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Co-Movements in Long-Term Interest Rates and the Role of PPP-Based Exchange Rate Expectations

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Abstract

This paper investigates international co-movement in bond yields by testing for uncovered interest parity (UIP). Existing work is supplemented by focusing on long instead of short-term interest rates and by employing exchange rate expectations derived from purchasing power parity (PPP) instead of actual outcomes. Among the major currencies during 1975-97, the paper does not find a further increase in co-movement beyond that associated with the wave of financial market liberalization in the early 1980s. Given the similarity between PPP-based UIP tests and those employing actual exchange rate outcomes, the value added of the former lies mainly with data availability.

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I. Introduction

Developments in bond markets over the past several years, in particular the decline in international yield spreads, have raised questions about the relations between long-term interest rates, globalization of financial activities, and the exchange rate regime. It is generally accepted that under fixed exchange rates, globalization of financial activities implies a high degree of convergence of long-term interest rates and synchronization of their movements over time. Interest rates are determined by conditions in the fixed exchange rate region as a whole, rather than in individual countries, and there is correspondingly reduced scope for independent monetary policy by a single country.[2] Under floating exchange rates, interest rates differ across countries, in both real and nominal terms, because the existing pressures on financial markets are absorbed by movements in the countries' (expected) exchange rates. A rise in one country's interest rate relative to that of a partner is effectively offset by an expected future depreciation or a rise in the relative risk premium on its currency.

Notwithstanding the greater importance of long-term interest rates for the business cycle in many industrialized countries, the empirical literature on international interest rate linkages focussing on short-term assets generally outnumbers the studies using long-term rates.[3] This reflects the fact that financial instruments traded at the long end of the market generally lack a well-developed forward market. While it is true that forward markets are most developed at the three-month maturity and hardly exist at maturities greater than two years, even among well-traded currencies, in the 1980s the currency swap market became sufficiently developed to hedge exchange rate risk for long-term investments as well (Popper, 1990, 1993, and Fletcher and Taylor, 1996). In addition to the gradual worldwide abolition of capital controls, this financial innovation may have increased the linkage of international long-term interest rates.[4]

Financial innovations such as currency swaps should contribute to international capital market integration by increasing capital mobility. In this paper we investigate whether this increased mobility has led to a measurable increase in co-movement of national bond yields. The focus on long-term interest rate differentials is by itself an important contribution to the

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[2]This does not apply to the anchor country in an asymmetric, fixed but adjustable exchange rate regime like Bretton-Woods or the Exchange Rate Mechanism (ERM) of the European Monetary System (De Grauwe, 1994).

[3]Recent studies involving long-term rates are Kasman and Pigott (1988), Howe and Pigott (1991), Pigott (1993), Throop (1994), Fell (1996), Kirchgässner and Wolters (1996), Fase and Vlaar (1998), and Meredith and Chinn (1998).

[4]Deviations from covered interest parity appear to be somewhat larger among long-term assets than among short-term assets, but for the major currencies the differences remain fairly limited (Popper, 1993; Fletcher and Taylor, 1996).

existing literature, which, as mentioned earlier, mainly concentrates on short-term instruments. A second but related novelty is that we propose to use a PPP-based measure of exchange rate expectations instead of the more common 'rational' expectations, which allows us to directly assess the uncovered interest parity (UIP) relationship. Our hypothesis is that increases in international capital mobility may be expected to result in a greater likelihood of validating the UIP relationship, as, under these circumstances, the scope for interest rate arbitrage increases and the impact on interest rate differentials of factors other than (expected) exchange rate movements and time-varying risk premia should steadily diminish.

Empirical tests of UIP thus far have typically been combined with the rational expectations hypothesis, allowing actual exchange rate movements to proxy for expected ones. The bulk of this evidence indicates not just that exchange rate changes fail to move one-for-one with interest rate differentials, but rather that these changes are substantial and in the *opposite* direction to that implied by UIP. Froot (1990) has calculated an average estimated regression coefficient across some 75 published estimates of *minus* 0.88. A few estimates are positive, but all of them are distinctly smaller than the null hypothesis of +1. McCallum (1994) and Meredith and Chinn (1998) have argued that for relatively short horizons, failure of UIP results from risk-premium shocks in the face of endogenous monetary policy. In the longer term, in contrast, exchange rate movements are driven by 'fundamentals', leading to a relation between interest rates and exchange rates that should be more consistent with UIP.

The remainder of the paper is organized as follows. In the next section we present some stylized facts. In Section III we derive the UIP relation from a portfolio model of interest rate determination, followed by a description of data and methodology in Section IV. Section V puts the theoretical predictions to an empirical test. Section VI concludes.

II. Recent Developments in International Bond Markets: A First Look

Figure 1 displays the development of long-term (ten year) government bonds yields for the United States, Germany, France, United Kingdom, Switzerland, and Japan from 1975 onward.[5] In the aftermath of two oil crises during the second half of the 1970s many countries experienced strong upward pressure on long-term interest rates, when inflation pressures rose and, around the turn of the decade, monetary policy was tightened quite aggressively. After reaching record highs in the early 1980s, bond yields have gradually come down and there has been a marked convergence of nominal (and also real) long-term interest rates among the

