A New Interpretation of the Exchange Rate - Yield Differential Nexus*

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Abstract

Empirical studies have had difficulty in establishing the long run relationship between real exchange rates and real yield differentials predicted by sticky price exchange rate models. We revisit this issue in a nonstationary panel regression framework. This facilitates estimation of a long run parameter even when the underlying relationship is subject to permanent shocks or the variables do not cointegrate. The slope coefficient estimate from a sample of 23 industrialized countries 1973M1-1998M12 has the correct sign and is statistically significant for both short and long term yields. These results support fundamentals-based models of exchange rate behaviour while permitting real factors to play a role. Moreover they indicate that capital markets integration is more advanced than hitherto believed.

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1 Introduction

In an important contribution, Meese and Rogoff (1988) found no evidence of a stable relationship between real exchange rates (RERs) and real interest differentials for a monthly sample of three US dollar RERs 1974-1985. Such a result went against the predictions of most monetary and portfolio balance models of the exchange rate and, significantly, their model provided no improvement over a random walk in terms of predictive power. They concluded that their findings are consistent with real disturbances, such as productivity shocks, being the major source of exchange rate volatility.

Despite this, the debate on the exchange rate-yield differential nexus in recent years has by and large been dominated by the cointegration approach that precludes or downplays a role for precisely such shocks. A cointegrating relationship by definition excludes permanent disturbances since these would lead to inconsistent coefficient estimates in a levels OLS regression of exchange rates on yield differentials. The verdict from extant time-series cointegration tests is on balance unfavorable to qualified (Campbell and Clarida 1987; Meese and Rogoff 1988; Hunter 1992; Edison and Pauls 1993; Juselius 1995; Edison and Melick 1999; Wu 1999). However the evidence from panel cointegration tests tends to be more positive. For instance, Chortareas and Driver (2001) using the Pedroni (1997) and Kao (1999) tests find support for cointegration for a sample of 11 open economies though not for their full panel of 18 OECD economies.

In this paper we redress the balance. The paper's first contribution is that it employs a framework that can accommodate permanent as well as temporary shocks to the RER. There are several rationales for according a role to permanent shocks besides that provided by Meese and Rogoff (1988). The residual term in a levels regression of exchange rates on yield differentials minimally captures the expected equilibrium exchange rate. If the typical real macroeconomic fundamentals underlying the latter — such as relative productivity growth, current account imbalances or GDP differentials — have stochastic trends, the regression error will also be nonstationary. This provides one obvious rationale for permanent shocks impacting on the RER.

Another rationale for accommodating a nonstationary regression error is that it is difficult to establish with certainty the true time series properties of variables for finite sample spans such as the post-Bretton Woods era. More specifically, unit root tests may be unable to reject the null under particular circumstances. For instance, A.M. Taylor (2001) shows that temporal aggregation and nonlinearities can make mean-reverting variables appear very persistent or nonstationary. He demonstrates the latter in the context of regressions of spot exchange rates on price differentials.[1] Nakagawa (2002) estimates a nonlinear time-series model along these lines to show that the relationship between real exchange rates and real yield differentials holds for large but not for all deviations of the RER from its equilibrium value.[2] This finding is consistent with Baxter (1994) who demonstrates an association between the medium- to low- (but not the high-) frequency components of exchange rates and relative yields. In other words, the link is strongest at business cycle and trend frequencies. More importantly her model predicts that exchange rates and yield differentials should not cointegrate.

The paper's second contribution it that it provides consistent estimates of the long run, real exchange rate-yield differential association — which is a reflection of a common stochastic trend — within a new econometric framework which permits real exchange rates to be subject to idiosyncratic permanent shocks. In doing so it builds on the recent nonstationary panel literature which shows that long-run effects are not exclusively associated with cointegrating relationships (Pesaran and Smith 1995; Phillips and Moon 1999; Kao 1999; Coakley, Fuertes and Smith, 2001). The former three studies develop asymptotic theory to demonstrate that, by adding independent cross-section information, it is possible to estimate consistently a long run coefficient even in the absence of time series cointegration.[3] The intuition is that, by averaging across individuals, the noise — the covariance between the nonstationary regressor and nonstationary error — that swamps the signal is alleviated.

Coakley et al. (2001) show via Monte Carlo simulations that these asymptotic results are relevant for the finite sample panel dimensions typical of post-Bretton Woods studies. Since our linear regression framework accommodates permanent disturbance, effectively it permits factors other than real yield differentials to be responsible for part of the observed persistence in RER. In this regard, our contribution may serve to bridge the gap between cointegration studies and those that advocate an exclusive role for real disturbances.

[1] The view that trade or other frictions can induce nonlinearities in real exchange rate behaviour — by inducing a no-arbitrage band for small deviations around equilibrium — has garnered particular support. See Obstfeld and A.M. Taylor 1997; O'Connell 1998; Coakley and Fuertes 2001; A.M. Taylor 2001; M.P. Taylor, Peel and Sarno 2001. Relatedly, Obstfeld and Rogoff (2000) posit that trade frictions giving rise to nonlinear relationships may provide the key to resolving longstanding puzzles in international finance.

[2] This 'band of inaction' is in line with some positive evidence on cointegration between exchange rates and yield differentials deriving from the enhanced power properties of the tests deployed (Chortareas and Driver 2001; Edison and Melick, 1999).

[3] See Baltagi and Kao (2000), Phillips and Moon (2000) and Smith (2000) for surveys.

The …nal contribution is that our empirical analysis provides evidence of a signi…cant long run relationship between exchange rates and both short term and long term yield di¤erentials for 23 industrialized countries over the 1973-1998 period. While existing studies have documented a degree of integration between commodity markets (RERs) and bond markets (long term yields), to the best of our knowledge our study is the …rst to suggest integration between commodity and money (short term yields) markets. The latter adds to the recent evidence on a high degree of capital mobility from a variety of related perspectives (Fujii and Chinn 2001; Lane and Milese-Ferretti 2001; Coakley, Fuertes and Spagnolo 2001).

The plan of the paper is as follows. Section 2 presents a theoretical model motivating the real exchange rate-yield di¤erential relation and discusses the statistical framework. Section 3 describes the data and analyses the empirical results. A …nal section concludes.

2 Reduced Form RER Model

2.1 Theoretical framework and issues

The original Meese and Rogo¤ (1988) study was based on a quasi-reduced form, RER model embodying both sticky prices or slow adjustment and rational expectations. Essentially their model was a real version of the rational expectations monetary models of the exchange rate of Dornbusch (1976)[4] and Hooper and Morton (1982). This has become the workhorse of most cointegration-based studies of the real exchange rate-yield di¤erential relation. Let q_t denote the RER which is de…ned as:

$$q_t \equiv s_t - (p_t - p_t^*) \tag{1}$$

where s_t is the current nominal exchange rate or the domestic price of foreign currency, p_t and p_t^* are domestic and foreign prices, respectively, and all variables are in natural logarithms.

Meese and Rogo¤ made three crucial assumptions. The …rst derives from sticky price models and posits that deviations of q_t from its long run equilibrium level $\bar{q}_t$ are monotonically eroded:

$$E_t(q_{t+m} - \bar{q}_{t+m}) = \phi^m(q_t - \bar{q}_t), \quad 0 < \phi < 1 \tag{2}$$

where $E_t(\cdot)$ denotes time t expectations. Note that this embodies adaptive expectations and the absence of further shocks. Second they assumed ex

[4] See Rogo¤ (2002) for a recent overview of the impact of the Dornbusch model.

ante PPP holds:

$$E_t(\bar{q}_{t+m}) = \bar{q}_t \tag{3}$$

One implication is that the equilibrium RER is also an I(1) process. Using (3), (2) can be rearranged as:

$$q_t = \bar{q}_t + \beta(E_t q_{t+m} - q_t) \tag{4}$$

where $\beta = 1/(\phi^m - 1)$. If q_t is an I(1) process, then (4) clearly implies that deviations from equilibrium $(q_t - \bar{q}_t)$ are stationary or that q_t and $\bar{q}_t$ cointegrate. At this point it is worth noting that a nonstationary RER does not necessarily preclude PPP. Long run relative PPP or the stationarity or deviations from relative PPP, i.e. $\Delta q_t \sim I(0)$, suggests that price di¤erentials are re‡ected one-for-one in nominal exchange rates while the latter may be subject to other permanent shocks inducing nonstationary real exchange rates (see Coakley, Flood and Taylor, 2001).

