Reinterpreting the Real Exchange Rate - Yield Differential Nexus*

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Abstract

This paper proposes a new interpretation of the real exchange rate - real yield differential nexus. Instead of relying on a cointegration framework, it provides estimates of the long run relationship in a panel regression whose residual term may be subject to permanent shocks. 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 findings are interesting since they support fundamentals-based models of exchange rate behaviour but also permit real factors to play a role. Moreover they indicate that capital markets integration is more advanced than hitherto believed.

JEL classification: C23; F31

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1 Introduction

The debate on the real exchange rate - real yield differential nexus in recent years has by and large been dominated by the cointegration approach.[1] On balance the verdict from time-series cointegration tests is 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) but 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 evidence of cointegration for a sample of 11 open economies though not for their full panel of 18 OECD economies. This failure to find clear evidence of cointegration has provided a basis for criticisms of the monetary and portfolio balance models of exchange rate determination and for a defense of the predominant role of real disturbances such as productivity differentials or current account imbalances.

While cointegration studies have shed important light on the underlying issues, they exclude a role for permanent disturbances in a regression of exchange rates on yield differentials. This is because, although non-stationary disturbances may be plausible from a theoretical viewpoint, they lead to inconsistent OLS time-series estimators. One rationale for such disturbances is that the residual term of a regression of exchange rates on yield differentials at a minimum captures the expected future exchange rate. If the typical real macroeconomic fundamentals underlying the latter — relative productivity growth, current account imbalances or GDP differentials — have stochastic trends, the regression error will also be nonstationary.

A second rationale for nonstationary errors relates to the idea that temporal aggregation and nonlinearities can make linear regression residuals appear nonstationary. A.M. Taylor (2001) demonstrates the latter in the context of regressions of spot exchange rates on price differentials.[2] Nakagawa (2002) estimates a nonlinear time-series model along these lines to show that the relationship between real exchange rates and yield differentials holds for large deviations of the real exchange rate from its equilibrium value.[3] This finding

[1] Hereafter these variables are defined in real terms unless otherwise specified.

[2] 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.

[3] This is in line with some positive evidence on cointegration between exchange rates

is consistent with Baxter (1994) who demonstrates the association between the temporary components of exchange rates and relative yields. 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.

This paper revisits the exchange rate-yield differential debate from a new econometric perspective. It is distinctive in that it does not seek to answer the question of whether the two variables cointegrate. Rather the goal is to measure consistently the long run effect of relative yields on exchange rates while permitting the latter to be subject to permanent shocks. In doing so it builds on recent nonstationary panel data studies which show that long-run effects are not exclusively associated with cointegrating relationships (Pesaran and Smith 1995; Phillips and Moon 1999; Kao 1999). These studies demonstrate that by adding cross-section information it is possible to estimate consistently a long run coefficient even in the absence of time series cointegration.[4] 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, Fuertes and Smith (2001) show via Monte Carlo simulations that these asymptotic results are relevant for the panel dimensions typical of post-Bretton Woods studies.

Our empirical analysis for 23 industrialized countries 1973-1998 indicates that yield differentials have a significant long run impact on exchange rates. Since our statistical framework accommodates nonstationary regression errors, these findings do not imply that yield differentials suffice to explain all the observed persistence in exchange rates. In this regard, this contribution may serve to bridge the gap between conflicting results in the literature. Futhermore, the findings are consistent with a high degree of integration of world financial markets. In this respect this study adds to the recent evidence on a high degree of capital mobility (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 differential relation and discusses the statistical framework. Section 3 describes the data and analyses the empirical results. A final section concludes.

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and yield differentials deriving from the enhanced power properties of the tests deployed (Chortareas and Driver 2001; Edison and Melick, 1999).

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

2 The exchange rate-yield differential nexus

2.1 Theoretical framework

Define the current nominal exchange rate level, S_t, as the value that equates the expected return on assets denominated in different currencies

$$(1 + i^*_{t,t+1})\frac{E_t S_{t+1}}{S_t} = (1 + i_{t,t+1})(1 + \lambda_t) \tag{1}$$

where S_t is defined as the domestic currency price of foreign currency, $E_t(\cdot)$ denotes time t expectations, $i_{t,t+1}$ ($i^*_{t,t+1}$) is the time t return on a domestic (foreign), one-period asset and λ_t is an exogenous risk premium reflecting the less-than-perfect substitutability of the foreign and domestic assets. Equation (1) defines risk-adjusted uncovered interest parity (UIP). It embodies the forward-looking asset view of exchange rate determination characteristic of the monetary and portfolio balance approaches. It can be written alternatively as $S_t = \frac{E_t(S_{t+1})}{1-(i_{t,t+1}-i^*_{t,t+1}+\lambda_t)}$ to state that the current exchange rate is the present value of the future rate where the discount factor is the risk-adjusted interest differential.

Using logarithms (denoted by lower case) and ignoring cross terms, the UIP relation can be rewritten as

$$s_t = E_t s_{t+1} - (i_{t,t+1} - i^*_{t,t+1}) - \lambda_t \tag{2}$$

The Fisher equation is used to decompose the nominal interest rate into a real rate and an expected inflation term

$$i_{t,t+1} = r_{t,t+1} + E_t\left[\pi_{t,t+1}\right] \tag{3}$$

where $\pi_{t,t+1}$ is the domestic inflation rate for the period t to $t+1$. Adding an expected inflation differential to both sides of (2) and noting the (log) real exchange rate definition $q_t \equiv s_t - (p_t - p^*_t)$ it follows that

$$q_t = E_t\left[q_{t+1}\right] - (r_{t,t+1} - r^*_{t,t+1}) - \lambda_t \tag{4}$$

The latter equation is known as the real uncovered interest parity (RUIP) relationship. Since the coefficient on the (annualised) yield differential depends on the underlying asset maturity m, this relation can be written more generally as

$$q_t = E_t\left[q_{t+m}\right] + \varphi_m(r_{t,t+m} - r^*_{t,t+m}) - \lambda_t \tag{5}$$

where $\varphi_m < 0$ is an increasing function of the maturity of the interest rate. Equation (5) implies that the current real exchange rate level is a funtion of the expected future rate, a real yield differential and a risk premium.

