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FIN DE SIÈCLE REAL INTEREST PARITY

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ABSTRACT

We evaluate the recent evidence for real interest parity, focusing on long-term yields. Examining the data on financial instruments of various maturities across the G7 countries, we find substantial differences in the degree of real interest equalization measured at different horizons. In general, real interest parity holds better at long horizons than at short. This empirical result is robust to alternative ways of modeling expected inflation rates. Considering the relevance of long-term yields for the investment decisions of firms, our findings imply that the degree of capital mobility among the G-7 economies may be greater than previously thought.

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1. Introduction

Over a decade ago, in his survey of international capital mobility, Frankel concluded:

> "...a currency premium remains, consisting of an exchange risk premium plus expected real currency depreciation. This means that, even with the equalization of covered interest rates, large differentials in real interest rates remain." (Frankel, 1989: 253)

The persistence of real interest differentials, even among the developed economies, suggested that capital mobility had yet to reach the stage where rates of return in terms of physical goods are equalized across national borders.

Given the omnipresent view that "globalization" has swept away many of the barriers to the free flow of goods as well as capital, it seems appropriate to investigate the validity of these conclusions at the turn of the new century. In this study we evaluate the recent evidence for equalization of real interest rates among the G7 economies, using data on yields of instruments with various maturities. Both ex ante and ex post real interest rate movements are examined at short (up to one-year) and long (five- and ten-year) horizons.

To anticipate the results, we find that there are substantial differences in the degree of capital mobility measured at different horizons. As in numerous previous studies (Cumby and Obstfeld, 1984; Cumby and Mishkin, 1986; Mishkin, 1984; Mark, 1985; Taylor, 1991), the real interest parity (RIP) hypothesis is decisively rejected with short horizon data. At five to ten-year horizons, however, the empirical evidence becomes far more supportive and in some cases the RIP hypothesis is not rejected. In general, RIP, up to a constant, holds better at long horizons than at short.[1] These results are robust to alternative ways of modeling expected inflation rates.

The more positive results that accompany the use of yields on long-term debt instruments are not without cost. These instruments are more heterogeneous than the offshore deposit rates that have typically been used in analyses of capital mobility. Moreover, it is not appropriate to characterize long-term bonds as zero discount bonds, so the reported interest rate data provide only approximate measures of the true returns that investors obtain. (A third issue is that ex post real interest rates are only measurable when the prices have been realized; hence the 2000Q1 ex post 10 year real interest rate will not be observed until 2010Q1.) Yet, in many ways, these long-term instruments are more appropriate for testing capital mobility. First, firms do not usually make their investment decisions on the basis of short-term yields; in fact, depending upon the market structure of the economy, firms may rely on bank debt or equity. However, to the extent that firms borrow in bond markets, long-term bond yields will be the most informative series. Second, also from the investors' point of view, the long-term real rates are most relevant since they more closely measure rates of return expressed in terms of physical goods. Finally, if our aim is to assess the equalization of returns in differing political jurisdictions, then on-shore -- rather than off-shore -- rates are once again more appropriate.

The remainder of this study is organized in the following manner. Section 2 provides a theoretical discussion. Section 3 describes the data and presents the results of the preliminary analysis on the ex post real interest rate behavior. In section 4 we examine how ex ante real interest rates have moved over the past two decades among the G-7 economies at various horizons. Section

[1] Since the government bonds examined in this study are not necessarily of identical default characteristics, RIP may not be observed with exactly zero real interest differentials. Also, small
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5 concludes.

2. Financial Integration and Real Interest Rates

Financial market integration refers to the ease with which assets are traded across borders and currency denominations. A decomposition of the nominal interest differential on instruments of comparable default attributes is helpful at this point.

$$\begin{aligned} i_t^k - i_t^{k*} &\equiv [i_t^k - i_t^{k*} - fd_t^k] + [fd_t^k - (s_{t,t+k}^e - s_t)] + (s_{t,t+k}^e - s_t) \\ fd_t^k &\equiv f_t^k - s_t \end{aligned} \tag{1}$$

where i_t^k is the nominal interest rate for a domestic debt instrument of maturity k, i_t^{k*} is the foreign counterpart, s_t is the (log) spot exchange rate at time t, f_t^k is the (log) forward exchange rate at time t (in the US dollar/foreign currency units) for a trade at time $t+k$, and $s_{t,t+k}^e$ is the (log) spot exchange rate expected for period $t+k$, as of time t. The term fd_t^k is the forward discount, and ($s_{t,t+k}^e - s_t$), the rate of expected depreciation. The first two terms on the right hand side of (1) are referred to as a covered interest differential and an exchange risk premium, respectively.

The existence of a covered interest differential is often taken as a manifestation of "political risk", caused either by capital controls, or the threat of their imposition. In the absence of these barriers, such differentials should not exist because they imply unlimited arbitrage profit opportunities. Frankel (1989) terms the condition of a zero covered interest differential "perfect capital mobility."

The absence of an exchange risk premium constitutes "perfect capital substitutability".

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deviations of the constant term from zero are expected due to Jensen's inequality.

This condition arises when government bonds, denominated in differing currencies, are treated as perfect substitutes. Investors will act this either when they are risk neutral, or when government bonds are actually identical in all important aspects.

A plethora of studies too numerous to mention have examined both issues for the G-7 countries. It is generally found that the covered interest differential can essentially be ignored from the 1980s onward for most of the countries in the sample. One exception is Canada; onshore interest rates apparently deviated from covered interest parity (Chinn and Frankel, 1994), due to uncertainty regarding the prospects for Canadian federalism. However, for other countries the differential is essentially zero.

While financial capital apparently moves with ease to locations where the rate of return is highest, it is not so clear that movements of capital are sufficient to equalize real rates of return. To see that this is a more stringent requirement, consider the situation where uncovered interest parity (UIP) holds,

$$\Delta s^{e}_{t,t+k} = (i^{k}_{t} - i^{k*}_{t}) \tag{2}$$

Suppose further that goods prices are also equalized, up to a constant. In particular, assume relative purchasing power parity (PPP) holds in expectation:

$$\Delta s^{e}_{t,t+k} = (\pi^{e}_{t,t+k} - \pi^{e*}_{t,t+k}) \tag{3}$$

so that expected depreciation equals the expected inflation differential. Equating (2) and (3), and rearranging yields:

$$(i^{k}_{t} - \pi^{e}_{t,t+k}) = (i^{k*}_{t} - \pi^{e*}_{t,t+k}) \tag{4}$$

Equation (4), RIP, states that ex ante real interest rates should be equalized, or alternatively the difference between the two expected real interest rates should be zero. This is the definition of capital mobility we adopt in this study. Note that RIP involves the conditions in both financial and goods markets, and can be interpreted as equalization of expected rates of return in terms of physical goods. In the subsequent sections, we evaluate the recent evidence of capital mobility among the G-7 economies by the degree to which the equalization is attained.

