Tests of covered interest parity

$$_tF_{t+j} = \alpha_0 + \alpha_1 \left({}_ti_{t+j} - {}_ti_{t+j}^* \right) + e_{1t}$$

	Full Sample	Sub-samples (1) 79m3-83m3	(2) 83m4-87m10	(3) 87m11-90m6
United Kingdom				
Strong Form	13.03*	26.62*	22.45*	19.14*
Semi-strong Form	10.25	10.00	7.60	8.20
France				
Strong Form	24.63*	110.83*	9.33*	4.62
Semi-strong Form	5.30	5.90	2.40	14.50
Belgium				
Strong Form	4.01	3.50	15.41*	2.91
Semi-strong Form	4.33	1.31#	12.80	5.00
Italy				
Strong Form	11.05*	14.32*	1.97	118.40*
Semi-strong Form	2.47	1.09#	3.07	4.90
Ireland				
Strong Form	5.82	4.10	7.46**	8.90**
Semi-strong Form	1.23#	0.62#	5.22	0.89#

The strong form test of covered interest parity is reported as a Wald test statistic of the joint null hypothesis $\alpha_0 = 0$; $\alpha_1 = 1$, and is distributed as a $x^2(2)$ with critical values of 9.21 at 1% and 5.99 at 5%. A * (**) indicates a rejection of the null at 1% (5%).

The semi-strong form test of covered interest parity is a test of the null $\alpha_1 < 1/2$, and is distributed as a (one-sided) 't' with critical values of +2.32 at 1% and +1.65 at 5%. A # (##) indicates a failure to reject the null at 5% (1 %), ie evidence against covered interest parity.

Table 5: Tests for risk premium effects

$$E_t(\Delta e_{t+j}) = \beta_0 + \beta_1 \, {}_tF_{t+j} + e_{2t}$$

	Full Sample	Sub-samples (1) 79m3–83m3	(2) 83m4–87m10	(3) 87m11–90m6
United Kingdom				
Strong Form	143.82*	203.29*	199.01*	160.96*
Semi-strong Form	-10.44#	-12.58#	-11.55#	-9.50#
France				
Strong Form	16.50*	6.96**	16.64*	150.96*
Semi-strong Form	5.92	6.14	1.89##	-3.64#
Belgium				
Strong Form	11.33*	2.29	30.53*	51.03*
Semi-strong Form	3.54	2.64	0.75#	-2.36#
Italy				
Strong Form	73.57*	36.34*	71.92*	215.64*
Semi-strong Form	1.08#	-0.625#	-0.68#	-1.48#
Ireland				
Strong Form	83.70*	135.53*	110.65*	6.89**
Semi-strong Form	-4.80#	-7.45#	-6.47#	-0.72#

The strong form test of risk premium effects is reported as a Wald test statistic of the joint null hypothesis $\beta_0 = 0$; $\beta_1 = 1$, and is distributed as a $x^2(2)$ with critical values of 9.21 at 1% and 5.99 at 5%. A * (**) indicates a rejection of the null at 1% (5%).

The semi-strong form test for risk premia is a test of the null $\beta_1 < 1/2$, and is distributed as a (one-sided) 't' with critical values +2.32 at 1% and +1.65 at 5%. A # (##) indicates a failure to reject the null at 5% (1%), ie evidence in favour of significant risk premium effects.

Table 6: Tests of ex-ante purchasing power parity

$$E_t(\Delta e_{t+j}) = \chi_0 + \chi_1 ({}_t\pi_{t+j} - {}_t\pi_{t+j}^*) + e_{3t}$$

	Full Sample	Sub-samples (1) 79m3-83m3	(2) 83m4-87m10	(3) 87m11-90m6
United Kingdom				
Strong Form	52.15*	39.03*	16.09*	0.72
Semi-strong Form	-5.67#	-5.00#	0.17#	-0.21*
France				
Strong Form	0.62	3.53	10.02*	14.02*
Semi-strong Form	1.91##	2.77	-1.45#	-1.52#
Belgium				
Strong Form	0.53	0.73	21.27*	82.69*
Semi-strong Form	1.54#	1.15#	-0.75#	-0.54#
Italy				
Strong Form	172.83*	190.83*	36.88*	113.35*
Semi-strong Form	-3.13#	-5.00#	-1.79#	-3.56#
Ireland				
Strong Form	157.68*	333.72*	123.47*	22.31*
Semi-strong Form	-6.38#	-6.36#	-6.27#	-3.03#

The strong form test of *ex-ante* PPP is reported as a Wald test statistic of the joint null hypothesis $\chi_0 = 0$; $\chi_1 = 1$, and is distributed as a $\chi^2(2)$ with critical values of 9.21 at 1% and 5.99 at 5%. A * (**) indicates a rejection of the null at 1% (5%).

The semi-strong form test of *ex-ante* PPP is a test of the null $\chi_1 < ½$ is distributed as a (one-sided) 't' with critical values +2.32 at 1% and +1.65 at 5%. A # (##) indicates a failure to reject the null at 5% (1%), ie evidence against *ex-ante* purchasing power parity.