Their third and …nal assumption is uncovered interest parity (UIP). The risk-adjusted version can be written as:

$$E_t(s_{t+1} - s_t) = i_{t,t+1} - i^*_{t,t+1} + \lambda_t \tag{5}$$

where $i_{t,t+1}$ $(i^*_{t,t+1})$ is the time t return on a domestic (foreign), one-period asset and λ_t is an exogenous risk premium re‡ecting the less-than-perfect substitutability of the foreign and domestic assets.[5] It embodies the forward-looking asset view of exchange rate determination. The …nal building block is the Fisher E¤ect formulated as $i_{t,t+1} = r_{t,t+1} + E_t[\pi_{t,t+1}]$ where $r_{t,t+1}$ $(\pi_{t,t+1})$ is the real interest rate (domestic in‡ation rate) for the period t to $t+1$. Subtracting in‡ation from the nominal variables on both sides of (5) gives:

$$E_t(q_{t+1} - q_t) = r_{t,t+1} - r^*_{t,t+1} + \lambda_t \tag{6}$$

Finally using (4) and (6) we have:

$$q_t = \bar{q}_t + \beta(r_{t,t+1} - r^*_{t,t+1}) + \lambda_t \tag{7}$$

where the slope coe¢ cient, β, embodies a negative relationship between the RER and yield di¤erentials as in Frankel (1989) and is a decreasing function

[5] This is derived as follows. De…ne the nominal exchange rate level, S_t, as the value that equates the expected return on assets denominated in di¤erent currencies $(1 + i^*_{t,t+1})\frac{E_t S_{t+1}}{S_t} = (1 + i_{t,t+1})(1 + \lambda_t)$. Using logarithms (denoted by lower case) and ignoring cross terms, the UIP relation follows. While Meese and Rogo¤ (1988) did not accommodate a risk premium, later empirical studies have extensively done so.

of the maturity of the interest rate. This relationship is called real uncovered interest parity (RUIP). Although this model has been extensively employed as the basis of cointegration tests, it does have a statistical imbalance problem as highlighted by Engel (2000). It is clear from (4) that $q_t - \bar{q}_t$ must be stationary if q_t is I(1) and this together with equation (7) implies that $r_{t,t+m} - r^*_{t,t+m}$ is also stationary. Then by definition q_t cannot cointegrate with $r_{t,t+m} - r^*_{t,t+m}$ and this would undermine the basis for cointegration studies.

As Engel (2000) puts it "...it might make sense to treat $q_t - \bar{q}_t$ and $r_{t,t+m} - r^*_{t,t+m}$ as if they were nonstationary variables in examining dynamics, if the error term in $q_t = \bar{q}_t + \beta(r_{t,t+1} - r^*_{t,t+1})$ were much less persistent than either $q_t - \bar{q}_t$ or $r_{t,t+m} - r^*_{t,t+m}$." In our formulation the risk premium, which has been argued in the literature to be either a non-zero constant or a time-varying but stationary process, plays the role of residual term. Hence, in seeking to redress the balance deviations from equilibrium are permitted to be very persistent or, equivalently, the adjustment parameter ϕ in equation (2) is allowed to be very close to unity. If $\phi = 1$ in the limit, deviations from equilibrium are permanent and $q_t - \bar{q}_t$ is nonstationary but then the parameter β in (4) and subsequent equations is not defined. However if one assumes that ϕ is arbitrarily close to unity, say 0.98, $q_t - \bar{q}_t$ and by implication yield differentials will be observationally equivalent to $I(1)$ processes. Such an assumption can capture the marked inertia of RERs or their persistent deviations from long-run equilibrium. Possible explanations for the latter include the influence of chartists or non-fundamental traders[6], real disturbances such as productivity shocks shifting the real exchange rate permanently (Caporale and Pittis 2002) and trade frictions leading to a no-arbitrage band.

This new formulation of RUIP implicitly assumes that real interest rate differentials may be nonstationary contra the original Meese and Rogoff (1988) model. However it is in line with the fact that, despite ongoing financial market integration, most empirical studies have found at best ambiguous evidence on real interest parity. There are several explanations for this. Slowly changing stances of monetary policy (Hoffman and MacDonald, 2001) and asymmetric feedback rules reflecting opportunistic central back behavior (Coakley and Fuertes, 2002) can make real interest rates vir-

[6] See Taylor and Alllen (1992) on the role of chartists in foreign exchange rate markets. More generally, the recent behavioral finance literature highlights the limits to arbitrage as the basis for the persistence of deviations of financial variables from their equilibrium levels. See Shleifer's (2000) critique of Friedman (1953) who interestingly used the foreign exchange market to argue for the efficacy of arbitrage.

tually indistinguishable from integrated processes in typical …nite samples. Moreover, persistent deviations from a constant yield di¤erential have been rationalized for long horizons as arising from the lack of homogeneity or liquidity in government bonds (Meese and Rogo¤, 1988) or from relative commodity price movements arising from real shocks to the economy.

Since setting ϕ arbitrarily close to unity makes …nite-sample RERs and yield di¤erentials observationally equivalent to I(1) processes, cointegration analysis may appear to be the next logical step from (7). However a body of research has posited $\bar{q}_t$ as a function of fundamental variables such as the cumulated current account balance to GDP ratio which may be represented by a random walk (Edison and Melick 1999). Other recent work also suggests that the equilibrium RER path may be in‡uenced by productivity di¤erentials, saving-investment decisions or GDP di¤erentials (Lane and Milesi-Ferreti 2000; Ho¤mann and MacDonald 2001). In these studies $\bar{q}_t$ is thus assumed to be nonstationary since the fundamentals themselves may be subject to permanent shocks.

This paper does not seek to unravel the role of the I(1) fundamental variables or other driving forces behind $\bar{q}_t$. The implication is that q_t and $r_{t,t+m} - r^*_{t,t+m}$ may not cointegrate even if both are I(1) processes. This is the basis for our long-run regression analysis that permits the unobservable equilibrium RER, $\bar{q}_t$ to enter the error term. E¤ectively, this means that although RERs and real yield di¤erentials may share a common contributory nonstationary source (a common stochastic trend), we view real shocks as an unobservable, nonstationary idiosyncratic factor variable in the determination of RERs.

2.2 Nonstationary panel regression approach

Rewriting (7) in panel regression form we have:

$$q_{it} = \alpha_i + \beta_i (r_{it} - r^*_{it}) + u_{it}, i = 1, ..., N, t = 1, ..., T \tag{8}$$

where i is the country (group) index. The error term u_{it} therefore at a minimum captures the risk premium and the expected equilibrium RER. In addition, a MA term due to overlapping data may underlie u_{it}. In either case, the systematic variability of the equilibrium RER will swamp the stationary behavior of the risk premium and any serial dependencies due to temporal aggregation or other reasons and will induce I(1) behavior in u_{it}. Hence, the main issue becomes how to estimate the long run coe¢ cient β when the errors are observationally equivalent to I(1) processes.

Phillips and Moon (1999) and Kao (1999) demonstrate that some panel datasets o¤er the prospect of overcoming the spurious regression problem of pure time series. More particularly, they show that in large N, large T panels one can obtain consistent estimates of a long-run average parameter even if there is no time-series cointegration at an individual level or, equivalently, when the error term as well as the variables are I(1). The intuition is that the averaging over cross-section units lessens the noise in the relationship — the covariance between the I(1) error and the I(1) regressor — that induces the spurious regression problem. In e¤ect the latter means that panel regressions may lead to a stronger overall signal than pure time-series regressions.

Take the simple data generating process:

$$y_{it} = \alpha_i + \beta_i x_{it} + u_{it}, i = 1, ..., N, t = 1, ..., T \tag{9}$$

where y_{it} and x_{it} are both I(1) and suppose that u_{it} is also I(1) so that y_{it} and x_{it} are not cointegrated. The mean group (MG) or unweighted average panel estimator of Pesaran and Smith (1995) is de…ned by

$$\widehat{\beta}^{MG} = N^{-1} \sum_{i=1}^{N} \left(\frac{\sum_{t=1}^{T} \widetilde{y}_{it} \widetilde{x}_{it}}{\sum_{t=1}^{T} \widetilde{x}_{it}^2} \right) \tag{10}$$

where $\tilde{x}_{it} = x_{it} - \bar{x}_i$ and $\bar{x}_i = T^{-1} \sum_{t=1}^{T} x_{it}$ and similarly for $\tilde{y}_{it}$. The …xed e¤ects (FE) or weighted average estimator is given by

$$\hat{\beta}^{FE} = \sum_{i=1}^{N} w_i \left(\frac{\sum_{t=1}^{T} \widetilde{y}_{it} \widetilde{x}_{it}}{\sum_{t=1}^{T} \widetilde{x}_{it}^2} \right) = \beta + \frac{\sum_{i=1}^{N} \sum_{t=1}^{T} \tilde{x}_{it} u_{it}}{\sum_{i=1}^{N} \sum_{t=1}^{T} \tilde{x}_{it}^2} \tag{11}$$

where $w_i = S_i / \sum_i S_i$ with $S_i = \sum_t \widetilde{x}_{it}^2$. In a time-series setup the noise, $\sum_t \tilde{x}_{it} u_{it}$, swamps the signal and hence the OLS estimator will not converge to the true β_i (but to a nondegenerate random variate instead) even when T becomes large. However, this problem is alleviated in a panel context by averaging over i and so a consistent long-run coe¢cient estimate can be obtained as $N \to \infty$ and $T \to \infty$.[7]