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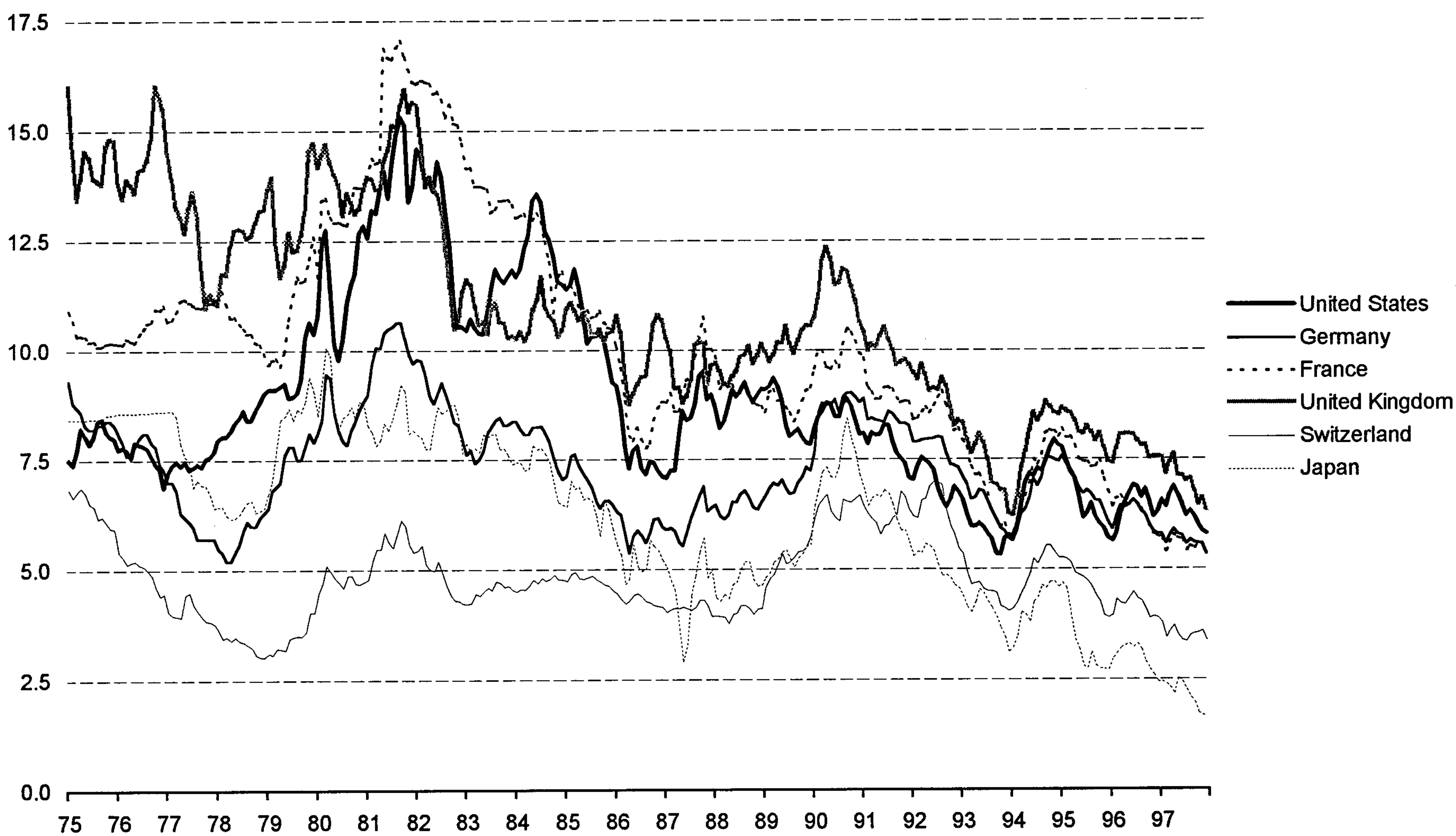
[5]We used monthly data on so-called benchmark bonds, obtained from BIS. It is important that the properties of these bond contracts are identical (so that they differ only in currency of denomination), and benchmark bonds come closest to this ideal. Differences are however likely to remain, for example stemming from variations in the degree of liquidity of underlying markets as well as differences in tax treatment, payment of coupons, and in the method by which the contracts are sold (Holmes and Wu, 1997).

Figure 1. Long-Term Government Bond-Yields



Source: BIS.

countries under investigation (perhaps less so for Japan and Switzerland). This trend broadly corresponds to the declining trend in inflation during this period and the simultaneous reduction in inflation divergences across the major regions. In most countries, this process was interrupted in the periods surrounding the stock market collapse of 1987 and the tightening of U.S. monetary policy in 1994. With the benefit of hindsight, the causes of these bond market corrections are well known. In both phases, substantial concerns about the return of inflationary pressures in the U.S. economy developed, and bond yields rose accordingly.

Table 1 documents the major swings which have taken place in international bond markets since 1975 in more detail.[6] Especially during the 1970s and the early 1980s, long-term interest rates across the major countries have been pushed apart, often with accompanying movements in exchange rates. Countries that had experienced the largest inflation increases toward the end of the 1970s, such as the United States, the United Kingdom, and France, had to allow for higher yields than Germany and Japan around the turn of the decade, while their currencies generally fell vis-à-vis the deutsche mark and the yen. U.S. long-term interest rates rose between spring 1983 and mid-1984 while at the same time interest rates in Europe and Japan were generally falling, due in part to the effects of U.S. fiscal expansion in raising the demand for domestic savings relative to its supply. The pressures resulting from this divergence in long-term interest rates were also reflected in the sharp appreciation of the dollar in this period, as well as in the reversal of that appreciation that began in 1985 and occurred as U.S. bond yields were falling back.

For the 1990s, however, Table 1 would seem to reveal a somewhat larger synchronization among government bond yields than in the previous decades. Apart from the correction after the monetary tightening by the Federal Reserve in 1994 and an additional hiccup during the first half of 1996,[7] the 1990s have been characterized by declining bond yields worldwide. In view of the experience of the 1980s, it is the circumstances surrounding the interest rate developments, and not the mere fact that they moved together, that would seem somewhat unusual. The 1994 long-term interest rate run-ups in Europe and Japan followed a period of declining short-term interest rates and weak economic activity. The experience of the 1990s thus suggests that the degree to which long-term interest rates are free to vary with national conditions should not be overstated, even under floating exchange

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[6]The methodology which has been adopted to classify a major swing in the bond market is similar to that used by the NBER to date business cycles. Peaks and troughs were identified using a 12-month moving window. A 'bull' market is defined as at least 6 consecutive months of declining bond yields and a 'bear' market is defined in the opposite way.

[7]This hiccup is not identified in Table 1, as the number of consecutive months with unidirectional movements was limited to 5 instead of the required 6 (see previous note).

Table 1: Peak to Trough Analysis of Major Swings in International Bond Markets

Market	United States		Germany		France		United Kingdom		Switzerland		Japan	
type	Start	Change	Start	Change	Start	Change	Start	Change	Start	Change	Start	Change
bull	75/09	-156	75/01	-410	75/01	-80			75/04	-384		
bear	76/12	+845			75/09	+126	75/03	+260			75/01	+19
bull					78/02	-172	76/10	-505			76/12	-245
bear			78/04	+544	79/04	+741	77/10	+374	78/12	+310	78/04	+389
bull							79/12	-160			80/03	-223
bear							80/10	+283			81/03	+139
bull	81/09	-494	81/09	-322	81/09	-928	81/10	-719	81/09	-191	81/09	-631
bear	83/05	+318	83/03	+96					83/01	+72		
bull	84/06	-648	83/12	-301					85/03	-117		
bear	87/01	+244	86/04	+367	86/08	+299	86/04	+204			87/05	+552
bull	87/10	-168			87/10	-239	86/11	-160				
bear	89/12	+105			89/08	+211	88/04	+314	88/04	+292		
bull	90/09	-356	90/10	-339	90/09	-481	90/04	-611	90/04	-264	90/09	-531
bear	93/10	+262	94/01	+194	94/01	+252	94/01	+257	93/12	+149	93/12	+160
bull	94/11	-343	95/01	-355	95/01	-400	94/09	-381	94/09	-286	94/10	-390
end	97/12		97/12		97/12		97/12		97/12		97/12	

Source: Authors' own calculations.