2.2 An alternative long run approach

The real exchange rate-yield differential literature has been bedevilled by two empirical issues. One is the marked inertia of real exchange rates or their persistent deviations from long-run equilibrium. Apart from overshooting — which rests on the joint hypothesis of price stickiness and a predominance of monetary shocks — possible explanations are 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. Recently, empirical evidence has been adduced showing that the nonlinear relation between nominal exchange rates and the underlying fundamentals — arising because of the interaction between transaction costs and agents' uncertainty about the 'true' equilibrium level — lies at the heart of the slow mean reversion of real exchange rates.

A second issue is that, contra ongoing financial market integration, most empirical studies have found at best ambiguous evidence on real interest parity or the stationarity of real interest rate differentials. 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 virtually indistinguishable from integrated processes in typical finite samples. Moreover, persistent deviations from a constant yield differential have been rationalized for long horizons as arising from the lack of homogeneity or liquidity in government bonds (Meese and Rogoff, 1988) or from relative commodity price movements arising from real shocks to the economy. The upshot is that real exchange rates and yield differentials seem observationally equivalent to nonstationary integrated processes. This has prompted a number of studies during the 1990s to investigate the nexus within a cointegrating framework and accordingly to test whether a linear combination of these variables is stationary.

The workhorse of most cointegration-based studies is equation (5) together with an expectations equation such as

$$E_t(q_{t+m} - \bar{q}_{t+m}) = \phi^m(q_t - \bar{q}_t), 0 < \phi < 1 \tag{6}$$

which embodies monotonic adjustment towards a long run equilibrium $\bar{q}_t$. One problem with some of these studies is that they rest on the assumption of a constant $\bar{q}_t$ — the long-run PPP pillar of most monetary and portfolio balance models — which is at odds with accepting I(1) real exchange rate behavior to conduct cointegration tests and, in turn, with the implicit mean-reverting behavior of $q_t - \bar{q}_t$ in (6). Another strand of research has posited $\bar{q}_t$ as a function of some fundamental variables, most frequently the cumulated current account balance to GDP ratio (Edison and Melick 1999). More recent work suggests that the equilibrium path may be also influenced by productivity differentials, saving-investment decisions or GDP differentials (Lane and Milesi-Ferreti 2000; Hoffmann and MacDonald 2001). The main tenet of these studies is non-stationary equilibrium real exchange rates since the fundamentals themselves may be subject to permanent shocks.

This paper does not seek to unravel the role of the above I(1) fundamental variables or other driving forces behind the expected future exchange rate level, $E_t[q_{t+m}]$. Instead it permits the latter to enter the residual term in our regression framework. The implication is that q_t and $r_{t,t+m} - r^*_{t,t+m}$ are not required to cointegrate. Rewriting (5) in panel regression form (and dropping the subscript m to simplify the notation) we have

$$q_{it} = \alpha_i + \beta_i(r_{it} - r^*_{it}) + u_{it}, i = 1, ..., N, t = 1, ..., T \tag{7}$$

where i is the country (group) index. The error term u_{it} therefore captures the risk premium and the expected future real exchange rate. Some studies are suggestive of a non-zero but constant risk premium while others argue in favour of a time-varying albeit stationary component. In either case, the systematic variability of $E_t[q_{t+m}]$ 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 measure the long run coefficient β if the errors are I(1).

Phillips and Moon (1999) and Kao (1999) argue that some panel datasets offer the prospect of overcoming the spurious regression problem of pure time series. More particularly, they demonstrate 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 i lessens the noise in the relationship — the covariance between the I(1) error and the I(1) regresssor — that induces the spurious

regression problem. Take the simple regression model

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

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 MG or unweighted average panel estimator is defined by

$$\widehat{\beta}^{MG} = N^{-1} \sum_{i=1}^{N} \left(\frac{\sum_{t=1}^{T} \widetilde{y}_{it} \widetilde{x}_{it}}{\sum_{t=1}^{T} \widetilde{x}_{it}^2} \right) \tag{8}$$

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 fixed effects (FE) or weighted average estimator is

$$\hat{\beta}^{FE} = \sum_{i=1}^{N} w_i \left(\frac{\sum_{t=1}^{T} \widetilde{y}_{it} \widetilde{x}_{it}}{\sum_{t=1}^{T} \widetilde{x}_{it}^2} \right) = \frac{\sum_{i=1}^{N} \sum_{t=1}^{T} \tilde{x}_{it} \tilde{y}_{it}}{\sum_{i=1}^{N} \sum_{t=1}^{T} \tilde{x}_{it}^2} = \beta + \frac{\sum_{i=1}^{N} \sum_{t=1}^{T} \tilde{x}_{it} u_{it}}{\sum_{i=1}^{N} \sum_{t=1}^{T} \tilde{x}_{it}^2} \tag{9}$$

where $w_i = S_i / \sum_i S_i$ with $S_i = \sum_t \widetilde{x}_{it}^2$ and $\tilde{x}_{it}$ and $\tilde{y}_{it}$ are defined as above. 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 as T becomes large. However, this problem is alleviated in a panel context by averaging over i and so a consistent estimate of β can be obtained as $N \to \infty$ and $T \to \infty$.[5]

These asymptotic results are complemented by Coakley, Fuertes and Smith (2001) who investigate the small sample properties of two pooled estimators — FE and pooled OLS (POLS) — and the mean group (MG) estimator of Pesaran and Smith (1995) in a non-stationary regression setup.[6] Their Monte Carlo simulations confirm that the limit theory is 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).[7]

[5] 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 effect of I(1) errors and between-group dependence.