Although one does not observe the expected real interest rates, they can be approximated in a variety of manners in empirical analyses. The first is to use the unbiasedness hypothesis again, and calculate ex post real interest differentials. The second is to model inflationary expectations as a time series process. Most studies have adopted the former approach. Cumby and Mishkin (1986), Blundell-Wignall and Browne (1991) and Taylor (1991) tested equation (4) by regressing one ex post real interest rate upon the other:[2]

$$r_t^k = \mu + \lambda r_t^{k*} + \zeta_{t+k} \tag{5}$$

Cumby and Mishkin (1986) examine monthly data on three month offshore rates for eight industrialized countries over the1973:06-1983:12 period, and generally reject the hypothesis that $\lambda = 1$. Blundell-Wignall and Browne (1991) use onshore rates to test equation (5) over data extending up to the second quarter of 1990. They find that RIP is again often rejected, but that linkage ($\lambda = 1$) is not rejected in four cases: Italy, UK, Netherlands and Switzerland against the US. Taylor (1991) tested for a similar relationship for EMS countries and found that linkage was

[2] An alternative approach to assessing the joint uncovered interest parity/relative PPP condition is to regress inflation differentials on interest differentials. This approach imposes the Fisher hypothesis, which requires instantaneous incorporation of inflation into interest rates. To the extent that this condition does not hold, this approach is more likely to reject. See, for instance, Cumby and Obstfeld (1984) and Mark (1985).

rejected even for intra-EMS country pairs. The results of those previous studies suggest that the evidence for RIP has been rather limited when short horizon data are used. In the following sections we investigate if the same conclusion is obtained when the yields on longer maturity instruments are considered.

3. Data and Preliminary Analysis

3.1 Data

The yields of financial instruments with various maturities are considered for the G7 countries. The short-term interest rates we examine are the 3-, 6-, 12-month maturity eurocurrency yields. There are two sets of long-horizon interest rate data. The first is the end-of-month yields on outstanding government bonds with ten-year maturity at the date of issuance, used by Edison and Pauls (1993). The second data set, available only for selected countries, consist of the synthetic "constant maturity" five- and ten-year yields that are interpolated from the yield curve of outstanding government securities. For the price series, we use both the consumer price index (CPI) and the wholesale price index (WPI) provided by IMF's *International Financial Statistics* (IFS).[3] All data are at a quarterly frequency. The short-horizon interest rate data are generally available for 1976Q1-2000Q1. The benchmark sample period for the long-horizon interest data is 1973Q1-2000Q1. After allowing the maximum of 10-year horizon, that is allowing k to be 40, the available estimation period is 1973Q1-1990Q1. In some cases the sample period is constrained further by limitation of the interest rate data (see the Data Appendix for details).

[3] Due to the limitation of the WPI data, only CPI is used for France and Italy.

3.2 Realized Real Interest Rates

As a preliminary analysis, we examine the ex post or realized real interest differentials (RRID) between the U.S. and the remaining G7 countries:

$$RRID_{t,t+k} \equiv (i_t^k - \pi_{t,t+k}) - (i_t^{k*} - \pi^*_{t,t+k}) \qquad (6)$$

Panels A and B of Table 1 report the mean and standard deviation of the RRID in annualized percentage terms at short and long horizons, respectively. Note that, except for a three-month horizon, using the quarterly frequency data leads to (k-1) periods of overlapping observations, thus, the longer-horizon series are artificially smoothed and obtain smaller standard deviations. To see if the mean of the RRID is significantly different from zero, we also report the corrected standard errors (Hansen, 1982). When the CPI is used to measure the price levels, the RRID of Canada, France, and Italy are significantly different from zero up to a one-year horizon. In other words, on average realized real interest rates are not equalized at short horizons. At five to ten-year horizons, however, the mean RRID for Canada becomes undistinguishable from zero. For France and Italy, the RRID diminish in absolute values at a ten-year horizon although still significantly negative on average. Similar to Canada, Japan's mean RRID becomes insignificant at longer horizons. In general, over five- to ten-year horizons the real interest rates of Canada and Japan are *ex post* equalized with that of the U.S.[4] The same pattern does not apply, however, to Germany and the U.K. The RRID between these two countries and the U.S. are statistically equal to zero for all horizons. When the CPI is replaced by the WPI as the price series, the RRID generally becomes far

4 The corrected standard errors become larger at longer-horizons for Germany and Japan, but not for Canada and the U.K. Hence, at least for the latter two countries, the insignificance of the mean RRID is not a mere consequence of insufficient power due to large standard errors.

more volatile as indicated by the increased standard deviations. Consequently in all cases but the U.K. at a five-year horizon, the mean of the RRID is not statistically distinguishable from zero.

To further examine the relationship between the realized real interests, we estimate:

$$i_t^k - \pi_{t,t+k} = \alpha + \beta(i_t^{k*} - \pi_{t,t+k}^*) + \varepsilon_{t+k} \tag{7}$$

by using the Generalized Method of Moments (GMM) estimator of Hansen (1982) to correct the standard errors for MA(k-1) terms in the residuals.[5] The results with the CPI and the WPI are summarized in Tables 2 and 3, respectively. Panel A of each table contains the short horizon results, and Panel B the long horizon ones. The CPI results, given in Table 2, exhibit a remarkably clear pattern. The value of the slope coefficient estimate increases and becomes closer to unity as the maturity is extended. At five-year and ten-year horizons, the slope coefficient estimates are not statistically different from the theoretical value of unity in most cases. Furthermore, there are five cases (with Canada, Germany, and the U.K.) in which we fail to reject the joint hypothesis of α=0 and β=1, indicating that the realized real interest rates were equal in these cases. The results with the WPI in Table 3 also exhibit a similar tendency of the increasing slope coefficient estimates toward unity as the maturity extends. In no case is the joint hypothesis of α=0 and β=1 rejected at the five-year horizon. These findings suggest that in general *ex post* real interest rates tend to be equalized at long-horizons but not at short-horizons. Nevertheless, it should be emphasized that

[5] We pre-tested the ex post real interest rates using the ADF-GLS test of Elliott, Rothenberg and Stock (1996). For the short-horizon data the unit root hypothesis is rejected unanimously. For the long-horizon data, however, the sample period is substantially shorter, and consequently the test often fails to reject. Yet some of the long-horizon series also fail to reject the null hypothesis of stationarity against a unit root by the LM test of Kwiatkowski, Phillips, Schmidt and Shin (1992). While the test results do not provide unambiguous evidence, it is unlikely for real rates of return to possess unbounded mean and variance. Thus, we treat all ex post real interest rates as stationary
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the findings are based on the realized values, and do not necessarily inform us on whether *expected* real interest rates are equal. In the following sections, we investigate whether the ex ante real interest rates show a similar tendency.