Notes: Changes are measured in basis points. Peaks and troughs were identified using a 12-month moving window. A 'bull' market is defined as at least 6 consecutive months of declining bond yields and a 'bear' market is similarly defined for rising bond yields.

rates. Globalization of portfolio trading strategies may have led to a more profound influence of aggregate international conditions with resulting spillovers among markets.[8]

More formal evidence on the degree to which long-term interest rates have become more synchronized with globalization can be extracted from bilateral correlations of their changes against those in the major market. In this respect, the U.S. bond market is clearly the world's dominant market, with a liquidity about three times that of Germany, the second largest market, for instance. For the five countries under investigation, Figure 2 displays 5-year rolling correlations of monthly changes in domestic bond yields vis-à-vis U.S. yields since the mid-1970s. Apart from the stylized fact that co-movements in long-term interest rates among the major industrialized economies were stronger in the 1980s and 1990s than during the 1970s (see among others Frankel, 1989), by this measure there seems to have been no systematic further increase in the bilateral synchronization since the early 1980s. Furthermore, correlations have varied considerably over time. While interest rates were indeed relatively highly correlated over the most recent couple of years (again perhaps with the exception of Switzerland), generally they were not more so than in the 1980s.

Finally, Figure 3 looks into the co-movements in bond yields from a multilateral perspective, by depicting the proportion of variance among the various interest rates explained by the first four principal components of the changes in bond yields in the countries under investigation. Following the suggestion of Gagnon and Unferth (1995), we interpret the first principal component—which is optimal in the sense that no other linear combination accounts for more of the total variance of the interest rates—as the 'world' interest rate. The more intense capital market integration, the higher the correlations between interest rates and the greater the proportion of variance explained by the first component. Figure 3 would suggest that the international integration of bond markets did indeed increase somewhat further since the mid-1980s, albeit very gradually.[9]

All in all, the message from Figures 1-3 and Table 1 is rather mixed. This should not be too much of a surprise, since the analysis thus far has been confined to bond markets only, without any reference to contemporaneous developments in foreign exchange markets. As is well known, however, exchange rates among the largest industrialized countries have fluctuated quite substantially in the period under investigation. In the remainder of this paper, we shall therefore introduce exchange rate movements into the analysis.

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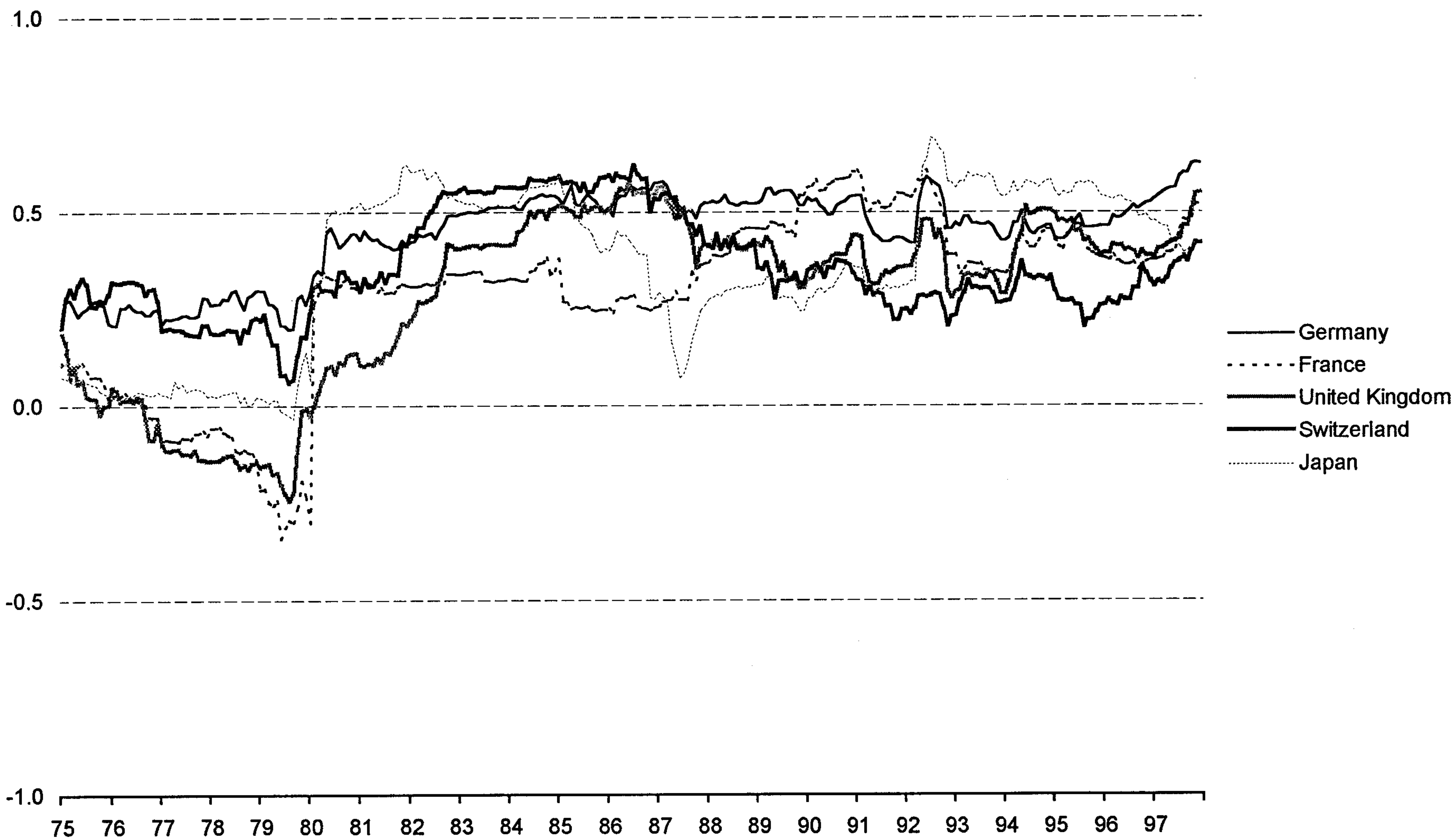
[8]In this respect, Browne and Tease (1992) find little tendency for long-term interest rates to vary systematically with the business cycle.