[6] Asymptotic results have not been established for the MG estimator but the estimator has been shown to be unbiased and correctly sized in finite samples (typical of PPP studies) in Coakley, Fuertes and Smith (2001).

[7] One contrasting feature between the FE and MG estimators is that the standard errors of the former will be biased in the I(1) error case and inference based on them may be misleading. However, correct standard errors can be computed (Phillips and Moon 1999).

3 Empirical analysis

3.1 Summary statistics

The sample comprises 23 industrialized countries from which the US is chosen as numeraire. 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.[8]

To gauge the sensitivity of the results to the inflation measure we use both *ex post* and *ex ante* real yields calculated from a static expectations assumption, $E_t(\pi_{t,t+m}) = \pi_{t-m,t}$, and two smoothing procedures. The latter involve a 7-point two-sided moving average (MA) filter and a Holt-Winters (HW) filter which averages past and present values and generalizes the single exponential smoother by adding linear trend and seasonal components.[9] 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.[10] The latter contrasts with most existing studies that, due to sample size constraints, deflate LT rates using a long MA smoother of ST (usually three-month-ahead) inflation rates. Nevertheless, to facilitate comparisons with the literature and since the latter seems to work quite well in practice we also deflate the LT yields using a 25-month two-sided MA smoother of three-month-ahead inflation.

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

[8]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.

[9]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).

[10]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 over the term of the bonds may produce real yield differential measures which are closer to the relevant ones.

German short term interest differential.

[Figure 1 around here]

The real exchange rate and ex ante yield differential 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 exchange rate overshooting or may reflect what Obstfeld and Rogoff (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 differential quite well for some countries over particular periods and more so for the long-term securities.

[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 single-equation augmented Dickey-Fuller (ADF) test. Table 1 reports the results for the CPI- and PPI-based real exchange rates and ST and LT yield differentials based on the static expectations and 7-point MA proxies for CPI inflation.[11]

[Table 1 around here]

They indicate that in a majority of cases it is not possible to reject the non-stationarity null for the real exchange rate and yield differential series. The conclusion for the latter may be questioned against a backdrop of highly integrated capital markets. However, most empirical studies — using both standard single-equation and multivariate or panel approaches — fail to find cogent evidence of stationary behavior in LT or ST real yield differentials (Baum and Barkoulas 2002; Chortareas and Driver 2001; Hoffmann and MacDonald 2001; Edison and Melick 1999). On balance therefore the results add to the consensus view that real exchange rate and yield differentials are observationally equivalent to nonstationary processes over the post Bretton-Woods sample period.

[11]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, real yield series constructed using the HW filter and PPI measures, are qualitatively similar and are available on request.

3.2 Long run panel estimates

This study uses both the MG and the FE panel estimators which have been shown to provide consistent measures of a long run coefficient in the context of nonstationary disturbances. These estimators also allow for country heterogeneity to varying degrees. The FE estimator permits heterogeneous intercepts but imposes equality of slopes, $\beta_i = \beta$, in (7). It is computed using (9) 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 (8). 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. For large N panels, heterogeneity may be an important issue and in this regard the MG estimator may be the more appropriate.

Table 2 reports the estimation results for the real exchange rate-yield differential relationship using ST interest rates and CPI inflation.

[Table 2 around here]

The first two columns report the MG and FE long run coefficient estimates for the different inflation proxies considered. Although there is some variation, the estimates are correctly signed and statistically significant in all cases.[12] Countries with individual estimates more than two standard deviations from the mean are trimmed to control for excessive heterogeneity. The resultant estimates reported in the final two columns are closer to their theoretical value and significantly negative.[13]

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

[12] This is inferred from the MG estimates and standard errors. The magnitude of the FE standard errors 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.

[13] 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).

in magnitude — especially for the CPI panels where the FE estimates are approximately half the value of their MG counterparts — which underlines the importance of country heterogeneity. The hypothesis that the slope coefficient is equal to its theoretical value — at φ_m = -0.25 for three month maturity and annualized yield rates — is however rejected. This may stem from the 'noise' in the inflation measure or possibly from omitted variable bias.

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

[Table 3 around here]

Since the static expectations and ex post inflation proxies produce very close results, only those for the former are reported. The table also contains the case of LT yields deflated by a 25-point MA filter of 3-month-ahead inflation to compare our results with those in the literature. The coefficient estimates for the LT yield differentials differ from those in Table 2 in two respects. First, in line with the theoretical priors, their absolute value is up to twelve times larger than that in the analogous ST case indicating a term structure relationship in yield differentials. Second, the FE estimates are much closer to the MG estimates suggesting that slope coefficient heterogeneity is mitigated for the LT yields.[14] The PPI panels give qualitatively similar results.

While the supporting evidence for the LT yield differential case is in line with some existing studies such as that of Chortareas and Driver (2001), the significant long run relationship found between real exchange rates and ST yield differentials represents a novel finding in this context.[15] More generally, our findings 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.[16]

[14]We also computed real yield differentials using a 7-year-ahead inflation measure since the Macaulay duration of the observed 10-year, coupon-paying bonds roughly corresponds to that for 7-year pure discount bonds, but this makes little difference 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.