In the context of nonstationary variables such as:

$$\begin{pmatrix} x_{it} \\ y_{it} \end{pmatrix} = \begin{pmatrix} x_{i,t-1} \\ y_{i,t-1} \end{pmatrix} + \begin{pmatrix} \varepsilon_{x,it} \\ \varepsilon_{y,it} \end{pmatrix}$$

[7] One caveat is in order. In line with most panel data work this asymptotic theory rests on the assumption of uncorrelated disturbances across groups. Little is known about the joint e¤ect of I(1) errors and between-group dependence.

where $\boldsymbol{\varepsilon}_{it} = (\varepsilon_{x,it}, \varepsilon_{y,it})'$ are I(0) innovations, Phillips and Moon (2000) show that it is possible to extend the classical notion of a regression coe¢cient to a long-run coe¢cient even in the absence of time series cointegration. Suppose that the two underlying variables are generated as follows:

$$\begin{aligned} y_{it} &= \beta_i z_{it} + w_{it} \\ x_{it} &= z_{it} \\ z_{it} &= z_{i,t-1} + \varepsilon_{z,it} \end{aligned}$$

where $\varepsilon_{z,it}$ is an I(0) innovation and w_{it} is a stochastic process. Let

$$E(\boldsymbol{\varepsilon}_{it}\boldsymbol{\varepsilon}_{it}') = \quad _i = \begin{bmatrix} \ ^i_{xx} & \ ^i_{xy} \\ \ ^i_{yx} & \ ^i_{yy} \end{bmatrix} \qquad (12)$$

denote the long-run covariance matrix of $\boldsymbol{\varepsilon}_{it}$ with $\ ^i_{xy} = \sum_{s=-\infty}^{\infty} E(\varepsilon_{x,i0}\varepsilon_{y,is})$. The classical regression coe¢cient can be extended to this nonstationary setup by de…ning $\beta_i = \ ^i_{xy}/\ ^i_{xx}$, the ratio of the long-run covariance between y and x to the long-run variance of x. Irrespective of whether w_{it} is I(0) or I(1) the latter measures a long run association (the nonstationary common factor z_{it}) between x_{it} and y_{it}.

If w_{it} is I(0) or equivalently, if $\ _i$ has full rank, then β_i has the traditional interpretation of a cointegrating coe¢cient. By contrast, β_i is not a cointegrating coe¢cient if $y_{it} - \beta_i x_{it}$ is nonstationary due to the in‡uence of a nonstationary idiosyncratic factor w_{it}. However in the latter case, since the two series have a common driving nonstationary factor z_{it}, they will be correlated in the long run and this is precisely what β_i measures. The pooled FE estimator in this context is consistent for $\beta = E(\ ^i_{xy})/E(\ ^i_{xx})$, the ratio of the average across units of the long-run covariance between y and x to the average across units of the long-run variance of x. Phillips and Moon (1999) called this the long-run average regression coe¢cient. By contrast, the MG estimator measures the average of the long-run covariance to long-run variance ratio, $\tilde{\beta} = E(\ ^i_{xy}/\ ^i_{xx})$, and can be therefore termed an average long-run coe¢cient. These two estimators are identical when the variance of x_{it} is the same across units.

These asymptotic results are complemented by the Monte Carlo analysis in Coakley et al. (2001) that explores the small sample properties of the FE and pooled OLS (POLS) estimators as well as the MG estimator in nonstationary regressions.[8] Their experiments con…rm that the limit theory is

[8] Asymptotic results have not been established for the MG estimator but the estimator has been shown to be unbiased and correctly sized in …nite samples (typical of PPP studies) in Coakley et al. (2001).

relevant for panel dimensions typical of annual and monthly post-Bretton Woods studies. In particular the above panel estimators appear unbiased with dispersion that falls at rate $\sqrt{N}$ even when the error term is I(1). One contrasting aspect of the FE and MG estimators is that the standard errors of the former are incorrect in the presence of autocorrelated I(0) errors and I(1) errors. Asymptotically valid inferences in the latter case can be made using the non-standard covariance matrix derived in Phillips and Moon (1999) or computing empirical p-values from sieve bootstrap distributions as in Fuertes (2003).

3 Empirical analysis

3.1 Summary statistics

The sample comprises a panel of 23 industrialized countries whose members were dictated by data availability, particularly of yield series. The US is chosen from amongst these as numeraire country given the leading role of the dollar in international trade and capital flows. Four different panels are constructed combining the consumer price index (CPI) and producer price index (PPI) measures with short- (ST) and long-term (LT) yields 1973M1-1998M12. Data definitions and sources are detailed in Appendix A.[9]

The real yield differentials require inflation measures and these are constructed as follows. We use both ex post and ex ante real yields. The latter are calculated from a static expectations assumption, $E_t(\pi_{t,t+m}) = \pi_{t-m,t}$, and two smoothing procedures. One involves a 7-point two-sided moving average (MA) filter and the other is a Holt-Winters (HW) filter which averages past and present values and generalizes the single exponential smoother by adding linear trend and seasonal components.[10] A value of 0.1 for the level, trend and seasonal damping factors seems a reasonable compromise for the HW filter for all countries on the basis of the one-step-ahead, root-mean-squared error loss function.

Inflation rates are computed over the span of the ST and LT yields.[11]

[9] The CPI-ST panel ($N = 19$ countries) excludes Luxembourg, Australia and South Africa due to lack of data. The CPI-LT panel ($N = 18$) excludes the latter two countries, Greece and Iceland. The PPI-ST panel ($N = 16$) excludes Belgium, Iceland, Italy, Luxembourg, New Zealand and Portugal. The PPI-LT panel ($N = 15$) excludes the latter six countries as well as Greece.

[10] We also employed a 25-point MA filter but this makes little difference to the results. For a discussion of forecasting with smoothing techniques see Harvey (1989).

[11] This implies eliminating roughly one third of our sample (time series dimension) for the LT panels. However, as noted by Meese and Rogoff (1988), computing inflation rates

The latter contrasts with most existing studies that, due to sample size constraints, de‡ate LT rates using a long MA smoother of ST (usually three-month-ahead) in‡ation rates. Nevertheless, we also de‡ate the LT yields using a 25-month two-sided MA smoother of three-month-ahead in‡ation both to facilitate comparisons with the literature and because the latter seems to work quite well in practice.

The MA procedure generates the smoothest series for the ex ante yield di¤erential while the static expectations proxy lies at the other extreme. As an illustration, Figure 1 presents these two measures for the Canadian and German short term interest di¤erential.

[Figure 1 around here]

The RER and ex ante yield di¤erential from the 7-point MA proxy are depicted for each country in Figures 2 and 3 for the LT and ST bonds, respectively. The plots show substantial short term deviations that may relate to the limits to arbitrage stemming from market sentiment. An alternative explanation is what Obstfeld and Rogo¤ (2000) call the 'exchange rate disconnect puzzle' to describe the weak high frequency links between the exchange rate and the rest of the economy. However, the trend behavior of the real exchange rate seems to track that of the real yield di¤erential quite well for some countries over particular periods and especially for the long-term securities. This is in line with Baxter's (1994) …nding that the strongest relationship between these two variables may be in the business cycle and trend components.

[Figures 2 and 3 around here]

This provides prima facie evidence that the two variables are related over the sample period. This conjecture is now assessed more formally.

In keeping with the literature the time series properties of each variable are examined using the augmented Dickey-Fuller (ADF) test. Table 1 reports the results for the CPI- and PPI-based real exchange rates and ST and LT yield di¤erentials based on the static expectations and 7-point MA proxies for CPI in‡ation.[12]

over the term of the bonds may produce real yield di¤erential measures which are closer to the relevant ones.

[12] The lag length is selected using Ng and Perron (1995) testing-down approach starting from $\bar{k} = 12$. The test results for the remaining cases — those using real yield series constructed using the HW …lter and PPI measures — are qualitatively similar and are available on request.

[Table 1 around here]

They indicate that it is not possible to reject the non-stationarity null for either the RER or yield di¤erential series in a majority of cases.