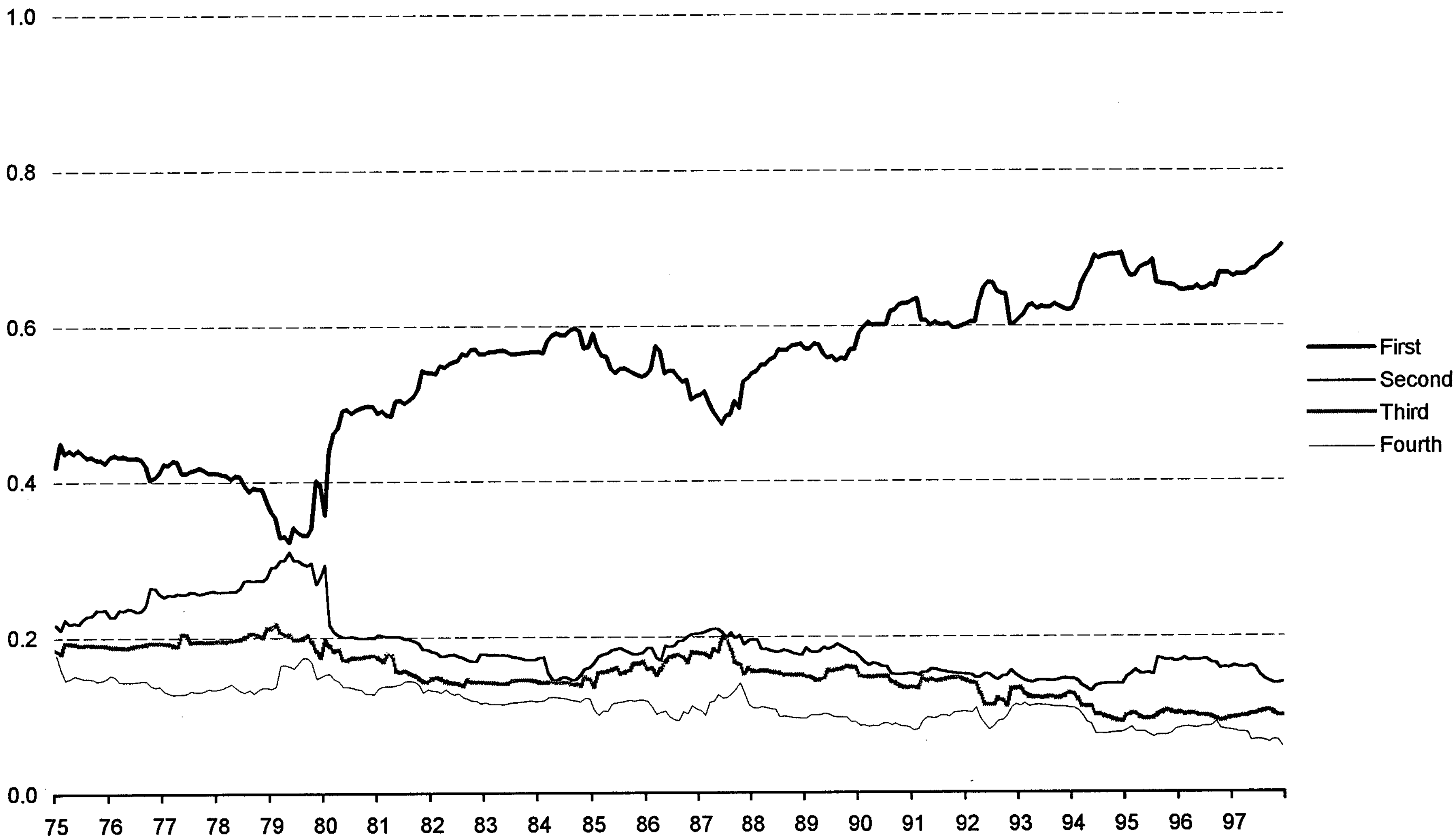
[9]Fase and Vlaar (1998) however suggest that this was mainly a European matter.

Figure 2. Correlations of Movements in Domestic Bond Yields vis-a-vis the U.S.



Source: Authors' own calculations

Figure 3. Proportion of Variance Explained by the First Four Principal Components



Source: Authors' own calculations

III. A Portfolio Model of Interest Rate Differentials

In this section we look at the effects of exchange rate risk on the interest rate differential between two open economies, assuming perfect capital mobility. Following Knot and De Haan (1995) we consider a discrete-time two-asset model in which representative investors have to decide whether to invest in domestic or in foreign securities. While the nominal return on both assets is known with certainty, the real return will be subject to the risk emanating from the possibility of exchange rate adjustments.

A domestic investor is assumed to maximize a utility function defined over the conditional expectation (E) and variance (Var) of real wealth in the next period (w_{t+1}):

$$Max\ U\{E_t[w_{t+1}],\, Var_t[w_{t+1}]\}, \qquad U_1>0,\ U_2<0. \tag{1}$$

Furthermore, he allocates a fraction θ of his initial nominal wealth W_t to the domestic security DB, and a fraction $(1-\theta)$ to the foreign security FB, so that:

$$\begin{aligned} \theta_t W_t &= DB_t \\ (1-\theta_t)W_t &= S_t FB_t, \end{aligned} \tag{2}$$

where S_t represents the initial price of foreign currency measured in units of domestic currency. Moreover, assuming that the foreign country is the country with the lower inflation rate, its price level is normalized to one so that—imposing some form of relative purchasing power parity—the exchange rate varies proportionally with the domestic price level P. Hence, $P=S$ and next period real wealth w_{t+1} equals nominal wealth divided by S_{t+1}:

$$w_{t+1} = \frac{(1+r_t)\theta_t W_t}{S_{t+1}} + \frac{(1+r_t^*)(1-\theta_t)W_t}{S_t}, \tag{3}$$

where $1+r_t$ is the gross nominal rate of return during the time to maturity and an asterisk indicates a corresponding foreign variable.

Likewise, the objective function of a representative foreign investor can be written as:

$$Max\ U^*\{E_t[w_{t+1}^*],\ Var_t[w_{t+1}^*]\}, \qquad U_1^*>0,\ U_2^*<0. \tag{4}$$

To this end he allocates a fraction θ^* of his initial nominal wealth W_t^* to the domestic security DB, and a fraction $(1-\theta^*)$ to the foreign security FB:

$$\begin{aligned} \theta_t^* W_t^* &= \frac{DB_t^*}{S_t} \\ (1-\theta_t^*) W_t^* &= FB_t^*. \end{aligned} \tag{5}$$

Next period real wealth for foreign investors can thus be written as:

$$w_{t+1}^* = \frac{(1+r_t)\theta_t^* W_t^* S_t}{S_{t+1}} + (1+r_t^*)(1-\theta_t^*)W_t^*. \tag{6}$$

The exchange rate in the next period is assumed to be determined by:

$$\frac{1}{S_{t+1}} = \frac{1+\eta_{t+1}}{S_t}, \tag{7}$$

where η is a stochastic variable distributed with expected value $E_t[\eta_{t+1}]$ and variance $Var_t[\eta_{t+1}]$. If η_{t+1} is positive, the domestic currency appreciates in value, whereas a negative value of η_{t+1} points to a depreciation.