[15]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 in favour of the expectations hypothesis at the long end than there is at the short end (Campbell and Shiller 1991).

[16]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), re-

4 Conclusions

This paper revisits the real exchange rate-yield differential parity relationship linking international asset and commodity markets. Prior work formulated largely within the cointegration framework has fallen short of establishing consistent evidence of a long run nexus between these variables. The alternative approach proposed in this paper to assessing the exchange rate-yield differential link allows us to go beyond a statistical cointegration relationship. Rather it permits nonstationary regression errors to accommodate real factors having an impact on the real exchange rate. In so doing it builds on recent advances in nonstationary panel data theory which demonstrate that long-run effects are not exclusively associated with cointegrating relationships. This rapidly emerging area of theoretical research proffers an opportunity directly to uncover evidence of long-run associations which hitherto have proven elusive.

A panel dataset for 23 industrialized countries 1973M1-1998M12 yields significantly negative slope coefficient estimates irrespective of the inflation proxy used. A comparison across panels using short- and long-term bonds reveals a clear term structure relationship. While the supporting evidence for the LT yield differential case is in line with the findings of Chortareas and Driver (2001), 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 as well as long term international capital flows We conclude that the data provide support for sticky-price theories of exchange rate determination without precluding real factors from playing a role in the persistence and volatility of real exchange rates. In this regard, our contribution may serve to bridge a gap between conflicting results in the literature. 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 this statistical framework to a heterogeneous dynamic panel regression which incorporates short-run movements also and then to reexamine the evidence on the long-run nexus.

spectively, for the ST yields in the naive expectations case after discarding Italy as an outlier.

Appendix A: Data Sources

The data cover the period 1973M1 to 1998M12. End-of-month bilateral exchange rates via-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 OECD or IMF (line 61) series on 10-year bellwether government bond yields.

	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 the real exchange rates, T=309 (static expectations) and T=306 (MA) for the.ST interest rate differentials and T=192 (static) and T=186 (MA) for the LT differentials. Real interest rate differentials are based on CPIs. *significant at the 5 % level. **significant at the 1% level.

Table 2 Slope coefficient estimates for short term yield differentials

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

[a]Outlier countries whose individual estimates are more than two standard deviations away from the mean.

Table 3 Slope coefficient estimate for long term yield differentials

	Full panel					Outliers excluded	
	$\hat{\beta}^{MG}$	$\hat{\beta}^{FE}$	$\text{Min}_{\hat{\beta}_i}$	$\text{Max}_{\hat{\beta}_i}$	$Z > 2$	$\hat{\beta}^{MG}$	$\hat{\beta}^{FE}$
Static[a]	-3.68	-3.48	-5.28	-.46	2.54	-3.87	-3.60
	(.299)	(.076)	(OE)	(CN)	(CN)	(.246)	(.078)
MA_7	-4.95	-4.43	-7.51	-.828	2.20	-5.19	-4.56
	(.441)	(.087)	(LX)	(CN)	(CN)	(.391)	(.083)
HW	-4.01	-3.58	-5.95	-.684	2.15	-4.20	-3.69
	(.364)	(.080)	(LX)	(CN)	(CN)	(.326)	(.082)
MA_{25}^{b}	-3.15	-2.83	-8.03	-.019	-2.52	-2.82	-2.74
	(.484)	(.081)	(FR)	(CN)	(FR)	(.383)	(.082)

Notes: The effective sample period for the cases based on 10-year-ahead inflation measures is shorter. This is 1985M1-1998M12 for the static and HW cases and 1985M4-1998M9 for the MA_7 case. [b]Three-month-ahead inflation is used. The effective sample is 1974M4-1997M12 and hence this case excludes Portugal and Spain since their yields are observed post-1975M12 and 1977M12, respectively.