The theoretical analysis of, for example, Giavazzi and Pagano (1988) suggests an important role for the real exchange rate as a disciplining device upon deviant inflationary countries. Our estimates provide relatively clear evidence of the real exchange rate having operated as such a disciplining device for ERM countries (who have, on average, experienced a real appreciation against the deutschemark), but typically not for the non-ERM country the United Kingdom (which has, on average, experienced a real depreciation). For example, in the Italian and Irish cases we see clear evidence of the important role played by the real exchange rate in the transitional period between 1979–83, with real appreciations of 9% and 7% respectively (see Giavazzi and Pagano (1988), Kremers (1990)); whereas for France the clearest evidence of the real exchange rate having acted as a disciplining device is to be found in the period between 1983–87. This suggests an important role for real exchange rate movements as self-equilibrating mechanisms between high and low inflation countries in semi-fixed exchange rate regimes, and hence provides one rationalisation for the observed failures of real interest parity. Similar theoretical arguments can be used to help justify the prevalence of risk premia and capital control influences.

IV Conclusions

The pattern of real interest rate differentials in the EMS suggests that most are due to the imperfect integration of goods markets and to the presence of risk premia in foreign exchange markets. Not surprisingly, the importance of capital controls in accounting for these differentials has declined since the inception of the EMS. While there is some evidence of real interest rate convergence having occurred during the 1980s, significant differentials still persist for some countries, indicating that further convergence may be necessary. In the context of the debate over fiscal policy rules in a monetary union, our results, although tentative, suggest that real interest rates are not as yet sufficiently interdependent to support those who have argued that one country's fiscal deficit will necessarily affect fully real interest rates for all other member countries. However, to the extent that a monetary union entails a major regime change (for example, with countries unable to monetise fiscal deficits), questions of this type are difficult to answer satisfactorily using as a benchmark a non-monetary union regime such as the ERM. Quantifying the scale of such a regime shift in the expectations-formation processes of private sector agents when moving to a monetary union is likely to remain a fruitful line of future research in the context of on-going EMU discussions.

Appendix 1

In this Appendix we outline the Carlson and Parkin (1975) method of quantifying qualitative responses on inflation expectations. The responses are those of industrialists in all the major EC countries, who are asked about their 'selling price expectations over the coming months'. Apart from the United Kingdom (where the question explicitly refers to the next four months), we have taken the question to relate to inflation expectations over the next three months. There are three categories of response: the proportion of respondents who expect inflation to rise (R), those who expect inflation to be lower (F), and those who expect inflation to remain the same (S). The actual responses published are the balance statistic R-F, and S.

In order to derive a quantitative measure of inflation expectations from these qualitative responses, Carlson and Parkin make a number of assumptions regarding the underlying distribution from which the responses are drawn, and specifically about the meaning of the response 'prices expected to remain the same, (or unchanged)'. The distribution of responses is assumed to be normal. For those responding 'prices unchanged' it is assumed that there is a range of inflation outcomes (an indifference interval) which are sufficiently near the current value not to be considered significantly different from it. Furthermore, the indifference interval is taken to be symmetric and constant over time. These assumptions are fairly restrictive and in recent work (see Pesaran (1987), Pesaran and Wright (1991) and Batchelor (1986)) have been relaxed. However, since the expectations data available does not ask industrialists about their perceptions of current prices, we are forced to impose these restrictions.

There is one important difference between the Carlson and Parkin method and the one used here which is taken directly from Pesaran and Wright (1991). Carlson and Parkin assume that a proportion of the respondents reporting 'prices unchanged' are incapable of answering the question, and this proportion is derived from an auxiliary regression. Since the respondents here are industrialists being asked about their own selling prices, we assume that all of them are capable of answering the question.

If π_t is the current rate of inflation (three month change annualised), then, assuming a normal distribution for responses, we define the indifference interval around π_t as ranging from $-a_{t+1}$ to b_{t+1}. If expected next period inflation is greater than b_{t+1} then respondents are assumed to report a rise (R_{t+1}). Similarly, if expected inflation is less than $-a_{t+1}$ then respondents are assumed to report a fall (F_{t+1}). The cumulative probabilities of expected rises and falls are then given as:

$$Pr(\pi_{t+1} < -a_{t+1}) = F^e_{t+1} \quad \textbf{(a)}$$

$$Pr(\pi_{t+1} > b_{t+1}) = R^e_{t+1} \quad \textbf{(b)}$$

where $R^e{}_{t+1}$ is the proportion of respondents expecting a rise in prices over the next three months, ie the probability of the change in prices being greater than b_{t+1}. Similarly, $F^e{}_{t+1}$ is the probability of the change being less (more negative) than $-a_{t+1}$.