Combining equations (1), (3), and (7), it can readily be shown that the optimal share of initial wealth placed in domestic securities by domestic investors is given by (see Appendix I):

$$\theta_t = \frac{r_t - r_t^* + (1+r_t)E_t[\eta_{t+1}]}{\phi(1+r_t)^2 Var_t[\eta_{t+1}]}, \tag{8}$$

where $\phi = -2(W/S)U_2/U_1 > 0$ denotes the coefficient of relative risk aversion for domestic investors which is assumed to be constant. Similarly, foreign investors find it optimal to invest a share θ^* of initial wealth in the domestic security, which can be expressed as:

$$\theta_t^* = \frac{r_t - r_t^* + (1+r_t)E_t[\eta_{t+1}]}{\phi^*(1+r_t)^2 Var_t[\eta_{t+1}]}, \tag{9}$$

where the coefficient of relative risk aversion for foreign investors is represented by $\phi^* = -2W^*U_2^*/U_1^* > 0$, which is also assumed to be constant. The equilibrium condition for the domestic securities market can now be expressed by:

$$SB = \frac{r_t - r_t^* + (1+r_t)E_t[\eta_{t+1}]}{(1+r_t)^2 Var_t[\eta_{t+1}]}\left(\frac{W_t}{\phi} + \frac{S_t W_t^*}{\phi^*}\right), \qquad SB > 0, \tag{10}$$

where SB denotes the supply of domestic securities assumed to be constant, and the right-hand side of (10) represents the total demand for the domestic security, DB+DB*. Solving for a stable equilibrium (stable in the sense that $\partial(DB+DB^*)/\partial r > 0$) yields:

$$r_t - r_t^* = -(1+r_t)E_t[\eta_{t+1}] + \frac{SB(1+r_t)^2 Var_t[\eta_{t+1}]}{W_t/\phi + S_t W_t^*/\phi^*}, \quad for \quad 1+r_t < \frac{2(1+r_t^*)}{1+E_t[\eta_{t+1}]}. \tag{11}$$

From equation (11) we can infer that in addition to the uncovered interest parity condition $r_t+(1+r_t)E_t[\eta_{t+1}] = r^*_t$ there exists a (country-specific) risk premium, which is non-linear in the nominal rate of return, and further depends on the supply of domestic securities relative to the total amount of domestic as well as foreign wealth, the uncertainty concerning the expected exchange rate movement $Var_t[\eta_{t+1}]$, and the degree of risk aversion of market participants. Risk-averse investors require a higher premium to invest in a security with uncertain real return, that is the domestic security.

Thus far we have only considered a maturity of one period, without further specification. For a maturity of k years, equation (11) may be rewritten as:

$$i_t^k - i_t^{k,*} = (1+ki_t^k)\frac{E_t[s_{t+k}] - s_t}{k} + \rho_t, \quad where \ \rho_t = \frac{SB(1+ki_t^k)^2 Var_t[\eta_{t+1}]}{k(W_t/\phi + S_t W_t^*/\phi^*)}, \tag{12}$$

s_t denotes the natural logarithm of S_t, $\eta_{t+k} = s_t - s_{t+k}$, and $i_t^k = r_t/k$ denotes the (annualized) nominal interest rate on a financial instrument with a remaining life of k years.

IV. DATA AND METHODOLOGY

Following the seminal work of Engle, Lilien and Robins (1997) we allow for a time-varying risk premium in our empirical work by estimating the UIP-relationship (12) as the conditional mean specification in a ARCH-in-mean model. In order to preserve a reasonable and parsimonious specification for the conditional variance term, we set the order of the ARCH-process equal to 1 (month), a choice that performed reasonably well in the empirical analysis below.[10] We therefore estimate:

$$i_t^k - i_t^{k,*} = \alpha + \beta\left(\frac{E_t[s_{t+k}] - s_t}{k}\right) + \gamma \rho_t + \epsilon_t$$
$$\rho_t^2 = \delta + \psi \epsilon_{t-1}^2, \tag{13}$$

where, conditional on past information, the residuals ϵ_t are distributed normally with mean zero and variance ρ_t^2. In this specification, time-varying risk premia are proxied by the conditional standard deviation ρ_t that directly affects the interest rate differential, with an elasticity of γ. The specification of the conditional standard deviation conforms with a very simple model in which financial market participants predict this period's variance based on information about previous volatility. An unexpectedly large interest rate differential will increase the estimated variance for the next period.

We test for uncovered interest rate parity between the United States and five other countries, Germany, France, the United Kingdom, Japan, and Switzerland, by regressing the long-term interest rate differential between the United States and the foreign country in question on an intercept, the expected rate of depreciation of the U.S. dollar against their currency, and a time-varying risk premium. The treatment of the U.S. dollar as the domestic currency in our model is consistent with the observation that for most of the 'foreign' countries involved, inflation has been lower than in the United States (primary exceptions here being the United Kingdom and, until the early 1980s, also France). In our model it also implies that investors require a time-varying risk premium on U.S. securities, which might be reconciled with the United States' international status of a net borrowing country during most of the sample period. Intercepts are included to take into account a number of country-specific characteristics of the financial system that are less time-varying, such as for instance the liquidity of the U.S. bond markets vis-à-vis those in the other countries, or differences in capital income taxation (McCallum, 1994).

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[10]We also investigated more general (G)ARCH processes, but additional ARMA terms were generally insignificant and, hence, did not materially affect the estimated parameter of the risk premium either.

In general, only the interest rates are observable in equation (13). To overcome the identification problem with respect to the expected depreciation term, most authors resort to the rational expectations hypothesis (REH) to justify their use of actual exchange rate outcomes to proxy for unobservable expectations. Under this assumption, however, UIP can only be tested in conjunction with REH; in case of a subsequent rejection it remains unclear which leg of the joint hypothesis is not being supported by the data. Moreover, it has often been documented that participants in the foreign exchange market do make systematic forecast errors and do not always use all information efficiently when forecasting, thereby rejecting the rationality of these forecasts (Frankel and Froot, 1987, Takagi, 1991, Cavaglia et al., 1993, and Sobiechowski, 1996). In our view, the longer the horizon of the investment in question, the more debatable this practice of employing actual realizations becomes.