4. Real Interest Parity at Short and Long Horizons

In a simple linear regression framework, one may wish to test equalization of the expected real interest rates by estimating

$$i_t^k - \pi_{t,t+k}^e = \alpha + \beta(i_t^{k*} - \pi_{t,t+k}^{e*}) + \varepsilon_{t+k} \qquad (8)$$

and testing if α=0 and β=1. An obvious obstacle, however, is that the expected inflation rates are not observable, and thus, (8) cannot be directly estimated. Consequently, testing equalization of the ex ante real interest rates requires some additional assumptions regarding how inflationary expectations are formed. A commonly employed method is to utilize the actual inflation data while imposing the rational expectations (i.e. unbiasedness) hypothesis. Another is to model the inflation as a time series process, and use the forecasted series. We pursue each method below.

4.1 RIP under Rational Expectations

Assume that inflationary expectations are rationally formed so that:

$$\pi_{t,t+k} \equiv \pi_{t,t+k}^e + \xi_{t+k}, \quad \pi_{t,t+k}^* \equiv \pi_{t,t+k}^{e*} + \xi_{t+k}^* \qquad (9)$$

where $\xi_{t,t+k}$ and $\xi^*_{t,t+k}$ are the k-period ahead rational inflation forecast errors that are uncorrelated with any time t information. Substituting (9) into (8) obtains:

$$i_t^k - \pi_{t,t+k} = \alpha + \beta(i_t^{k*} - \pi_{t,t+k}^*) + \omega_{t+k} \qquad (10)$$

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series. These results, as well as those for inflation, are available upon request.

where $\omega_{t+k} \equiv \beta\xi^*_{t+k} - \xi_{t+k}$. The condition α=0 and β=1 in (10) means that the expected real interest rates are equal assuming that expectations are unbiased. Rejection of the condition can arise when either expected real interest rates are not equal or expectations are not unbiased, or both.

We estimate (10) by the GMM with the lagged three-month real interest rates at *t*-*k*-3 through *t*-*k* as the instruments.[6] The results are summarized in Table 4 for CPI real rates and Table 5 for WPI rates, respectively. The CPI results indicate that over short-horizons α is generally significantly different from zero while β is not. This suggests that the short-horizon ex ante real interest rates of the U.S. have no relationship with those of the other G-7 countries. With a five-year horizon, however, the results change dramatically. For Canada and the U.K., the estimates of β, reflecting the comovement of the expected real interest rates, are significantly positive and close to unity. On the other hand, the estimates of α, measuring the expected real interest rate differentials, are not significantly different from zero. In fact we fail to reject the composite hypothesis of RIP and rational expectations for the U.K. at a five-year horizon. The ten-year horizon results are somewhat less supportive of RIP. The constants are significantly different from zero except for the U.K., and the slope coefficients deviate from unity, although they are generally significantly positive. In Figure 1, the estimates of β for Germany and the U.K. are graphed by horizons. The stark contrast between the short and long horizon RIP results are well summarized by the figure.

The results with the WPI in Table 5 also exhibit a quite similar pattern to those with CPI

[6] Note that ordinary least squares (OLS) estimates of (10) will be biased since the regressor, the realized real interest rate, measures the expected real interest rate with an error as (9) indicates. Also, for the GMM estimation, the regressor in the current period should not be used as an instrument due to the correlation between $\pi^*_{t,t+k}$ and $\xi^*_{t,t+k}$. To avoid biased estimates, proper
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except that no supportive evidence of RIP is found for Germany, regardless of the horizons. Again, at a five-year horizon the composite RIP/rational expectations hypothesis is not rejected for Canada. In summary, independent of the choice of the price data, we find a rather unambiguous pattern in the test results that evidence for ex ante RIP becomes stronger at long horizons. This is a novel finding. While numerous previous studies rejected RIP, they are based almost exclusively on short horizon data. An important exception to this is Jorion (1996) who rejects RIP at 3-month to 5-year horizons using monthly data for the U.S., Germany, and the U.K. for 1973-1991. In obtaining the contrasting results, our study differs from Jorion (1996) in several important aspects. First, the data sets used in the two studies differ. Perhaps most importantly, our data has a longer sample period extending to the first quarter of 2000. Also, our long-horizon analyses incorporate ten-year horizon data in addition to five-year. Second, we test RIP by examining if two expected real interest rates have a tendency to move exactly one for one and if their difference is on average null. On the other hand, Jorion (1996) imposes one-for-one comovement between two expected real interest rates and investigates if their differentials are systematically related to currently available information. By finding that the current nominal interest differentials contain significant information about future real interest differentials, Jorion (1996) rejects RIP under rational expectations. While closely related, the two methods also use alternative instruments in the estimation procedure, and hence, the results need not be identical. Finally, we conduct a number of robustness checks, including allowing for alternative methods of modeling inflationary expectations.

4.2 RIP with Forecasted Inflation Rates

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instruments need to be lagged at least by *k* periods.

In the absence of the data on expected inflation rates, forecasted inflation series are often used as a proxy variable. We use the univariate time series forecast of inflation as a proxy for the unobserved expected inflation rates. Specifically, the actual quarterly inflation series are modeled as autoregressive (AR) processes.[7] The maximum order of the AR structure is set twelve, and the Schwarz-Bayesian criteria (SBC) is used in selecting the model specification. Once the models are selected, the expected inflation series are constructed by performing rolling regression and forecast exercises. We conduct fixed sample period estimations and out-of-sample forecast with the originating sample period being 1957Q2-1983Q1 (104 observations). With each estimation, the next three-month to ten-year inflation rates are forecasted. As we roll through each forecast period, the parameter estimates are updated with the addition of each new data point. One advantage of the out-of-sample forecast is that it will allow us to construct the expected inflation rates, both short and long, all through 2000Q1. Therefore, although 1973Q1-1983Q1 observations are subsumed into the originating sample period, there will be no loss in terms of sample size as we are able to estimate the expected real interest rates at 1983Q1 through 2000Q1.[8]

Table 6 presents the selected AR specifications. Using these model specifications, the future inflation rates are forecasted at short and long horizons and substituted into (8) as the expected inflation. The estimation results of (8) with the forecasted CPI and WPI inflation series are provided in Tables 7 and 8, respectively. The use of the forecasted CPI inflation rates yields a less clear distinction between the short and long horizon results. The significantly positive values

[7] The ADF-GLS tests reject the unit root hypothesis for the inflation series. The unit root test results (not reported) are available upon request.