Defining $Z_{t+1} = ((\pi_{t+1} - \pi^e{}_{t+1}) / \sigma^e{}_{t+1})$ as a standard normal variate, equations (a) and (b) can be written as:

$$Pr(Z_{t+1} < (-a_{t+1} - \pi^e{}_{t+1}) / \sigma^e{}_{t+1})) = F^e{}_{t+1} \tag{c}$$

$$Pr(Z_{t+1} < (b_{t+1} - \pi^e{}_{t+1}) / \sigma^e{}_{t+1})) = 1 - R^e{}_{t+1} \tag{d}$$

Additionally, we define:

$$(-a_{t+1} - \pi^e{}_{t+1}) / \sigma^e{}_{t+1} = f^e{}_{t+1} \tag{e}$$

$$(b_{t+1} - \pi^e{}_{t+1}) / \sigma^e{}_{t+1} = r^e{}_{t+1} \tag{f}$$

If the cumulative probability distribution of a normal variate is given by $\Phi(.)$, then $Pr(Z_{t+1} < - a_{t+1}) = \Phi(f^e{}_{t+1})$ and $Pr(Z_{t+1} < b_{t+1}) = \Phi(r^e{}_{t+1})$. Note that $f^e{}_{t+1}$ and $r^e{}_{t+1}$ are the values of $-a_{t+1}$ and b_{t+1} in the standard normal distribution.

In order to now solve for $\pi^e{}_{t+1}$, we eliminate $\sigma^e{}_{t+1}$ by combining (e) and (f). Recalling our earlier assumption that the indifference interval is symmetric, ie $a_{t+1} = b_{t+1}$, and constant over time (at say c) we can write:

$$\pi^e{}_{t+1} = c\,((f^e{}_{t+1} + r^e{}_{t+1}) / (f^e{}_{t+1} - r^e{}_{t+1})) \tag{g}$$

If the indifference interval is not constant over time, then by definition (g) will have time-varying parameters (see Pesaran and Wright (1991)). Once we have the values of $f^e{}_{t+1}$ and $r^e{}_{t+1}$, we need c in order to scale the expected series so that it can be expressed in terms of actual inflation. If we had information on industrialists' perceptions of current inflation we could get an estimate for c (c^*, say) from the current inflation analogue of (g), ie

$$\pi_t = c^*\,((f_t + r_t) / (f_t - r_t)) \tag{h}$$

However, since we do not have data on current inflation perceptions, we follow the method of Carlson and Parkin. The estimate of c is then given by:

$$c^{\#} = T^{-1} \Sigma_{t=1}{}^{T} \pi_t / T^{-1} \Sigma_{t=1}{}^{T} ((f^e{}_{t+1} + r^e{}_{t+1}) / (f^e{}_{t+1} - r^e{}_{t+1})) \tag{i}$$

The estimate of c from (i) is then used in (g) to calculate $\pi^e{}_{t+1}$, the measure of expected inflation.

Appendix 2

Our methodology in testing the parity conditions implied by the real interest differential draws upon the study by Gaab et al (1986). Specifically, our estimated equations for gauging the importance of, respectively, capital controls, risk premia and real exchange rate movements are:

$$ {}_tF_{t+j} = \alpha_0 + \alpha_1 ({}_ti_{t+j} - {}_ti_{t+j}^*) + e_{1t} \tag{i}$$

$$ E_t(\Delta e_{t+j}) = \beta_0 + \beta_1 \, {}_tF_{t+j} + e_{2t} \tag{ii}$$

$$ E_t(\Delta e_{t+j}) = \chi_0 + \chi_1 ({}_t\pi_{t+j} - {}_t\pi_{t+j}^*) + e_{3t} \tag{iii}$$

Following Gaab et al (1986) we distinguish strong and semi-strong forms of the hypotheses under consideration; for example *vis.* capital controls (covered interest parity) we have:

Strong form. $\quad {}_ti_{t+j} - {}_ti_{t+j}^* = {}_tF_{t+j}$

That is, observed values of the nominal interest differential and the forward discount coincide. Substituting into (i) gives us the joint restriction $\alpha_0 = 0$; $\alpha_1 = 1$ as our strong form test of covered interest parity.

Semi-Strong Form. We have from (i):

$$ \alpha_1 = Cov({}_tF_{t+j}, {}_ti_{t+j} - {}_ti_{t+j}^*) \, / \, \mathrm{Var}({}_ti_{t+j} - {}_ti_{t+j}^*) \tag{iv}$$

which, given a capital controls term (τ_t) in our uncovered interest parity condition, can be rewritten as:

$$ = (\mathrm{Var}({}_tF_{t+j}) + \mathrm{Cov}(\tau_t, {}_tF_{t+j})) \, / \, (\mathrm{Var}({}_tF_{t+j}) + \mathrm{Var}(\tau_t) + 2\,\mathrm{Cov}(\tau_t, {}_tF_{t+j})) $$

The semi-strong form test of covered interest parity is a variance inequality test: $\mathrm{Var}(\tau_t) < \mathrm{Var}({}_tF_{t+1})$. That is, the variance of obstacles to free capital movement is less than the variance of forward exchange rate movements. Given (iv), this therefore amounts to a test of the restriction $\alpha_1 > 1/2$.

Equivalent forms of these hypotheses were tested for risk premium and real exchange rate effects using (ii) and (iii) as a basis.

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