Instead, we propose a process of long-term expectations formation in which the exchange rate is assumed to be ultimately guided by its 'fundamental anchor' of purchasing power parity (PPP). Underlying this assumption is the notion that investors, while acknowledging the possibility that exchange rates may be misaligned for a couple of months, quarters, or even years, will regard the persistence of this phenomenon over the full ten-year term as rather unlikely. Questioning 200 chief London forex dealers, Allen and Taylor (1989) indeed found that at forecast horizons of one year and longer, nearly 30 percent of the respondents relied on pure fundamentals and 85 percent judged fundamentals to be more important than swings in market psychology as captured by so-called 'chartism'. Moreover, recent studies that either use long-horizon or panel data sets have been quite supportive of the notion that real exchange rates are in fact mean-reverting, with an estimated half-life for deviations from PPP of about three to four years (Abuaf and Jorion, 1990, Froot and Rogoff, 1995, and Wu, 1996). For a ten-year horizon it therefore seems plausible to expect the exchange rate to revert to 'current wisdom' regarding its long-run equilibrium, as measured by then prevailing estimates of PPP exchange rates.[11]

For this purpose, we gathered annual data on purchasing power parity exchange rates vis-à-vis the U.S. dollar from the OECD National Accounts database, that have been interpolated into monthly PPP observations based on developments in relative consumer

[11]Ideally, one should use the expected PPP rate at maturity as the anchor guiding exchange rate expectations under PPP. The difference between the expected PPP rate k years ahead, $E_t[s^{ppp}_{t+k}]$, and the current PPP rate, s^{ppp}_t, will normally correspond to the expected cumulative inflation differential between the United States and the foreign country in question over that time interval, $E_t[\pi_{t,t+k}-\pi^*_{t,t+k}]$, so that the latter expectation should be regarded as a missing variable in our regressions (recall $E_t[s^{ppp}_{t+k}]-s_t = E_t[s^{ppp}_{t+k}]-s^{ppp}_t + s^{ppp}_t-s_t = E_t[\pi_{t,t+k}-\pi^*_{t,t+k}] + s^{ppp}_t-s_t$). In terms of the empirical implementation, however, it is not so straightforward to come up with a proxy that would adequately capture this ten-year ahead expected inflation differential.

prices employing a method proposed by Boot, Feibes and Lisman (1967).[12] Averaged over the year the movement in the resulting series is identical to the original OECD series while movements within the year reflect movements in the CPI of the United States relative to the foreign country in question, which is in line with the nature of purchasing power parity. Annualized 'expected' rates of depreciation are then calculated by dividing the deviation of the actual exchange rate from its PPP-value by the term of the investment (ten years).

V. Empirical Results

Table 2 first presents estimates of the risk-adjusted UIP model (13) with respect to the entire sample, using maximum likelihood with heteroskedasticity-consistent standard errors (Bollerslev and Wooldridge, 1992).

Upon investigating Table 2, four features stand out. First, while four out of five parameters of the expected rates of depreciation have the expected positive sign—Japan being the single exception—for none of the countries is the coefficient near unity, as would be implied by strict UIP.[13] Second, despite the greater liquidity of the U.S. bond markets, the interest rate differentials against Germany, Switzerland, and Japan display positive intercepts. Switzerland and Japan are characterized by a relatively closed financial system, which may be have kept domestic bond yields artificially low.[14] Third, while the expected positive ARCH-in-mean elements are indeed documented for the interest rate differentials against Germany and Switzerland, negative risk premia emerge for the differentials vis-à-vis France, the United Kingdom, and Japan (albeit for the latter insignificantly so). In both France and the

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[12]The two-step procedure is as follows. First, by minimizing the sum of squares of the second differences a monthly series is interpolated, ensuring smoothness whilst preserving the movement of the original series. Subsequently, monthly observations on PPP are constructed as the fitted values of a regression of the smoothly interpolated monthly series on the quotient of monthly movements in the CPI between the United States and the foreign country.

[13]The fact that our PPP-based measure of expected depreciation is measured with error (see footnote 11) would in itself tend to bias downward its coefficient.

[14]Switzerland has often been referred to as an 'interest rate island', owing to the stylized facts that real interest rates have been lower in Switzerland than in most other industrialized countries and that Swiss interest rates seem to be decoupled from world interest rates to a greater extent than in other countries. These features may relate to an excess of national savings over investment for many years, that has resulted in a persistent current account surplus and the largest ratio of net foreign assets to GDP in the world (Mauro, 1995). Also the role of the Swiss franc as a 'safe haven' currency could have played a role in explaining the results for Switzerland.

Table 2. Testing for UIP: Entire Sample 1975-1997

	Germany	France	United Kingdom	Switzerland	Japan
Intercept	0.022	-0.008	-0.006	0.037	0.026
	(74.4)	(16.1)	(14.7)	(93.7)	(23.4)
PPP-based expected	0.705	0.175	0.269	0.342	-0.084
depreciation	(91.5)	(18.6)	(17.42)	(36.7)	(5.2)
Time-varying risk	0.116	-0.141	-0.606	0.273	-0.095
premium	(3.0)	(2.7)	(17.0)	(10.22)	(1.4)
Variance Equations					
Intercept	5.39 E-6	8.09 E-6	12.9 E-6	3.35 E-6	14.5 E-6
	(4.6)	(4.6)	(4.3)	(4.3)	(2.7)
ARCH(1)	1.053	1.064	0.975	1.279	0.944
	(18.1)	(13.6)	(18.7)	(28.0)	(18.0)
Diagnostics					
R^2-adjusted	0.37	0.02	0.16	0.19	0.01
LM[ARCH-8]	1.06	2.27	1.34	1.97	1.55
(p-value)	(0.40)	(0.03)	(0.22)	(0.05)	(0.14)
Q(32)	35.51	63.38	66.35	42.18	30.07
(p-value)	(0.31)	(0.00)	(0.00)	(0.11)	(0.57)
LM[NORM]	10.77	3.21	24.76	25.04	14.67
(p-value)	(0.01)	(0.17)	(0.00)	(0.00)	(0.00)
ADF tests on standardized residuals					
3 unit roots	23.59**	25.68**	26.38**	25.22**	24.48**
	(C,4)	(C,4)	(C,4)	(C,4)	(C,4)
2 unit roots	9.14**	9.92**	11.02**	11.50**	9.43**
	(C,4)	(C,4)	(C,4)	(C,4)	(C,4)
1 unit root	5.52**	7.15**	5.78**	5.20**	4.63**
	(N,2)	(N,2)	(N,1)	(N,1)	(N,1)
Conclusion	I(0)	I(0)	I(0)	I(0)	I(0)

Source: Authors' own calculations.

Notes: Absolute t-values, computed with heteroskedasticity consistent standard errors are in parentheses. LM[ARCH-8] is Lagrange-Multiplier test for ARCH-8; Q(32) is Ljung-Box Q-test; LM[NORM] is Jarque-Bera test. Critical values for an ADF-test on the residuals of a cointegration relation with 3 non-stationary variables are 4.35 (1%) and 3.78 (5%), respectively (Engle and Yoo, 1987). Significance of the ADF tests at the 5% (*) and 1% (**) level is indicated. The information within parentheses relates to the deterministic part of the ADF equation. C=Constant; N=No Constant; the number indicates the lags included.

United Kingdom inflation has not been consistently lower than in the United States (as was assumed for the 'foreign' country in our model), and the current account has swung into deficit every now and then. Finally, despite the restrictive character of the UIP plus ARCH-in-mean specification for a decomposition of interest rate differentials, the equations are reasonably well-behaved—aside from some excess kurtosis.