8 Note that RIP estimations in sections 4.1 and 4.2 have different sample periods, and hence their results can differ.

of the slope coefficient estimates are found at short as well as long horizons. However, these estimates of β are also significantly different from unity in all cases, and the joint hypothesis of α=0 and β=1 is universally rejected. Further, there appears no consensus across countries in the relationship between the size of the point estimates of β and the maturity.

More informative observations are found in the results with the forecasted WPI inflation summarized by Table 8. As in the CPI case, at a short horizon the slope coefficient estimates are far below unity though significantly positive except for Germany as seen in Panel A of the table. Further α is significantly different from null, indicating the expected real interest rates remain unequal. Consequently the joint hypothesis of α=0 and β=1 is rejected unanimously. At a long horizon, however, the RIP regression obtains remarkable results particularly for Germany and Japan. For Germany, all of the long maturity estimates yield a slope coefficient value statistically equal to unity. Also, the constant term is insignificant except for the synthetic ten-year rate. Consequently, in two out of three cases the joint hypothesis of α=0 and β=1 is not rejected for Germany at the conventional level of statistical significance. The results should be contrasted with those of the short-horizon estimates in which the RIP hypothesis is decisively rejected. As in the German case, the coefficient estimates for Japan are also statistically indistinguishable from the theoretically implied value of unity, and the estimated constants are not statistically significant. The hypotheses of α=0 and β=1, although rejected jointly, are not rejected individually. To highlight the difference between the short and long horizon results, the point estimates of β for Germany and Japan are graphed in Figure 2. Also note that supportive evidence for long horizon RIP is provided also by the synthetic five-year maturity data for the U.K. which does not reject the joint hypotheses of $\alpha=0$ and $\beta=1$.

While CPI and WPI results discussed above are rather different, it is not uncommon to obtain dissimilar results when using alternative price deflators.[9] When deflating yields, the CPI and the WPI generate conceptually different real rates of return. From a firm's viewpoint, the WPI may be more appropriate since the WPI is more likely to measure the price of the firm's output. From an investor's point of view, on the other hand, a better price index might be the CPI, as it better measures the price of a consumption bundle.

We also believe that the dissimilarity in the empirical results is due partly to the time series structure of the price series. The German and Japanese WPI -- deflators that yield positive evidence in support of RIP -- are adequately modeled as AR(1) processes, which makes inflation forecasting fairly straightforward.

Although the RIP results with the forecasted inflation are generally not as clear-cut as those with the unbiasedness hypothesis in the previous section, we still find a few cases where the short and long horizon estimates are drastically different. In these cases, the evidence for the RIP hypothesis is once again much stronger at a long horizon.

5. Conclusions

We have re-evaluated the evidence regarding capital mobility by examining equalization of real interest rates across the G7 countries. The definition we adopt is, admittedly, quite specific. It is a definition at once broader than financial capital mobility, and narrower than the unhindered

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9 See, for instance, Cumby and Obstfeld (1984).

flow of saving and physical investment over borders. We believe such a definition is in some ways the most important for economic behavior. First, onshore rates embody the political risk that is important for firm decisions regarding investment. Second, long-term rates are more directly linked to the rates at which firms borrow from the capital markets, and increasingly, as restrictions on government bond transactions are eliminated, the rates at which investors save. Third, if investors care for the rates of return to their investments in terms of physical goods, the rates measured at horizons over which goods prices can fully adjust would be the most relevant measure.

While RIP has been repeatedly rejected by the numerous preceding studies, our examination of the longer maturity yields obtains much more favorable evidence. We tested the RIP hypothesis first by assuming that expectations are rational, and then, by using time series forecasts of future inflation rates. With both methods we find cases where the hypothesis of equal ex ante real interests cannot be rejected for the selected G7 countries at a five- and/or a ten-year horizon. Clearly the long horizon interest rate data we adopt are not free from various distortions and shortcomings as already noted in section 1, and thus, the empirical results need to be interpreted with caution. Nevertheless, they suggest that by the end of the last century, real rates of return were virtually equalized among the key industrialized economies.

In presenting the results, we have refrained from discussing why real interest parity appears to hold better at long than short horizons. We believe that the result arises from two causes. First, as the recent studies by Meredith and Chinn (1998) and Alexius (1998) report, uncovered interest parity holds better at long than short horizons. Second, relative purchasing power parity appears to hold better at longer horizons. Our findings therefore add to the growing consensus that at long horizons, arbitrage conditions exert greater force on international goods and asset markets so that fundamentals matter (Flood and Taylor, 1997). For future work, it would be useful to examine a broader set of financial instruments, as such data become available, to see if a similar conclusion is obtained.

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Table 1

A. Short Horizon Realized Real Interest Differentials

	CPI-based Differentials		WPI-based Differentials	
	Mean	Standard Deviation	Mean	Standard Deviation
3-month rate				
Canada	-.82** (.27)	2.62	.23 (.40)	3.96
France	-1.30** (.30)	2.94	n.a.	n.a.
Germany	-.07 (.33)	3.17	.59 (.49)	4.80
Italy	-1.17** (.37)	3.46	n.a.	n.a.
Japan	.71* (.35)	3.28	.48 (.67)	6.25
U.K.	-.58 (.41)	4.06	.41 (.48)	4.75
6-month rate				
Canada	-.83** (.32)	2.39	.21 (.45)	3.34
France	-1.27** (.37)	2.73	n.a.	n.a.
Germany	-.02 (.37)	2.78	.65 (.58)	4.26
Italy	-1.12** (.43)	3.16	n.a.	n.a.
Japan	.80* (.36)	2.62	.59 (.77)	5.54
U.K.	-.51 (.40)	3.07	.50 (.55)	4.14
1-year rate				
Canada	-.83* (.40)	2.07	.22 (.55)	2.87
France	-1.26** (.46)	2.55	n.a.	n.a.
Germany	.06 (.44)	2.39	.72 (.74)	3.88
Italy	-1.15* (.55)	2.94	n.a.	n.a.
Japan	.91 (.47)	2.43	.77 (.85)	4.53
U.K.	-.45 (.49)	2.71	.57 (.65)	3.58