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

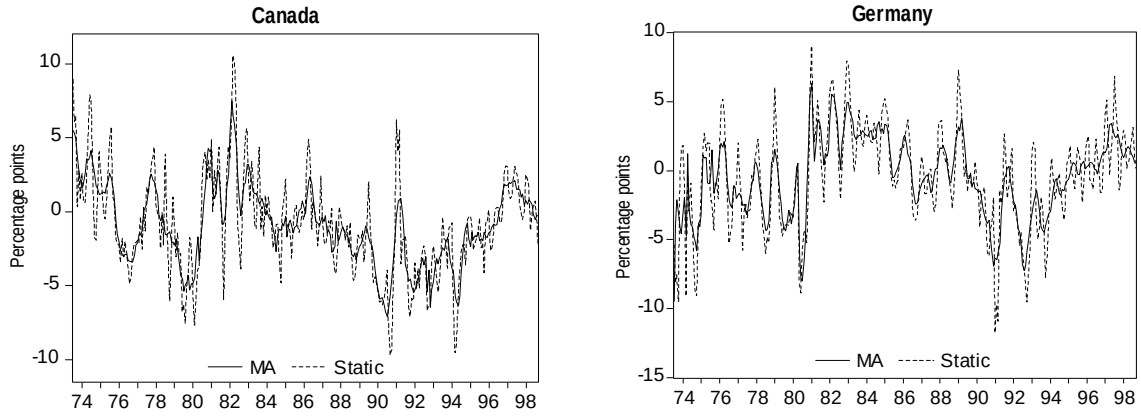
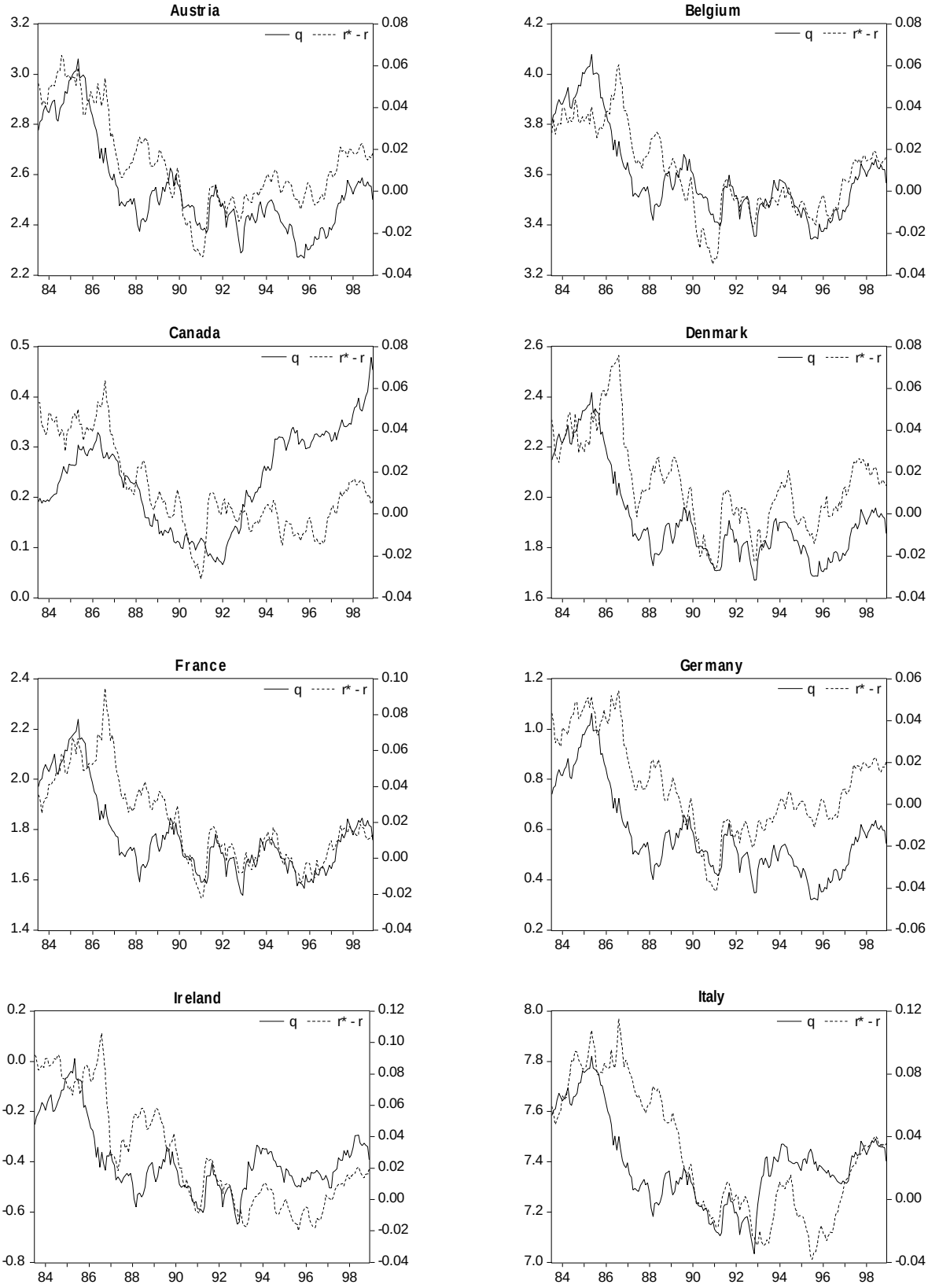
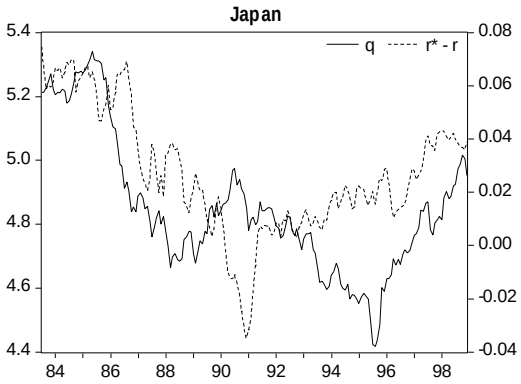


Figure 2 Real exchange rates (left scale) and ex ante long-term yield differentials (right scale)