The lower panel of Table 2 reports on a Engle and Granger (1987) type cointegration test for the UIP regressions estimated. From an investigation of the time-series characteristics of individual interest rate differentials and PPP-based expected rates of depreciation (not shown), no clear-cut picture across countries emerged as to the order of integration of these variables. Additionally, the estimates of ψ in the variance equations of Table 2 reveal a good deal of persistence in our measure of the risk premium, possibly suggesting non-stationarity for this generated regressor as well. We therefore proceed by investigating the stationarity properties of the standardized residuals of the risk-adjusted UIP-model (13).[15] Even if some of the individual series contain a unit root, stationary residuals would point to the existence of cointegration between the three stochastic variables in the risk-adjusted UIP model. From the final row of Table 2 it appears that for all five countries under investigation, non-stationarity of the residuals can safely be rejected, thereby validating our levels-based specification in (13).

In order to investigate whether the apparent increases in co-movement of bond yields across the decades—suggested in Section II—can be traced back to a declining prevalence of exchange rate misalignments, Table 3 reports on subsample estimates of the risk-adjusted UIP model (13). Since the advent of new financial instruments and the abolition of capital controls increased long-term capital mobility, our hypothesis is that in more recent decades there should be stronger evidence in favor of UIP than in earlier decades.

Table 3 contains a mixed bag of results.[16] For a start, empirical support for the UIP hypothesis has clearly increased in the 1980s compared to the 1970s. While the five estimated regression coefficients for the 1970s point at uncovered interest rate differentials moving consistently in the opposite direction to that implied by PPP-based exchange rate forecasts,

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[15]This standardization, common in the evaluation of (G)ARCH models, implies dividing the residuals from the model estimated in Table 2 by their conditional standard deviation, and ensures that the residuals are identically and independently distributed, with zero mean and unit variance. The critical values for the residuals-based test for cointegration depend on the number of I (1) variables (in this case, three) in the cointegrating regression (Engle and Granger, 1987, and Engle and Yoo, 1987).

[16]For the sake of parsimonity the diagnostics have been omitted from Table 3. In general, most diagnostics were clearly supportive of the risk-adjusted UIP specification (13) across the decades, except (again) for the presence of excess kurtosis in France and Switzerland during the 1980s and Japan during the 1990s.

Table 3. Testing for UIP across the Decades

	Germany	France	United Kingdom	Switzerland	Japan
a) 1975-1979					
Intercept	-0.021	-0.052	-0.038	0.001	-0.005
	(41.6)	(11.5)	(32.7)	(1.4)	(2.8)
PPP-based expected	-1.227	-0.961	-0.899	-1.380	-0.680
depreciation	(47.6)	(15.0)	(19.5)	(71.5)	(22.6)
Time-varying risk	-0.126	1.847	-0.234	-0.128	-0.063
premium	(1.7)	(2.0)	(1.9)	(2.4)	(0.2)
Variance Equations					
Intercept	3.37 E-6	16.1 E-6	16.8 E-6	0.246 E-6	13.1 E-6
	(2.8)	(5.1)	(2.8)	(2.8)	(2.9)
ARCH(1)	1.256	0.384	1.121	1.483	0.599
	(3.7)	(2.2)	(3.7)	(4.4)	(3.5)
b) 1980-1989					
Intercept	0.029	-0.007	-0.005	0.061	0.045
	(55.5)	(6.2)	(7.1)	(84.6)	(43.9)
PPP-based expected	0.317	0.174	0.265	0.344	0.066
depreciation	(23.1)	(9.8)	(14.2)	(19.5)	(3.6)
Time-varying risk	0.072	-0.239	-0.246	0.200	-0.303
premium	(0.8)	(2.2)	(3.2)	(3.3)	(3.6)
Variance Equations					
Intercept	9.23 E-6	16.0 E-6	14.3 E-6	10.4 E-6	13.2 E-6
	(3.2)	(2.8)	(4.0)	(3.0)	(3.3)
ARCH(1)	1.034	0.955	0.991	1.027	1.038
	(7.9)	(7.9)	(9.6)	(13.2)	(9.1)
c) 1990-1997					
Intercept	0.001	-0.005	-0.019	0.030	-0.004
	(1.8)	(7.3)	(54.6)	(24.1)	(3.9)
PPP-based expected	0.142	0.527	0.401	0.148	-0.517
depreciation	(6.5)	(14.3)	(16.3)	(6.6)	(33.9)
Time-varying risk	0.002	0.477	0.394	0.404	0.292
premium	(0.1)	(3.5)	(5.1)	(3.9)	(4.4)
Variance Equations					
Intercept	2.52 E-6	3.44 E-6	4.08 E-6	4.95 E-6	3.51 E-6
	(3.7)	(3.1)	(3.6)	(3.6)	(3.2)
ARCH(1)	1.020	1.039	0.958	0.863	1.086
	(10.4)	(9.3)	(7.6)	(8.8)	(8.7)

Source: Authors' own calculations.
Notes: See Table 2.

these parameters all adopt the expected positive sign during the 1980s. Nonetheless, there is no discernable tendency for these coefficients to further increase from the 1980s to the current decade. If anything, three out of five are *de*creasing. Only for France and the United Kingdom does Table 3 reveal the expected upwardly sloping pattern across the decades. For the other countries, the data seem to reject the notion of steady increases in UIP-validation over time.

All in all, one cannot simply conclude that our evidence supports the validity of UIP among all 'industrialized' currencies. It nevertheless appears that from the 1980s onward and with the exception of Japanese securities,[17] uncovered interest rate differentials consistently adjust in the same direction as suggested by expected rates of depreciation, albeit less than one-for-one. In this respect, our strategy of employing PPP-based data instead of ex-post realizations to proxy for (unobservable) exchange rate expectations constitutes a clear improvement over the various studies summarized in Froot (1990) and McCallum (1994).

In order to account for the relatively favorable results for the more recent decades, two explanations come to mind. One possibility is that the improvement is due to the rejection of the joint hypothesis of rational expectations in the specific form underlying the studies referred to above, as participants in the foreign exchange market do make systematic forecast errors and do not always use all information efficiently when forecasting (Frankel and Froot, 1987, Takagi, 1991, Cavaglia et al., 1993, and Sobiechowski, 1996). Alternatively, one could also point at the differential maturity of the assets under investigation. McCallum (1994) and Meredith and Chinn (1998) have shown that for relatively short horizons, failure of UIP may result from temporary risk-premium shocks in the face of an endogenous monetary policy that is characterized by 'leaning against the wind' of prevailing exchange rate movements.[18] In the longer term, in contrast, exchange rates are driven by 'fundamentals', leading to a relation between interest rate differentials and exchange rates that should be more consistent with UIP.