This conclusion for the yield series may be called into question against a backdrop of highly integrated capital markets. However, most empirical studies using both standard and multivariate or panel approaches fail to …nd cogent evidence of stationary behavior in LT or ST real yield di¤erentials (Baum and Barkoulas 2002; Chortareas and Driver 2001; Ho¤mann and MacDonald 2001; Edison and Melick 1999). On balance therefore the results are in line with the consensus view that RER and yield di¤erentials are observationally equivalent to nonstationary processes over the post Bretton-Woods sample period. Finally, the unreported Engle-Granger test results indicate that virtually all individual residual series from (8) are nonstationary I(1) with a few exceptions for the LT yield regressions (at the 10% level) such as Belgium, Denmark, Luxembourg and the UK. This motivates the next section.

3.2 Long run panel estimates

The MG and FE panel estimators allow for heterogeneity to varying degrees. The FE estimator permits heterogeneous intercepts α_i to allow for country-speci…c …xed e¤ects but imposes equality of slopes, $\beta_i = \beta$, in (8). It is computed using (11) and its standard error by $se(\hat{\beta}^{FE}) = s/\sqrt{\sum_{i=1}^{N}\sum_{t=1}^{T}\widetilde{x}_{it}^2}$ where s is the standard error of the within regression. By contrast, the MG estimator permits heterogeneity in both intercept and slope and is computed using (10). Its standard error is calculated by $se(\widehat{\beta}^{MG}) = \sigma(\hat{\beta}_i^{OLS})/\sqrt{N}$ where $\sigma(\hat{\beta}_i^{OLS})$ is the sample standard deviation of the individual OLS estimates. In the context of large N panels however, heterogeneity is an important issue and arguably the MG is more adept than the FE estimator at capturing this. Nonetheless we report the results from both.

Table 2 reports the estimation results for the RER-real yield di¤erential relationship using ST interest rates and CPI in‡ation.

[Table 2 around here]

The …rst two columns report the MG and FE coe¢ cient estimates for the different in‡ation proxies considered. Although there is some variation, the estimates are correctly signed and statistically significant in all cases.[13] Countries with individual estimates more than two standard deviations from the mean are trimmed to control for excessive heterogeneity. The resultant estimates given in the final two columns are also significantly negative.[14] The PPI panels also produce statistically significant slope coefficients. For instance, using static expectations the MG and FE estimates after discarding an outlier (Germany) are -0.426 (0.069) and -0.374 (0.023), respectively. Unreported results for the other two expected inflation proxies are qualitatively similar to their CPI counterparts.

The two panel estimates differ in magnitude, especially for the CPI panels where the FE estimates are approximately half the value of their MG counterparts. Since the expected sampling variability of the MG estimator is larger than that of the FE estimator, we pay particular attention to the latter (Coakley et al. 2001). It is worth noting that the FE estimates of the slope coefficient from the ST yield panels are not compatible with monotonic adjustment following shocks.[15] Accordingly an alternative RUIP hypothesis proposed by Edison and Melick (1999) that the slope coefficient is $\beta_i = -1/4$ (for three month maturity and annualized yield rates) is tested. This does not embody a specific notion of adjustment but rather assumes long run proportionality between real yield differential shocks and subsequent exchange rate adjustment.

Empirical p-values are computed using the nonparametric sieve bootstrap approximation proposed in Fuertes (2003) to test the hypothesis $\beta = -1/4$. The latter's Monte Carlo simulations suggest that this procedure is useful in eliminating the substantial size distortions of conventional FE t-tests based on the N(0,1) quantiles. In line with the earlier Engle-Granger test results and to allow for some degree of heterogeneity — rather than assuming $u_{it} \sim$I(1) for all i — this bootstrap method assumes that the integration order of each individual residual series is unknown and has to be pre-tested. The resampling scheme permits errors with general temporal de-

[13] This is inferred from the MG estimates and standard errors. The magnitude of the FE standard errors (near 1/7 times the value of their MG counterparts) is in line with the Monte Carlo findings in Coakley, Fuertes and Smith (2001) of oversized t tests from this estimator in the I(1) error case.

[14] We repeated the exercise excluding countries for which the ADF test gave some evidence of I(0) behavior for real yield differentials. The estimation results are qualitatively unaffected. For instance, for the MA filter, after excluding Denmark, Japan, New Zealand and Sweden, the MG and FE estimates (and standard errors) are -1.33(0.256) and -0.496(0.029).

[15] The theoretical slope of $\beta_i = 1/(\phi_i^m - 1)$ is not consistent with values of the estimated slope coefficient lying between 0 and -1.

pendence such as AR(I)MA and cross-section heterogeneity. Table 2 reports bootstrap p-values from 10,000 replications in brackets. Interestingly, they suggest that the long run RUIP hypothesis cannot be rejected for any of the in‡ation measures considered while, unsurprisingly, it is rejected using the dramatically undersized FE standard errors.

Table 3 reports the results for the CPI panels using LT yields.

[Table 3 around here]

Since the static expectations and ex post in‡ation proxies produce very close results, only those for the former are reported. The table also contains the case of LT yields de‡ated by a 25-point MA …lter of 3-month-ahead in‡ation for comparison with the literature. The coe¢ cient estimates for the LT yield di¤erentials di¤er from those in Table 2 in two respects. First, in line with the theoretical priors, their sample mean at about -4 is substantially larger (in absolute terms) than that in the analogous ST case. Indeed, t-statistics are computed to test the hypothesis that $\beta = -4$ for each coe¢ cient estimate. Using either the conventional standard errors for the average long run (MG) coe¢ cient or the sieve bootstrap p-values for the long run average (FE) coe¢ cient the hypothesis cannot be rejected in most cases. This indicates a clear term structure relationship in yield di¤erentials.

Second, the FE estimates are much closer to the MG estimates than in the ST panels which is in line with the fact that the variance of the yield di¤erential, $var(r_{it} - r_{it}^{*})$, is more homogeneous across countries for the LT maturities. It also suggests that cross-section heterogeneity in the long run RER - yield di¤erential nexus (di¤erent β_i) is less marked for LT yields.[16] There is a parallel here with the evidence relating to the expectations hypothesis of the term structure which suggests that in general there is greater evidence (a larger signal-noise ratio) in favour of the expectations hypothesis at the long end than at the short end of the maturity spectrum (Campbell and Shiller 1991). The PPI panels gave qualitatively similar results.

The supporting evidence for RUIP from the LT yield panels is in line with some existing studies such as that of Chortareas and Driver (2001). Interestingly, we can back out an estimate of the speed of adjustment of RERs for the case of LT yields. Given that $\hat{\beta} = -4$ on average for the

[16] We also computed real yield di¤erentials using a 7-year-ahead in‡ation measure since the Macaulay duration of the observed 10-year, coupon-paying bonds (at 6.543) roughly corresponds to that for 7-year pure discount bonds, but this makes little di¤erence to the results. The MG estimate (s.e.) is -2.052 (0.514) and -2.165 (0.533) for the static expectations and MA smoother, respectively.

different cases considered and since the frequency of our data is monthly, the implied $\phi^m = (1 + 1/\beta)$ for $m = 120$ months is 0.75. This gives an estimate of $\hat{\phi} = 0.9976$ which, as we surmised earlier, is very close to unity. Adopting Macaulay's duration (at 6.5 years) as a more adequate measure of effective maturity for coupon-paying bonds, then $\phi^{78} = 0.75$ which gives $\hat{\phi} = 0.996$. In other words, the RER seems indistinguishable from an I(1) process. This is consistent both with Dornbusch's (1976) overshooting model and with persistent RERs.

To conclude, the significant long run relationship found between real exchange rates and ST yield differentials represents a novel finding in the RUIP context. However, the LT yield differential results are more indicative of support for the Dornbusch (1976) hypothesis of monotonic albeit very persistent RER adjustment following overshooting in the wake of shocks. More generally, our findings in favour of RUIP using both ST and LT yield differentials are consistent with mobility in both short and long term international capital flows. They are also in line with the globalization and integration of financial markets in recent decades.[17]

4 Conclusions

This paper revisits the real exchange rate-yield differential parity relationship linking international financial and commodity markets. Prior work formulated within a cointegration framework has by and large failed to establish unambiguous evidence of a long run nexus between these variables. The alternative approach proposed in this paper permits nonstationary regression errors to accommodate real as well as temporary shocks impacting on the real exchange rate. In so doing it builds on recent advances in nonstationary panel data theory by Pesaran and Smith (1995), Kao (1999) and Phillips and Moon (1999). These demonstrate that by adding cross-section information one can consistently estimate long-run relationships between two nonstationary variables that may or may not cointegrate.

Our results from a panel of 23 industrialized countries 1973M1-1998M12 produce evidence of a significant long-run, inverse relationship between real exchange rates and real yield differentials. Moreover the regression results using long term yields support the monotonic real exchange rate adjustment

[17] These results carry over to a more recent 10-year span. For instance, the MG and FE estimates for the period 1988M1-1998M12 are -0.645 (0.111) and -0.545 (0.028), respectively, for the ST yields in the naive expectations case after discarding Italy as an outlier.

predicted by sticky price and portfolio balance exchange rate models and are in line with other recent findings such as those of Chortareas and Driver (2001). To the best of our knowledge, our study is the first to report a significant long run relationship between real exchange rates and short term yield differentials. The latter finding is plausible since it is consistent with mobility in short term as well as long term international capital flows.