B. Long Horizon Realized Real Interest Differentials

	CPI-based Differentials		WPI-based Differentials	
	Mean	Standard Deviation	Mean	Standard Deviation
Synthetic 5-year rate				
Canada	-.97 (.63)	1.62	-.02 (.24)	1.02
Germany	-1.26 (.96)	2.62	-.64 (.36)	1.61
U.K.	-.21 (.61)	2.06	2.73* (1.09)	3.55
Synthetic 10-year rate				
Germany	-.54 (1.22)	2.69	.31 (1.09)	2.68
Japan	-.09 (.94)	2.27	-.93 (.95)	2.29
U.K.	-0.42 (.44)	1.74	.58 (.50)	1.75
10-year rate				
Canada	-.61 (.34)	.88	.11 (.27)	.73
France	-.99* (.50)	1.29	n.a.	n.a.
Germany	-.47 (1.18)	2.64	.38 (1.05)	2.64
Italy	-.86* (.36)	1.67	n.a.	n.a.
Japan	-.24 (1.11)	2.51	-1.08 (1.12)	2.59
U.K.	.15 (.66)	2.17	1.15 (.74)	2.26

Notes: Means and standard deviations of the realized real interest rate differentials vis-à-vis the U.S. are reported in annual percentage terms. Panels A and B summarize the short horizon and long horizon data, respectively. The numbers in the parentheses are the corrected asymptotic standard errors for the means. ** and *denote statistical significance at 1 % and 5 % levels, respectively. Due to data limitation, WPI-based differentials are not available for France and Italy. The sample periods are as specified in the data appendix.

Table 2

A. The Short Horizon Realized Real Interest Parity Results with CPI Inflation

	α	β	β=1	α=0 and β=1	adjusted R^2
3-month rate					
Canada	.008 (.005)	.587** (.105)	15.462**	24.884**	.281
France	.009* (.004)	.479** (.090)	33.604**	53.035**	.304
Germany	.019** (.004)	.363** (.093)	28.059**	28.084**	.098
Italy	.020** (.005)	.302** (.093)	56.757**	74.411**	.150
Japan	.026** (.004)	.294** (.099)	51.279**	59.904**	.101
U.K.	.020** (.003)	.287** (.065)	121.512**	125.515**	.193
6-month rate					
Canada	.002 (.005)	.730** (.134)	4.068*	8.771*	.374
France	.007 (.005	.562** (.120)	13.340**	23.129**	.335
Germany	.012* (.005)	.602** (.177)	5.058*	5.458	.182
Italy	.017** (.005)	.393** (.107)	32.170**	35.072**	.209
Japan	.019** (.005)	.578** (.165)	6.578*	13.256**	.221
U.K.	.014** (.003)	.466** (.087)	37.344**	39.333**	.308
1-year rate					
Canada	-.003 (.006)	.877** (.145)	.727	4.624	.534
France	.004 (.008)	.646* (.149)	5.635*	12.593**	.404
Germany	.000 (.008)	1.025** (.268)	.009	.0198	.368
Italy	.013** (.005)	.488** (.106)	23.455**	23.455**	.307
Japan	.016* (.008)	.750** (.222)	1.268	5.170	.321
U.K.	.011** (.004)	.584** (.117)	12.644**	13.759**	.392

B. The Long Horizon Realized Real Interest Parity Results with CPI Inflation

	α	β	β=1	α=0 and β=1	adjusted R^2
Synthetic 5-year rate					
Canada	-.003 (.005)	.850** (.147)	1.040	2.606	.607
Germany	-.013 (.010)	1.255** (.251)	1.034	1.717	.650
U.K.	.000 (.007)	.888** (.233)	.230	.859	.547
Synthetic 10-year rate					
Germany	-.009 (.005)	1.395** (.071)	30.645**	98.671**	.779
Japan	.006 (.009)	1.161** (.213)	.5760	177.271**	.521
U.K.	-.004 (.008)	1.102** (.166)	.374	.4974	.506
10-year rate					
Canada	-.002* (.006)	.917** (.057)	2.118	6.006*	.797
France	-.012* (.006)	.901** (.069)	2.057	565.317**	.734
Germany	-.009* (.005)	1.398** (.068)	34.303**	105.502**	.783
Italy	.020** (.005)	.539** (.101)	20.782**	24.256**	.350
Japan	.010 (.007)	1.117** (.183)	.408	243.165**	.545
U.K.	.009 (.011)	.927** (.198)	.136	2.022	.376

Notes: Panels A and B summarize the estimation results of (7) in the text:

$$i_t^k - \pi_{t,t+k} = \alpha + \beta(i_t^{k*} - \pi^*_{t,t+k}) + \varepsilon_{t+k}$$

with CPI inflation at short and long horizons, respectively. The second and third columns of each panel contain the coefficient estimates with the corrected standard errors in the parentheses below. The fourth and fifth columns provide the Wald test statistics for the null hypothesis indicated in the top row. ** and *denote statistical significance at 1 % and 5 % levels, respectively. The sample periods are as specified in the data appendix.

Table 3

A. The Short Horizon Realized Real Interest Parity Results with WPI Inflation

	α	β	β=1	α=0 and β=1	adjusted R^2
3-month rate					
Canada	.015** (.004)	.692** (.074)	17.409**	22.208**	.477
Germany	.019** (.007)	.669 (.167)	3.941*	6.875*	.140
Japan	.035** (.005)	.321** (.059)	134.109**	141.530**	.164
U.K.	.021** (.005)	.587** (.093)	19.668**	24.577**	.229
6-month rate					
Canada	.014** (.004)	.723** (.085)	10.535**	16.637**	.561
Germany	.017* (.009)	.730** (.197)	1.875	4.058	.180
Japan	.035** (.006)	.362** (.064)	98.413**	107.001**	.226
U.K.	.019** (.006)	.654** (.099)	12.200**	14.066**	.279
1-year rate					
Canada	.013** (.004)	.773* (.108)	4.438*	9.136*	.645
Germany	.013 (.011)	.860** (.255)	.301	1.599	.257
Japan	.032** (.006)	.477** (.060)	75.176**	90.663**	.368
U.K.	.014 (.009)	.813** (.146)	1.647	2.308	.383

B. The Long Horizon Realized Real Interest Parity Results with WPI Inflation

	α	β	β=1	α=0 and β=1	adjusted R^2
Synthetic 5-year rate					
Canada	.000 (.003)	.966** (.096)	.121	.125	.829
Germany	-.003 (.006)	1.156** (.221)	.497	.517	.633
U.K.	-.007 (.004)	1.232** (.161)	2.075	3.397	.677
Synthetic 10-year rate					
Germany	-.003 (.005)	1.362** (.085)	18.048**	51.067**	.862
Japan	.021** (.006)	.789** (.086)	6.008*	28.628**	.509
U.K.	.001 (.003)	1.142** (.099)	2.046	9.4482**	.576
10-year rate					
Canada	.006** (.004)	.909** (.028)	10.688**	18.628**	.931
Germany	-.003 (.005)	1.368** (.084)	19.130**	58.294**	.866
Japan	.0220** (.005)	.775** (.078)	8.248**	46.929**	.514
U.K.	.008 (.005)	1.081** (.088)	.845	6.688*	.495

Notes: Panels A and B summarize the estimation results of (7) in the text:

$$i_t^k - \pi_{t,t+k} = \alpha + \beta(i_t^{k*} - \pi_{t,t+k}^*) + \varepsilon_{t+k}$$

with WPI inflation at short and long horizons, respectively. The second and third columns of each panel contain the coefficient estimates with the corrected standard errors in the parentheses below. The fourth and fifth columns provide the Wald test statistics for the null hypothesis indicated in the top row. ** and *denote statistical significance at 1 % and 5 % levels, respectively. The sample periods are as specified in the data appendix.