Japan
q
r* - r
5.4
5.2
5.0
4.8
4.6
4.4
0.08
0.06
0.04
0.02
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-0.02
-0.04
84
86
88
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92
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96
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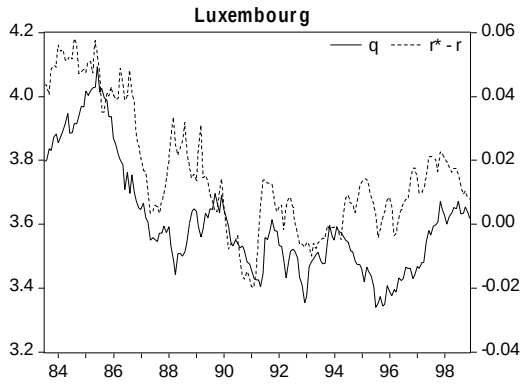
Luxembourg
q
r* - r
4.2
4.0
3.8
3.6
3.4
3.2
0.06
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-0.02
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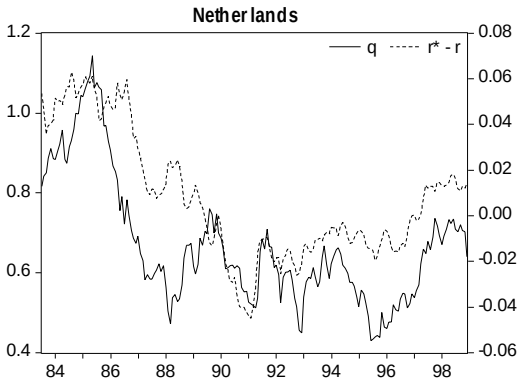
Netherlands
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r* - r
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0.8
0.6
0.4
0.08
0.06
0.04
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-0.04
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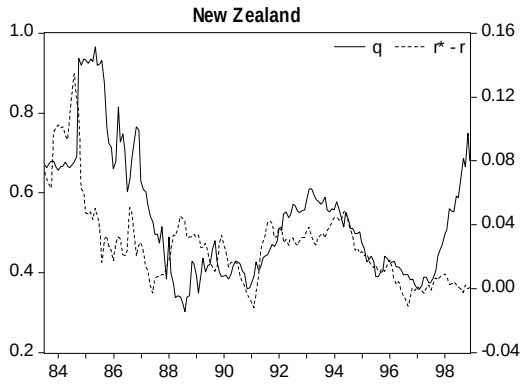