The overall conclusion is that, on the one hand, real interest differentials are reflected in real exchange rates in the long-run as sticky-price theories of exchange rate determination suggest. On the other, the latter does not preclude macroeconomic real shocks from playing the role of an idiosyncratic nonstationary factor in the persistence and volatility of real exchange rates. Our results add to other recent evidence that capital market integration is more advanced than hitherto believed. An avenue for further research is to extend our framework to a dynamic panel regression setup assuming identical long-run coefficients across panel members but allowing the short-run coefficients to differ and to reexamine the evidence in this context.

Appendix A: Data Sources

The data cover the period 1973M1 to 1998M12. End-of-month bilateral exchange rates vis-à-vis the US dollar and CPIs and PPI data are taken from Datastream. Since interest rate data sources are more diverse, they are detailed in the table below. Short term (3-month) rates are mostly call money market rates from the IMF (line 60b). Long term rates are average yields to maturity on bellwether government bonds with residual maturities between 9 and 10 years from the IMF (line 61) or OECD.

	ST interest rate	LT interest rate
AU[a]	Money market rate (IMF)	OECD
OE	Money market rate (IMF)	OECD, Bank of Austria
BG	Money market rate (IMF)	IMF
CN	Treasury bill rate (IMF)	OECD
DK	Money market rate (IMF)	IMF, Bank of Denmark
FR	Money market rate (IMF)	OECD
GE	Money market rate (IMF)	OECD
GR	Comm. banks deposits (IMF)	—
IC	Discount rate (IMF)	—
IR	Interbank rate (OECD)	OECD
IT	Money market rate (IMF)	OECD
JP	Banks bills rate (Bank of Japan)	OECD
LX	—	IMF
NH	Money market rate (OECD)	IMF
NZ	Banks bills rate (OECD)	OECD
NW	Money market rate (IMF)	IMF
PT	Discount rate (IMF)	IMF[b]
SA	Money market rate (IMF)	IMF
SP	Money market rate (OECD)	IMF[c]
SD	Money market rate (IMF)	OECD
SW	Euro-deposit rate (OECD)	IMF
UK	Money market rate (IMF)	OECD
US	Money market rate (IMF)	OECD

[a] Australia (AU), Austria (OE, Belgium (BG), Canada (CN), Denmark (DK), France (FR), Germany (GE), Greece (GR), Iceland (IC), Italy (IT), Japan (JP), Luxembourg (LX), New Zealand (NZ), Norway (NW), Portugal (PT), South Africa (SA), Spain (SP), Sweden (SD), Switzerland (SW). [b] Starts in 1976M1. [c] Starts in 1978M1.

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Table 1 Augmented Dickey-Fuller test results

	Real exchange rate		Real interest rate differential			
	Inflation measure		ST interest rates		LT interest rates	
	CPI	PPI	Static	MA	Static	MA
AU[a]	—	-2.57(0)[b]	—	—	—	—
OE	-2.33(11)	-2.43(11)	-2.05(9)	-1.87(9)	-1.76(10)	-2.02(1)
BG	-2.33(11)	—	-2.16(9)	-2.24(7)	-1.99(12)	-1.67(8)
CN	-.95(12)	-2.23(11)	-2.86(9)	-2.83(11)	-1.92(10)	-2.30(1)
DK	-2.36(11)	-1.59(6)	-3.28(10)*	-3.14(8)*	-2.66(12)	-2.21(1)
FR	-2.01(6)	-2.34(6)	-1.96(9)	-1.96(7)	-1.40(9)	-1.69(1)
GE	-2.43(11)	-2.26(11)	-2.32(10)	-2.38(11)	-1.51(10)	-1.89(1)
GR	-1.98(12)	-2.05(11)	-3.49(10)**	-2.51(11)	—	—
IC	-2.94(12)*	—	-2.07(9)	-1.80(11)	—	—
IR	-2.49(9)	-2.56(11)	-2.54(9)	-2.40(7)	-1.97(10)	-1.99(10)
IT	-2.50(11)	—	-2.29(9)	-2.54(7)	-1.35(10)	-1.31(1)
JP	-1.87(12)	-1.55(6)	-4.73(9)**	-5.60(8)**	-1.81(10)	-1.83(8)
LX	-2.35(11)	—	—	—	-2.36(12)	-2.14(1)
NH	-2.44(11)	-2.16(11)	-2.34(9)	-2.72(10)	-1.54(10)	-1.72(1)
NZ	-2.66(12)	—	-3.94(10)**	-3.81(6)**	-3.62(11)**	-3.28(9)*
NW	-2.10(9)	-1.90(9)	-2.22(9)	-2.43(10)	-2.28(9)	-2.22(0)
PT	-1.65(11)	—	-2.76(9)	-2.72(7)	-1.88(6)	-1.53(5)
SA	—	-2.27(9)	—	—	—	—
SP	-2.00(7)	-2.37(8)	-3.74(6)**	-2.70(11)	-1.56(10)	-1.64(1)
SD	-2.01(9)	-2.20(9)	-3.33(10)*	-3.39(8)*	-2.40(10)	-2.39(12)
SW	-2.56(11)	-2.45(11)	-3.16(9)*	-2.77(9)	-1.97(10)	-2.09(3)
UK	-2.58(11)	-1.70(11)	-2.16(9)	-2.23(11)	-2.17(10)	-2.36(7)

Notes: [a]See country codes in Appendix A. [b]The number of lags used is shown in parentheses. All test regressions include a constant term. The largest possible number of observations is used for each variable, T=312 for real exchange rates, T=309 (static expectations) and T=306 (MA) for ST interest rate differentials and T=192 (static) and T=186 (MA) for LT differentials. Real interest rate differentials are based on CPIs. *significant at the 5 % level. **significant at the 1% level.

Table 2 Slope coe¢ cient estimates for short term yield di¤erentials

	Full panel					Outliers excluded	
	$\hat{\beta}^{MG}$	$\hat{\beta}^{FE}$	Min$_{\hat{\beta}_i}$	Max$_{\hat{\beta}_i}$	$Z > 2^a$	$\hat{\beta}^{MG}$	$\hat{\beta}^{FE}$
Static	-.712 (.127)[b]	-.308 (.023)[b] [.244][c]	-1.90 (GE)	-.049 (NZ)	-2.14 (GE)	-.645 (.115)	-.294 (.023) [.360]
MA_7	-1.23 (.215)	-.508 (.028) [.184]	-3.19 (OE)	-.137 (IC)	-2.09 -2.05 (OE) (GE)	-1.00 (.165)	-.461 (.028) [.178]
HW	-.989 (.180)	-.428 (.025) [.162]	-2.30 (OE)	.215 (NZ)	—	-.989 (.180)	-.428 (.025) [.162]
Ex post	-.748 (.119)	-.353 (.022) [.234]	-1.83 (GE)	-.095 (IC)	-2.08 (GE)	-.688 (.109)	-.339 (.022) [.208]

[a]Outlier countries whose individual estimates are more than two standard deviations away from the mean. [b]MG and FE standard errors in parenthesis. [c]Sieve bootstrap p-values for the two-sided test $H_0 : \beta^{FE} = -1/4$ from 10,000 replications in brackets.

Table 3 Slope coe¢cient estimate for long term yield di¤erentials

	Full panel					Outliers excluded	
	$\hat{\beta}^{MG}$	$\hat{\beta}^{FE}$	Min$_{\hat{\beta}_i}$	Max$_{\hat{\beta}_i}$	$Z > 2$	$\hat{\beta}^{MG}$	$\hat{\beta}^{FE}$
Static[a]	-3.68	-3.48	-5.28	-.46	2.54	-3.87	-3.60
	(.299)[c]	(.076)[c]	(OE)	(CN)	(CN)	(.246)	(.078)
		[.132][d]					[.080]
MA$_7$	-4.95	-4.43	-7.51	-.828	2.20	-5.19	-4.56
	(.441)	(.087)	(LX)	(CN)	(CN)	(.391)	(.083)
		[.406]					[.416]
HW	-4.01	-3.58	-5.95	-.684	2.15	-4.20	-3.69
	(.364)	(.080)	(LX)	(CN)	(CN)	(.326)	(.082)
		[.088]					[.164]
MA$^b_{25}$	-3.15	-2.83	-8.03	-.019	-2.52	-2.82	-2.74
	(.484)	(.081)	(FR)	(CN)	(FR)	(.383)	(.082)
		[.000]					[.000]

[a] The e¤ective sample period for the cases based on 10-year-ahead in‡ation measures is shorter. This is 1985M1-1998M12 for the static and HW cases and 1985M4-1998M9 for the MA$_7$ case. [b] 3-month-ahead in‡ation is used. The e¤ective sample is 1974M4-1997M12 and hence this case excludes Portugal and Spain whose yields are observed post-1975M12 and 1977M12, respectively. [c] MG and FE standard errors in parenthesis. [d] Sieve bootstrap p-values for the two-sided test $H_0 : \beta^{FE} = -4$ from 10,000 replications in brackets.