Table 4

A. The Short Horizon Real Interest Parity Results with CPI Inflation

	α	β	β=1	α=0 and β=1	adjusted R^2
3-month rate					
Canada	.027* (.011)	.140 (.244)	12.392**	33.804**	.134
France	.039** (.010)	-.120 (.218)	26.491**	59.986**	.030
Germany	.042** (.009)	-.228 (.261)	22.148**	22.264**	-.008
Italy	.057** (.015)	-.471 (.292)	25.306**	77.274**	.014
Japan	.030** (.003)	.104 (.098)	83.796**	145.441**	.050
U.K.	.042** (.007)	-.195 (.132)	82.046**	122.073**	.015
6-month rate					
Canada	.012 (.018)	.495 (.406)	1.552	18.397**	.161
France	.039* (.011)	-.091 (.238)	21.050**	37.606**	.075
Germany	.042** (.011)	-.152 (.324)	12.623**	14.600**	-.001
Italy	.054** (.018)	-.357 (.341)	15.874**	41.085**	-.001
Japan	.030** (.004)	.192 (.219)	26.684**	67.112**	.123
U.K.	.056** (.016)	-.464 (.345)	18.049**	31.053**	.030
1-year rate					
Canada	.006 (.019)	.648 (.390)	.815	8.675*	.370
France	.037* (.018)	.034 (.369)	6.838**	15.659**	.202
Germany	.037** (.014)	.083 (.420)	4.773*	9.669**	.051
Italy	.010 (.040)	.537 (.728)	.404	29.799**	.050
Japan	.031** (.005)	.329 (.222)	9.145**	36.009**	.167
U.K.	.065* (.026)	-.557 (.526)	8.758**	12.874**	.062

B. The Long Horizon Real Interest Parity Results with CPI Inflation

	α	β	β=1	α=0 and β=1	adjusted R^2
Synthetic 5-year rate					
Canada	.006 (.008)	.624** (.167)	5.052*	6.825*	.607
Germany	.015 (.019)	.655 (.404)	.726	.727	.650
U.K.	.002 (.013)	.897** (.258)	.159	.950	.547
Synthetic 10-year rate					
Germany	-.076** (.015)	2.757** (.345)	25.880**	25.907**	.746
Japan	.097** (.030)	-.742 (.509)	11.686**	13.437**	.329
U.K.	-.069** (.021)	2.206** (.376)	10.321**	13.242**	.616
10-year rate					
Canada	.007* (.003)	.759** (.064)	14.146**	15.188**	.854
France	.019** (.007)	.498** (.102)	24.217**	117.688**	.828
Germany	-.052** (.009)	2.263** (.205)	38.048**	40.171**	.780
Italy	.046** (.015)	.193 (.239)	11.444**	13.393**	.553
Japan	.088** (.025)	-.571 (.437)	12.920**	13.772**	.407
U.K.	-.008 (.013)	1.328** (.268)	1.493	13.650**	.425

Notes: Panels A and B summarize the generalized method of moments estimation results of (10) in the text

$$i_t^k - \pi_{t,t+k} = \alpha + \beta(i_t^{k*} - \pi^*_{t,t+k}) + \omega_{t+k}$$

with CPI inflation at short and long horizons, respectively. For all cases, the instruments are the three-month real interest rates lagged by k through k+3, where k is the maturity in quarters. The second and third columns contain the coefficient estimates with the corrected standard errors in the parentheses below. The fourth and fifth columns provide the Wald test statistics for the null hypothesis indicated in the top row. The sample period is 1973Q1-1990Q1. ** and *denote statistical significance at 1 % and 5 % levels, respectively.

Table 5

A. The Short Horizon Real Interest Parity Results with WPI Inflation

	α	β	β=1	α=0 and β=1	adjusted R^2
3-month rate					
Canada	.027** (.007)	.463** (.136)	15.611**	17.715**	.264
Germany	.061** (.013)	-.246 (.272)	21.041**	22.651**	.021
Japan	.042** (.010)	.123 (.167)	27.730**	27.925**	.111
U.K.	.026 (.015)	.475 (.289)	3.311	3.312	.139
6-month rate					
Canada	.028** (.007)	.465** (.128)	17.533**	19.242**	.328
Germany	.067* (.015)	-.343 (.306)	19.236**	20.267**	.018
Japan	.046** (.011)	.100 (.199)	20.450**	20.662**	.172
U.K.	.035 (.019)	.349 (.322)	4.096*	4.129	.128
1-year rate					
Canada	.033** (.010)	.433** (.168)	11.418**	11.746**	.427
Germany	.078** (.021)	-.482 (.388)	14.608**	14.750**	.034
Japan	.060** (.012)	-.001 (.196)	26.018**	28.736**	.223
U.K.	.031 (.024)	.413 (.392)	2.238	2.813	.147

B. The Long Horizon Real Interest Parity Results with WPI Inflation

	α	β	β=1	α=0 and β=1	adjusted R^2
Synthetic 5-year rate					
Canada	.011 (.006)	.786** (.111)	3.699	4.414	.718
Germany	.097* (.049)	-.699 (.768)	4.889*	5.559	.427
U.K.	.050 (.026)	.155 (.374)	5.100*	5.564	.553
Synthetic 10-year rate					
Germany	.101* (.045)	-.513 (.758)	3.984*	7.616*	.856
Japan	.040** (.013)	.519** (.151)	10.192**	10.201**	.508
U.K.	-.009 (.015)	1.296** (.233)	1.614	45.878**	.662
10-year rate					
Canada	.019** (.003)	.726** (.059)	21.944**	46.208**	.930
Germany	.092* (.044)	-.0365 (.733)	3.462	7.497*	.857
Japan	.046** (.010)	.440** (.111)	25.412**	26.712**	.551
U.K.	.034 (.028)	.709 (.483)	.365	16.443**	.454

Notes: Panels A and B summarize the generalized method of moments estimation results of (10) in the text

$$i_t^k - \pi_{t,t+k} = \alpha + \beta(i_t^{k*} - \pi^*_{t,t+k}) + \omega_{t+k}$$

with WPI inflation at short and long horizons, respectively. For all cases, the instruments are the three-month real interest rates lagged by *k* through *k*+3, where *k* is the maturity in quarters. The second and third columns contain the coefficient estimates with the corrected standard errors in the parentheses below. The fourth and fifth columns provide the Wald test statistics for the null hypothesis indicated in the top row. The sample period is 1973Q1-1990Q1. ** and *denote statistical significance at 1 % and 5 % levels, respectively.