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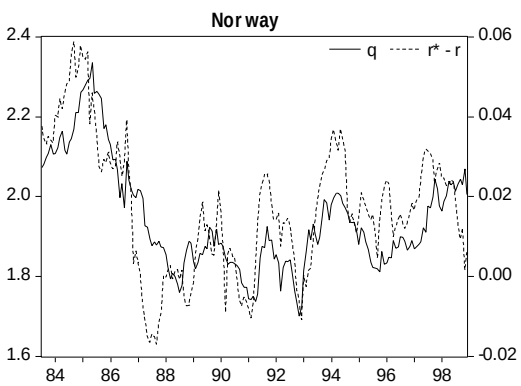
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r* - r
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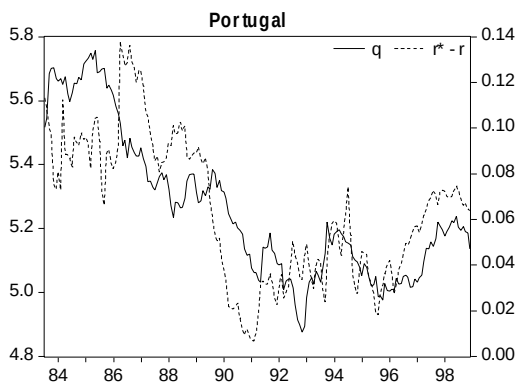
Portugal
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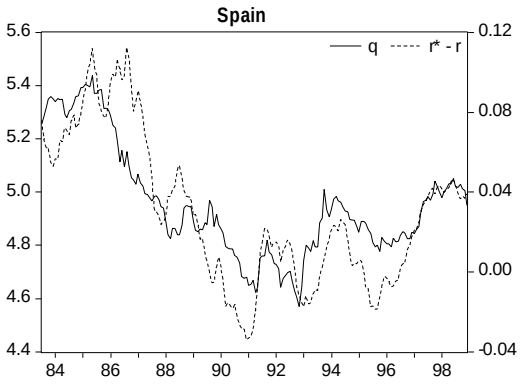
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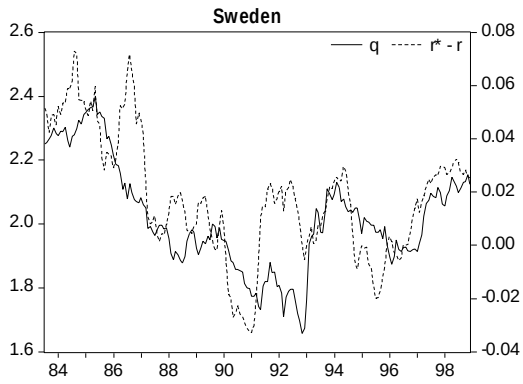
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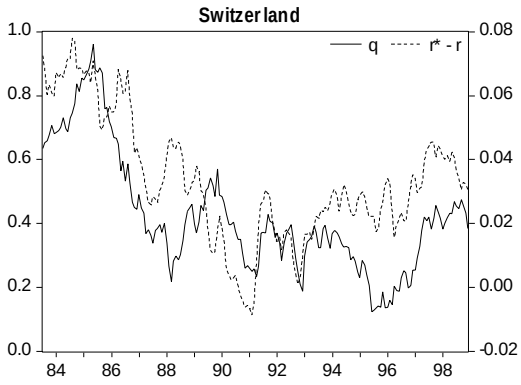
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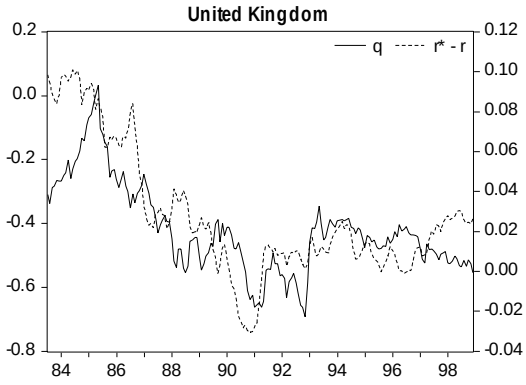
United Kingdom
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0.06
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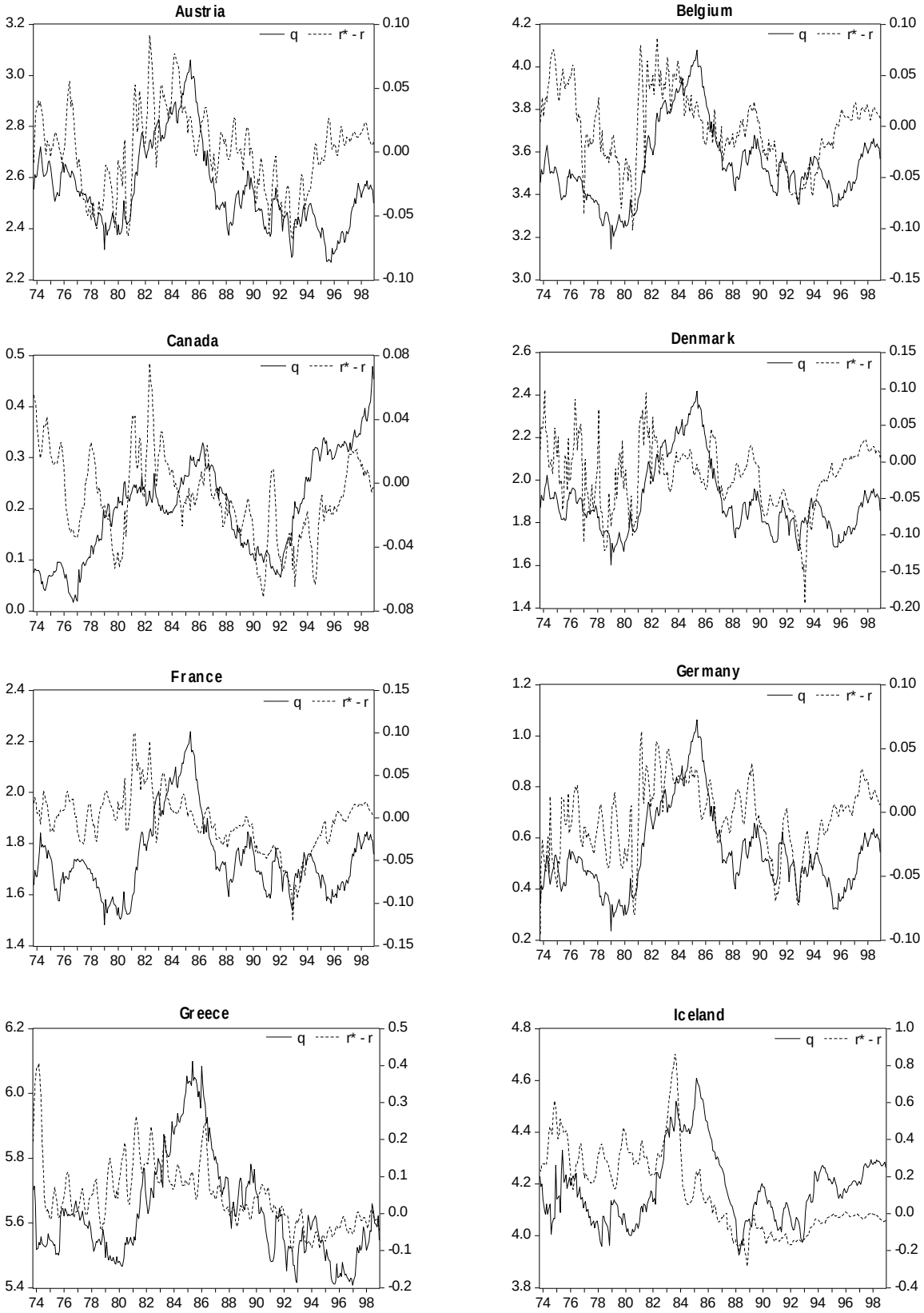
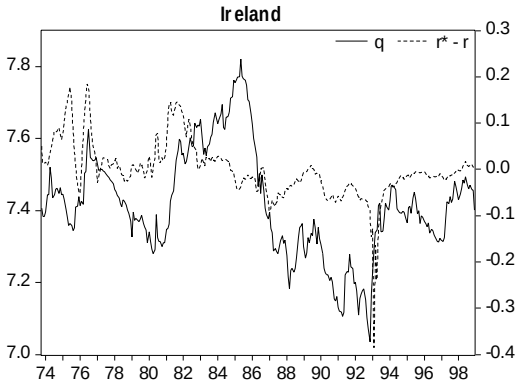
Fig 3 Real exchange rates (left scale) and ex ante short-term yield differentials (right scale)