Figure 1 Alternative measures of ex ante real short-term yield differential (inverse)

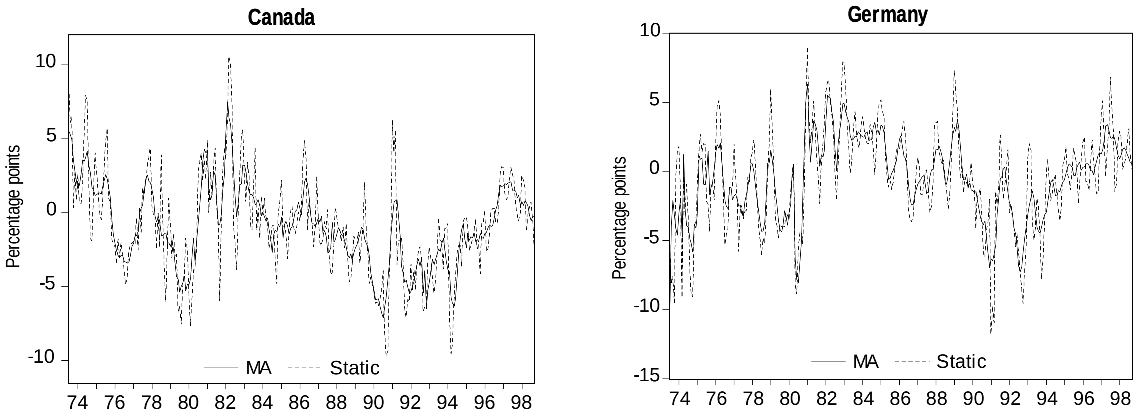
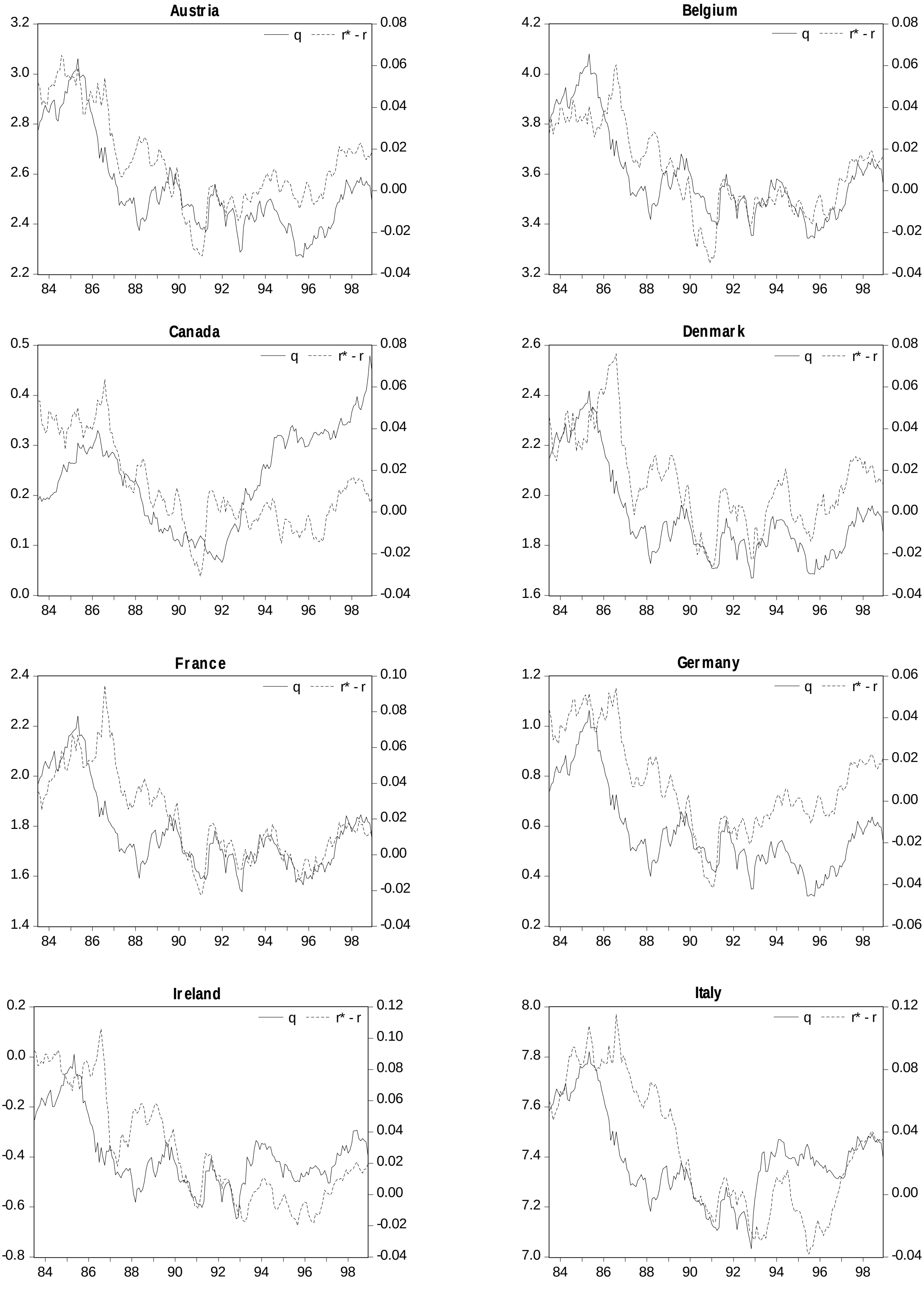
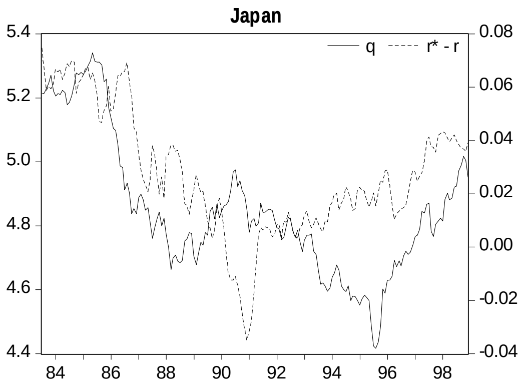
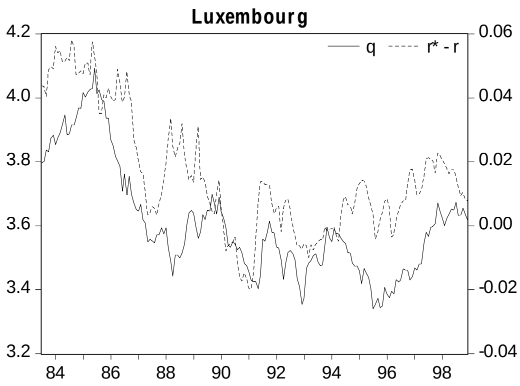


Figure 2 Real exchange rates (left scale) and ex ante long-term yield differentials (right scale)




Japan
q
r* - r
5.4
5.2
5.0
4.8
4.6
4.4
0.08
0.06
0.04
0.02
0.00
-0.02
-0.04
84 86 88 90 92 94 96 98


Luxembourg
q
r* - r
4.2
4.0
3.8
3.6
3.4
3.2
0.06
0.04
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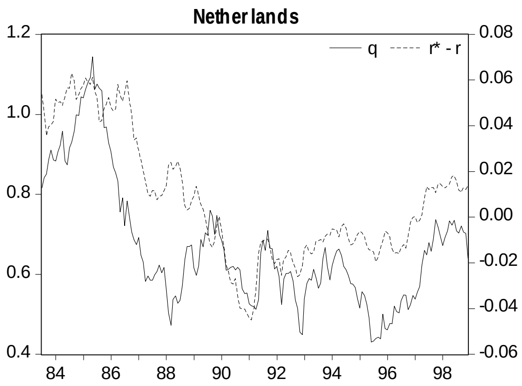
-0.02
-0.04
84 86 88 90 92 94 96 98