Table 6

The Selected Model Specifications for the Quarterly Inflation Series

	CPI Inflation	WPI Inflation
U.S.	AR(3)	AR(3)
Canada	AR(4)	AR(3)
France	AR(5)	n.a.
Germany	AR(4)	AR(1)
Italy	AR(2)	n.a.
Japan	AR(4)	AR(1)
U.K.	AR(4)	AR(3)

Notes: The Schwarz-Bayesian criteria is used to select the model specifications.

Table 7

A. The Short Horizon Real Interest Parity Results with Forecasted CPI Inflation

	α	β	β=1	α=0 and β=1	adjusted R^2
3-month rate					
Canada	.018** (.005)	.578** (.067)	39.313**	70.090**	.510
France	0.036** (.005)	.289** (.082)	74.828**	76.327**	.154
Germany	.060** (.005)	-.066 (.108)	97.084**	144.828** *	-.012
Italy	.029 (.007)	.312** (.084)	66.539**	266.123**	.201
Japan	.043** (.002)	.413** (.063)	86.283**	410.443**	.270
U.K.	.023** (.004)	.438** (.055)	104.906**	191.631**	.292
6-month rate					
Canada	.016* (.007)	.564** (.109)	16.085**	31.792**	.391
France	.032** (.006)	.286* (.115)	38.528**	36.612**	.128
Germany	.054** (.006)	-.067 (.131)	66.150**	77.532**	-.012
Italy	.028** (.009)	.291* (.137)	26.880**	104.245**	.133
Japan	.038** (.003)	.397** (.094)	41.283**	164.969**	.230
U.K.	.018** (.006)	.473** (.093)	32.456**	65.250**	.230
1-year rate					
Canada	.018* (.008)	.403* (.178)	11.281**	15.560**	.130
France	.031** (.007)	.129 (.186)	22.021**	22.088**	.001
Germany	.039** (.008)	-.035 (.193)	28.878**	28.901**	-.015
Italy	.035* (.009)	.059 (.185)	25.836**	30.254**	-.012
Japan	.029** (.004)	.327* (.157)	18.386**	41.788**	.120
U.K.	.012 (.008)	.516** (.179)	7.287**	14.258**	.112

B. The Long Horizon Real Interest Parity Results with Forecasted CPI Inflation

	α	β	β=1	α=0 and β=1	adjusted R^2
Synthetic 5-year rate					
Canada	-.051** (.009)	.224 (.138)	31.771**	60.018**	.107
Germany	-.061** (.014)	.112 (.088)	100.845**	101.292**	.006
U.K.	-.033** (.009)	.371** (.063)	100.573**	103.138**	.394
Synthetic 10-year rate					
Germany	-.144** (.023)	.133* (.068)	163.551**	167.418**	.016
Japan	-.145** (.014)	.592** (.189)	4.669*	529.406**	.160
U.K.	-.074** (.011)	.413** (.036)	270.392**	283.769**	.516
10-year rate					
Canada	-.122** (.006)	.235** (.088)	75.684**	451.170**	.156
France	-.150** (.022)	.096 (.126)	51.704**	57.537**	.005
Germany	-.159** (.017)	.064 (.069)	184.549**	213.372**	-.010
Italy	-.152** (.028)	.072 (.151)	37.697**	38.183**	-.006
Japan	-.152** (.010)	.581** (.207)	4.093*	584.430**	.155
U.K.	-.079** (.014)	.405** (.035)	284.778**	415.778**	.517

Notes: Panels A and B summarize the generalized method of moments estimation results of (8) in the text:

$$i_t^k - \pi_{t,t+k}^e = \alpha + \beta(i_t^{k*} - \pi_{t,t+k}^{e*}) + \varepsilon_{t+k}$$

with forecasted CPI inflation rates at short and long horizons, respectively. The second and third columns contain the coefficient estimates with the corrected standard errors in the parentheses below. The fourth and fifth columns provide the Wald test statistics for the null hypothesis indicated in the top row. ** and *denote statistical significance at 1 % and 5 % levels, respectively. The sample period is 1983Q1-2000Q1.

Table 8

A. The Short Horizon Real Interest Parity Results with Forecasted WPI Inflation

	α	β	β=1	α=0 and β=1	adjusted R^2
3-month rate					
Canada	.026** (.004)	.523** (.059)	65.280**	71.507**	.528
Germany	.069** (.006)	-.132 (.112)	101.575**	148.697**	.001
Japan	.047** (.002)	.377** (.050)	153.684**	491.844**	.273
U.K.	.026** (.004)	.445** (.049)	128.584**	201.779**	.340
6-month rate					
Canada	.029** (.005)	.481** (.090)	32.984**	33.029**	.416
Germany	.068** (.008)	-.134 (.155)	53.629**	69.259**	-.001
Japan	.045** (.003)	.358** (.066)	94.160**	182.870**	.253
U.K.	.027* (.006)	.450** (.082)	44.608**	64.505**	.251
1-year rate					
Canada	.033** (.006)	.421** (.120)	23.395**	32.542**	.263
Germany	.059** (.012)	-.015 (.229)	19.610**	22.876**	-.016
Japan	.040** (.006)	.375** (.103)	37.120**	45.244**	.216
U.K.	.029* (.012)	.452* (.175)	9.808**	12.410**	.119

B. The Long Horizon Real Interest Parity Results with Forecasted WPI Inflation

	α	β	β=1	α=0 and β=1	adjusted R^2
Synthetic 5-year rate					
Canada	.038** (.013)	.316** (.097)	50.026**	95.578**	.186
Germany	-.020 (.019)	1.070** (.340)	.043	2.351	.103
U.K.	.030* (.015)	.608* (.305)	1.651	4.601	.084
Synthetic 10-year rate					
Germany	-.043* (.017)	1.394* (.274)	2.065	7.850*	.121
Japan	-.013 (.028)	.918* (.379)	.047	6.664*	.174
U.K.	.037** (.012)	.590* (.237)	2.980	9.833**	.080
10-year rate					
Canada	.044** (.014)	.209** (.059)	181.969**	478.485**	.130
Germany	-.034 (.025)	1.194** (.420)	.212	4.787	.086
Japan	-.023 (.029)	1.030* (.413)	.005	10.003**	.225
U.K.	.039** (.013)	.389 (.219)	7.812**	9.934**	.031