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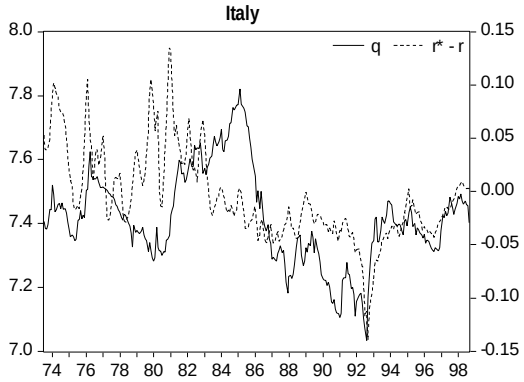
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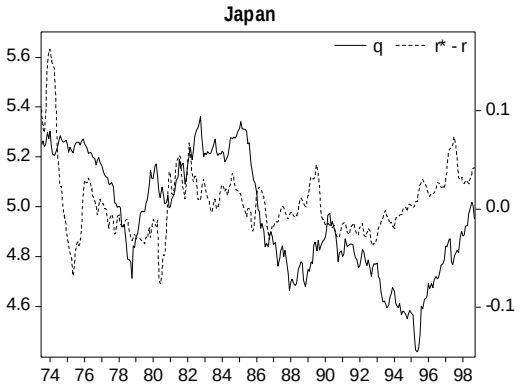
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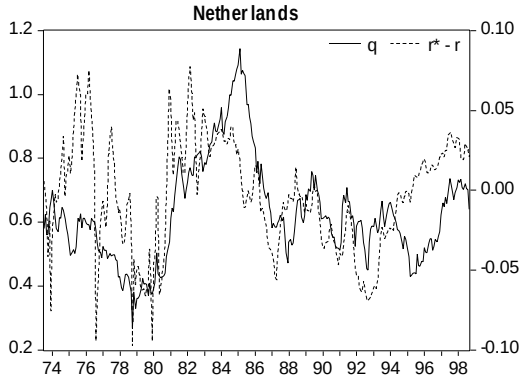
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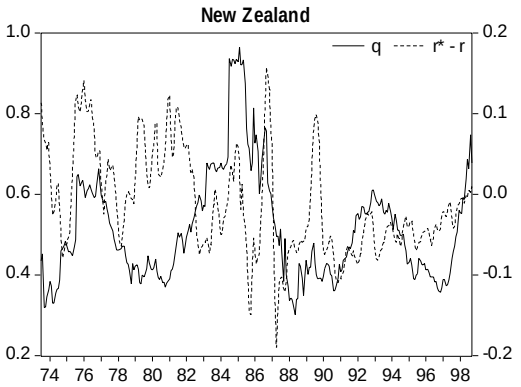
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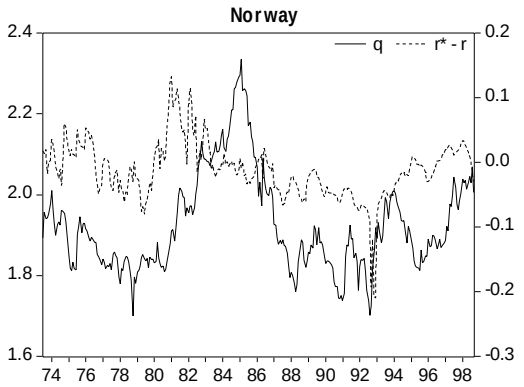
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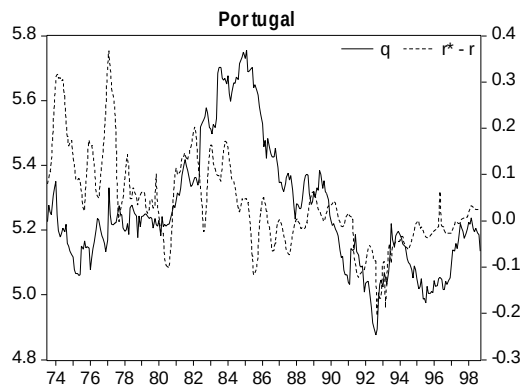
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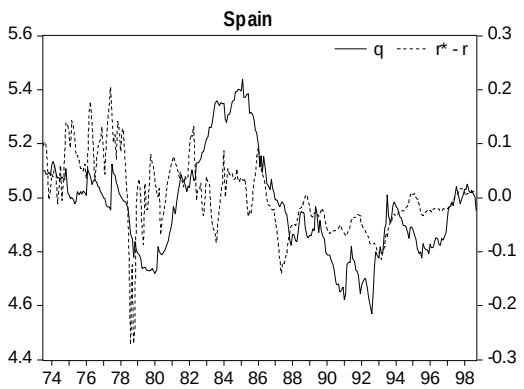
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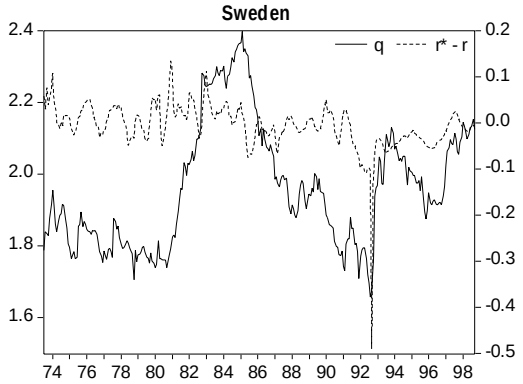
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r* - r
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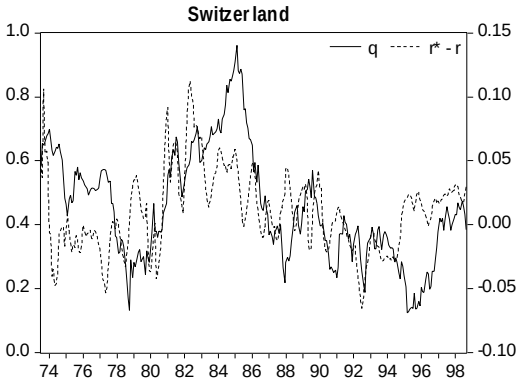
Switzerland
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r* - r
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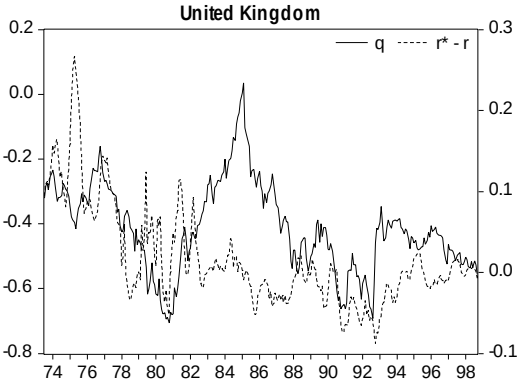
-0.10
74 76 78 80 82 84 86 88 90 92 94 96 98


United Kingdom
q
r* - r
0.2
0.0
-0.2
-0.4
-0.6
-0.8
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74 76 78 80 82 84 86 88 90 92 94 96 98