Netherlands
q
r* - r
1.2
1.0
0.8
0.6
0.4
0.08
0.06
0.04
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0.00
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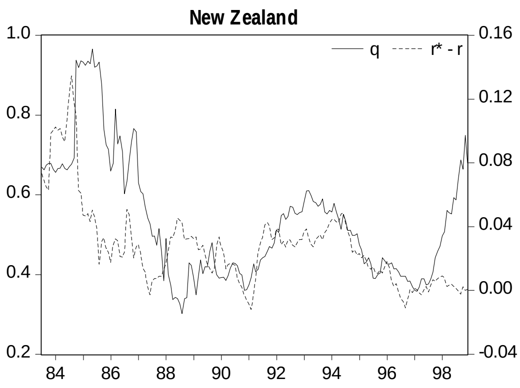
84 86 88 90 92 94 96 98


New Zealand
q
r* - r
1.0
0.8
0.6
0.4
0.2
0.16
0.12
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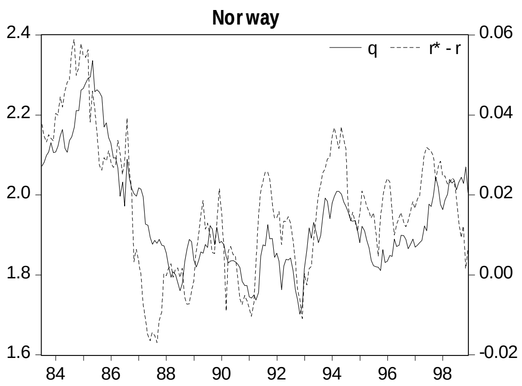
0.00
-0.04
84 86 88 90 92 94 96 98


Norway
q
r* - r
2.4
2.2
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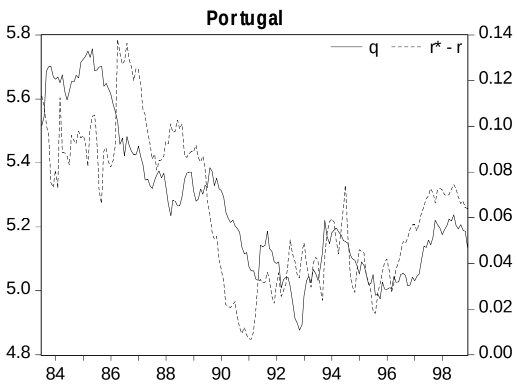
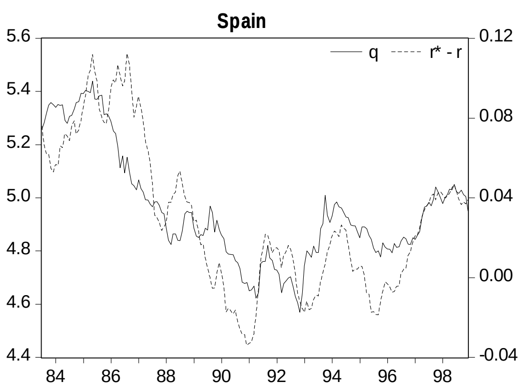
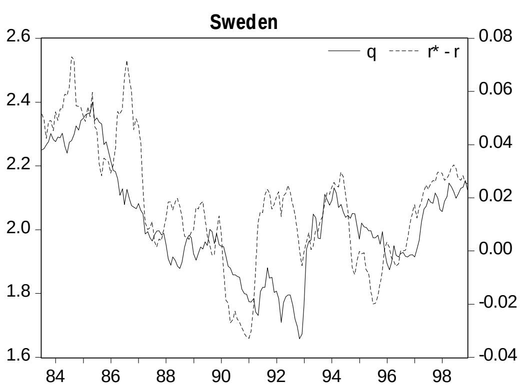
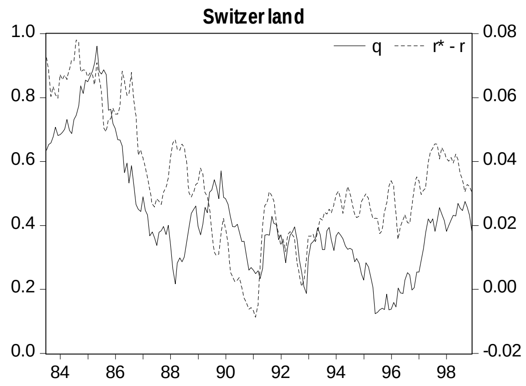
84 86 88 90 92 94 96 98


Portugal
q
r* - r
5.8
5.6
5.4
5.2
5.0
4.8
0.14
0.12
0.10
0.08
0.06
0.04
0.02
0.00
84 86 88 90 92 94 96 98


Spain
q
r* - r
5.6
5.4
5.2
5.0
4.8
4.6
4.4
0.12
0.08
0.04
0.00
-0.04
84 86 88 90 92 94 96 98


Sweden
q
r* - r
2.6
2.4
2.2
2.0
1.8
1.6
0.08
0.06
0.04
0.02
0.00
-0.02
-0.04
84 86 88 90 92 94 96 98


Switzerland
q
r* - r
1.0
0.8
0.6
0.4
0.2
0.0
0.08
0.06
0.04
0.02
0.00
-0.02
84
86
88
90
92
94
96
98