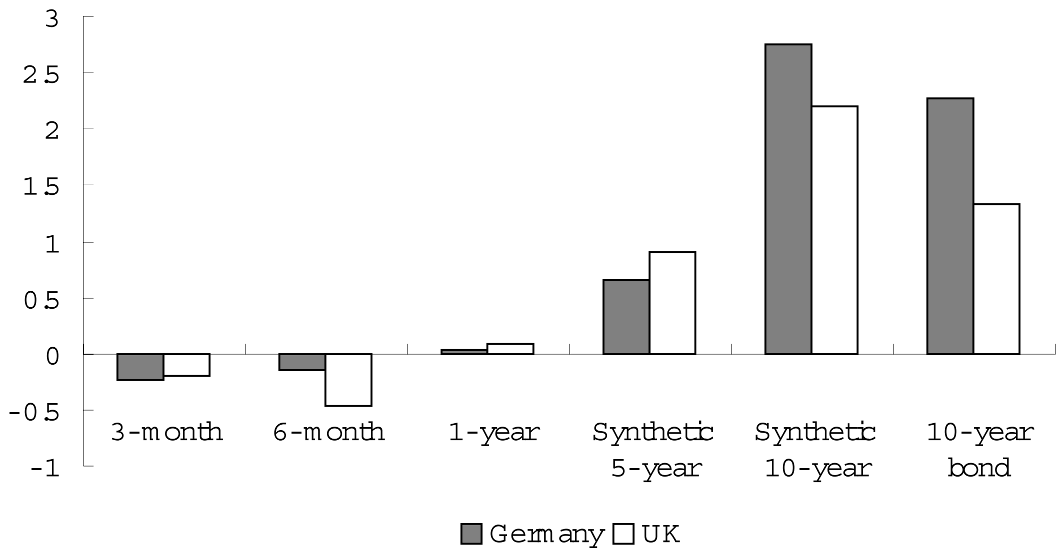
Notes: Panels A and B summarize the generalized method of moments estimation results of (8) in the text:

$$i_t^k - \pi_{t,t+k}^e = \alpha + \beta(i_t^{k*} - \pi_{t,t+k}^{e*}) + \varepsilon_{t+k}$$

with forecasted WPI inflation rates at short and long horizons, respectively. The second and third columns contain the coefficient estimates with the corrected standard errors in the parentheses below. The fourth and fifth columns provide the Wald test statistics for the null hypothesis indicated in the top row. ** and *denote statistical significance at 1 % and 5 % levels, respectively. The sample period is 1983Q1-2000Q1.

Figure 1

The Point Estimates of β by Horizons



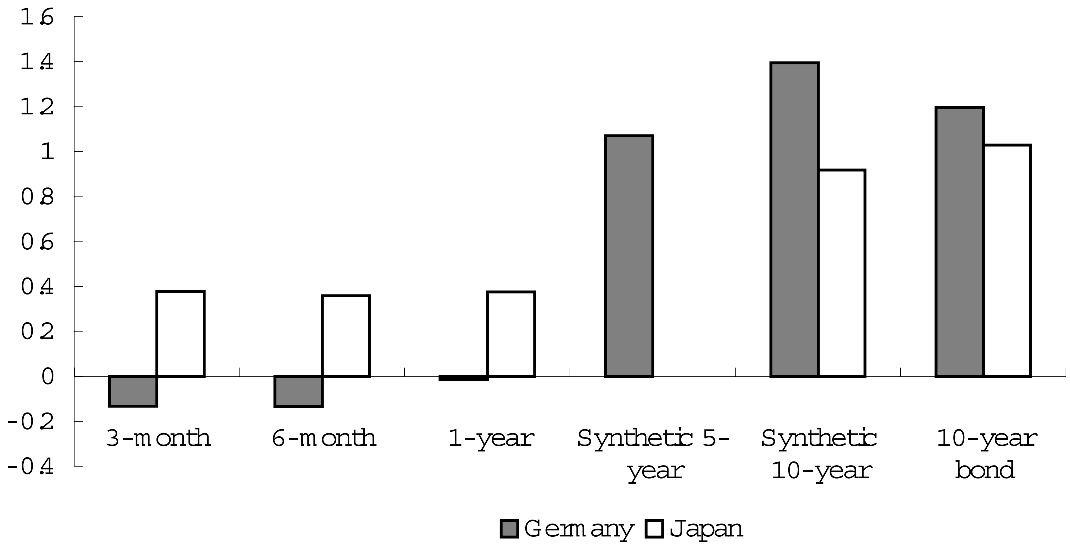
Notes: The graph shows the point estimates of β in (10) in the text:

$$i_t^k - \pi_{t,t+k} = \alpha + \beta(i_t^{k*} - \pi_{t,t+k}^*) + \omega_{t+k}$$

for the German and U.K data with CPI inflation by horizons. The sample period is 1973Q1-1990Q1.

Figure 2

The Point Estimates of β by Horizons with Forecasted Inflation



Notes: The graph shows the point estimates of β in (8) in the text:

$$i_t^k - \pi_{t,t+k}^e = \alpha + \beta(i_t^{k*} - \pi_{t,t+k}^{e*}) + \varepsilon_{t+k}$$

for the German and Japanese data with forecasted WPI inflation by horizons. The sample period is 1983Q1-2000Q1 for all cases. The synthetic five-year interest rate data is not available for Japan, and hence, the corresponding estimate of β is not graphed.

Data Appendix

1. Short-term interest rates

Short-term rates are end-of-month 3-, 6-, and 12-month maturity eurocurrency yields. The data are available for the following sample periods.
1976Q1-2000Q1 for the U.S., Canada, and the U.K.
1976Q1-1999Q2 for France and Germany
1978Q2-1999Q3 for Italy
1978Q2-2000Q1 for Japan

2. Long-term interest rates

a. The 10-year government bond rates
End-of-month yields on benchmark government bonds of ten-year maturity at the date of issuance used by Edison and Pauls (1993). The sample periods are,
1973Q1-1997Q4 for all except Italy
1977Q1-1997Q4 for Italy
b. Synthetic constant maturity five- and ten-year rates
End-of-month rates interpolated from the yield curve of outstanding government securities. The data are obtained from the International Monetary Fund country desks. The sample periods are as follows.
The five-year rates
1973Q1-2000Q1 for the U.S., Canada, Germany, and the U.K.
The ten-year rates
1973Q1-2000Q1 for the U.S., Germany, Japan, and the U.K.