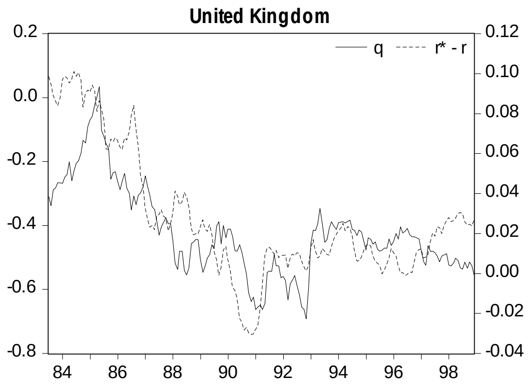
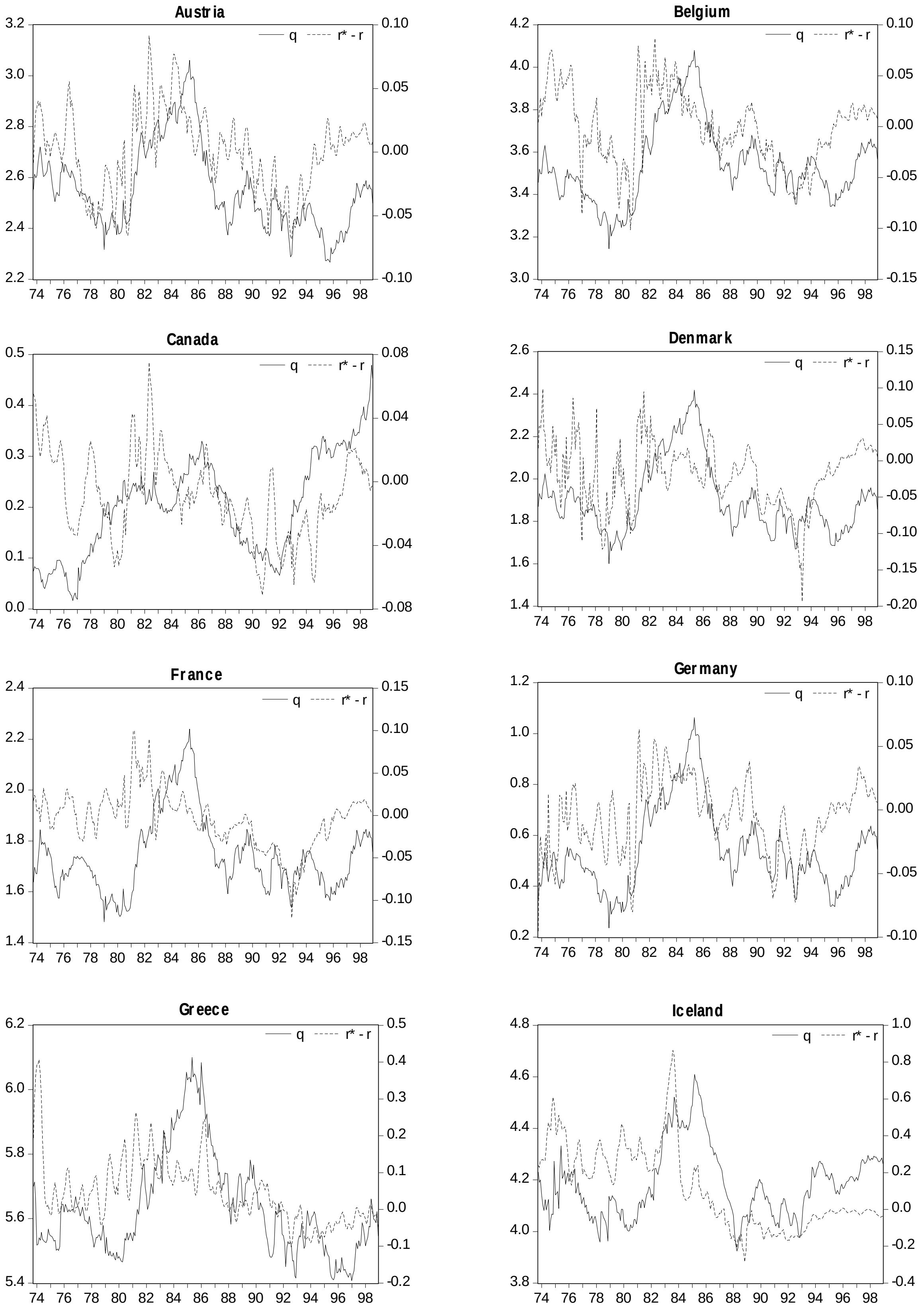
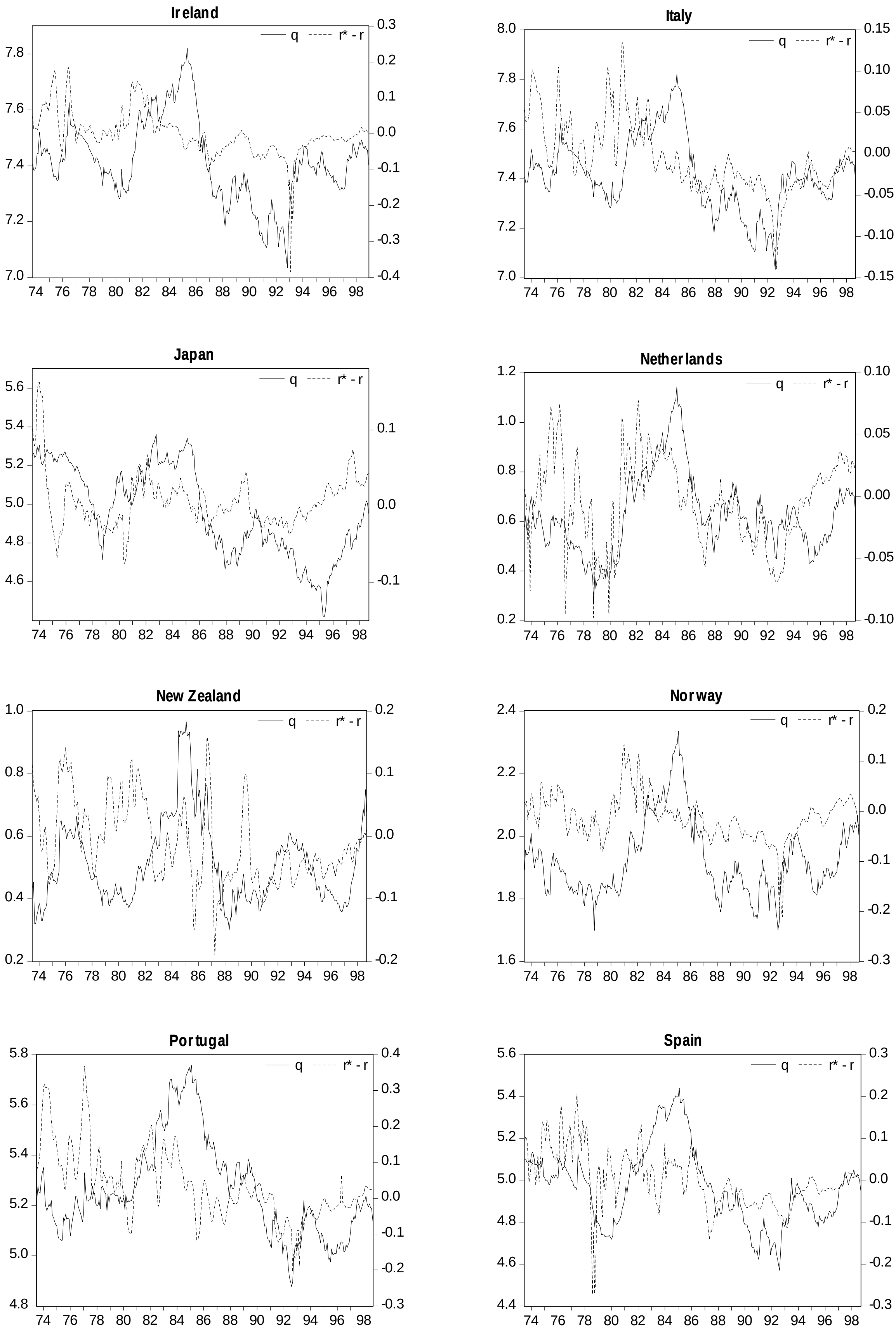
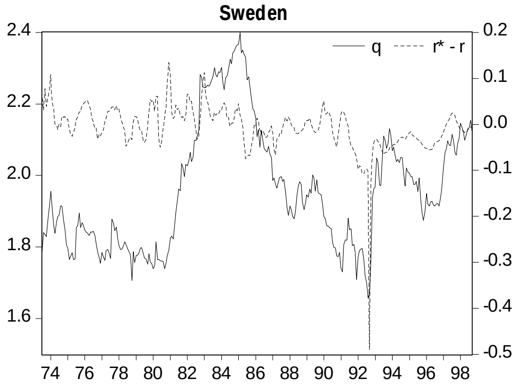
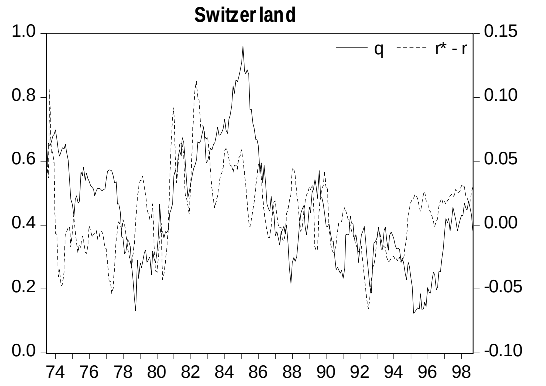

United Kingdom
q
r* - r
0.2
0.0
-0.2
-0.4
-0.6
-0.8
0.12
0.10
0.08
0.06
0.04
0.02
0.00
-0.02
-0.04
84
86
88
90
92
94
96
98

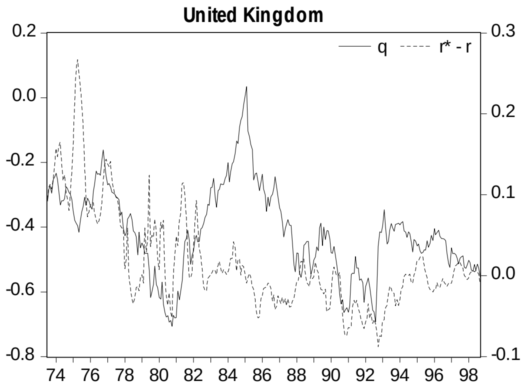
Fig 3 Real exchange rates (left scale) and ex ante short-term yield differentials (right scale)




Ireland
q
r* - r
7.0 7.2 7.4 7.6 7.8
-0.4 -0.3 -0.2 -0.1 0.0 0.1 0.2 0.3
74 76 78 80 82 84 86 88 90 92 94 96 98
Italy
q
r* - r
7.0 7.2 7.4 7.6 7.8 8.0
-0.15 -0.10 -0.05 0.00 0.05 0.10 0.15
74 76 78 80 82 84 86 88 90 92 94 96 98
Japan
q
r* - r
4.6 4.8 5.0 5.2 5.4 5.6
-0.1 0.0 0.1
74 76 78 80 82 84 86 88 90 92 94 96 98
Netherlands
q
r* - r
0.2 0.4 0.6 0.8 1.0 1.2
-0.10 -0.05 0.00 0.05 0.10
74 76 78 80 82 84 86 88 90 92 94 96 98
New Zealand
q
r* - r
0.2 0.4 0.6 0.8 1.0
-0.2 -0.1 0.0 0.1 0.2
74 76 78 80 82 84 86 88 90 92 94 96 98
Norway
q
r* - r
1.6 1.8 2.0 2.2 2.4
-0.3 -0.2 -0.1 0.0 0.1 0.2
74 76 78 80 82 84 86 88 90 92 94 96 98
Portugal
q
r* - r
4.8 5.0 5.2 5.4 5.6 5.8
-0.3 -0.2 -0.1 0.0 0.1 0.2 0.3 0.4
74 76 78 80 82 84 86 88 90 92 94 96 98
Spain
q
r* - r
4.4 4.6 4.8 5.0 5.2 5.4 5.6
-0.3 -0.2 -0.1 0.0 0.1 0.2 0.3
74 76 78 80 82 84 86 88 90 92 94 96 98


Sweden
q
r* - r


Switzerland
q
r* - r


United Kingdom
q
r* - r