To put the competing explanations to a more formal test, Table 4 replicates the middle panel of Table 3, employing REH-based expected rates of depreciations instead of the PPP-based ones used before. Given the lagged availability of ten-year ahead REH-outcomes, the most recent observation on the process of interest rate arbitrage under REH is March 1989.

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[17]The results for Japan may have been adversely affected because our method of gauging exchange rate expectations from current PPP rates does not take account of trends in PPP, such as the long-run upward trend that Japan has witnessed (see also footnote 11). There is also evidence that equilibrium exchange rates can be influenced by trends in overseas net assets, which may also have been potentially important for Japan.

[18]Unfortunately, the backward-looking specification of the time-varying risk premium in our model (13) is unable to directly pick up the type of contemporaneous risk-premium shocks that contaminate short-horizon tests of UIP.

Table 4. Testing for UIP with Rational Expectations: 1980 to March 1989

	Germany	France	United Kingdom	Switzerland	Japan
Intercept	0.025	-0.008	-0.003	0.048	0.042
	(33.8)	(8.6)	(4.6)	(82.0)	(59.9)
REH-based expected	0.190	0.188	0.325	0.290	0.043
depreciation	(14.9)	(16.0)	(14.7)	(24.6)	(3.9)
Time-varying risk	-0.101	-0.574	-0.169	0.173	-0.289
premium	(0.8)	(6.4)	(1.9)	(2.8)	(3.2)
Variance Equations					
Intercept	12.2 E-6	11.2 E-6	17.8 E-6	8.1 E-6	11.9 E-6
	(3.3)	(3.6)	(3.1)	(3.3)	(3.2)
ARCH(1)	0.902	1.101	0.841	1.041	1.039
	(8.2)	(7.8)	(8.2)	(12.4)	(8.6)

Source: Authors' own calculations.
Notes: See Table 2.

Table 4 displays a striking similarity between the UIP estimates obtained for REH-based expected rates of depreciation and those obtained for their PPP-based counterparts. Despite a small difference in sample period, all β-parameters but one (Germany) diverge by less than one-tenth from the corresponding estimates of the middle panel of Table 3, with no consistent pattern as to the direction of the deviation. Hence, it is fair to conclude that the relatively favorable results reported above are probably due to the longer horizon of the securities under investigation rather than to the use of an empirically 'superior' measure of exchange rate expectations.[19] In the process of asset valuation with the aid of UIP, the value added of PPP-based expected rates of depreciation would then predominantly be a matter of data availability. While the REH-based 'expected' rates of depreciation underlying Table 4 can only be inferred ex-post, our PPP-based measures are readily available at the time investors have to decide on the diversification of their portfolios.

VI. Concluding Remarks

Empirical tests of the uncovered parity relationship traditionally used short-term interest rates, because until the mid 1980s hardly any financial instruments existed that enabled market participants to hedge for long-term interest rate positions. This situation changed in the 1980s, when cross-currency interest rate swaps created the possibility of hedging longer-term currency positions (Popper, 1993). Arbitrage conditions which previously applied only to short-term assets now facilitate long-term financial capital to become international mobile across distinct bond markets.

This paper has investigated the uncovered interest parity relationship between bond yield differentials for the United States vis-à-vis Germany, France, the United Kingdom, Switzerland, and Japan, with the aid of PPP-based exchange rate expectations against the U.S. dollar. Globalization and integration of financial markets would seem to imply a tendency toward co-movement of bond yields, as reflected in a greater likelihood of validating the UIP relation. A visual inspection of the major swings in international bond markets would indeed suggest such a tendency. However, we could not find empirical evidence supporting the notion of further increases in co-movement beyond that associated with the wave of financial market liberalization and deregulation in the early 1980s.

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[19]Using REH-based expected rates of depreciation over the period 1973-88, Meredith and Chinn (1998) also report quite strong support for UIP at the ten-year horizon. Owing to the 'inverse' specification of the UIP-relation, the absence of a time-varying risk premium, and the use of quarterly instead of monthly data, caution has to be exercised when comparing results. Meredith and Chinn (1998) report β-parameters that are either insignificantly different from one (Canada, France, and Germany), or at least significantly positive but smaller than one (Italy, Japan, and the United Kingdom). Translated to an inverse specification such as (13), however, the latter estimates would imply that changes in exchange rate expectations will be reflected in adjustments to interest rate differentials in a *more*-than-one-for-one fashion.

Nonetheless, the relatively strong support we documented for at least partial validation of the UIP relation suggested that our strategy of employing PPP-based forecasts instead of ex-post realizations to proxy for (unobservable) exchange rate expectations might constitute a clear improvement over the existing literature. Further investigation revealed, however, that our relatively favorable results were due to the longer horizon of the securities under investigation rather than to the use of an empirically 'superior' measure of exchange rate expectations. The value added of PPP-based expected rates of depreciation for this purpose mainly lies with their ready availability at the time investors have to allocate their funds across assets of variable denomination.

APPENDIX I: DERIVATION PORTFOLIO COMPOSITIONS (8) AND (9)

Next period real wealth of domestic investors can be written as:

$$w_{t+1} = \frac{(1+r_t)\theta_t W_t}{S_{t+1}} + \frac{(1+r_t^*)(1-\theta_t)W_t}{S_t}, \tag{A1}$$

so that

$$E_t[w_{t+1}] = (1+r_t)\theta_t W_t E_t[S_{t+1}^{-1}] + (1+r_t^*)\frac{(1-\theta_t)W_t}{S_t}, \tag{A2}$$

which, by means of equation (7), can be rewritten as:

$$E_t[w_{t+1}] = (1+r_t)(1+E_t[\eta_{t+1}])\frac{\theta_t W_t}{S_t} + (1+r_t^*)\frac{(1-\theta_t)W_t}{S_t}. \tag{A3}$$

Furthermore:

$$\begin{aligned}
Var_t[w_{t+1}] &= E_t[w_{t+1} - E_t[w_{t+1}]]^2 \\
&= E_t\left[(1+r_t)\frac{\theta_t W_t}{S_{t+1}} - (1+r_t)(1+E_t[\eta_{t+1}])\frac{\theta_t W_t}{S_t}\right]^2 \\
&= (1+r_t)^2\theta_t^2 W_t^2\left\{E_t[S_{t+1}^{-2}] - \frac{2(1+E_t[\eta_{t+1}])}{S_t}E_t[S_{t+1}^{-1}] + \frac{(1+E_t[\eta_{t+1}])^2}{S_t^2}\right\} \\
&= \frac{(1+r_t)^2\theta_t^2 W_t^2}{S_t^2}\{E_t[\eta_{t+1}^2] - E_t^2[\eta_{t+1}]\} \\
&= \frac{(1+r_t)^2\theta_t^2 W_t^2}{S_t^2} Var_t[\eta_{t+1}].
\end{aligned} \tag{A4}$$

Utility maximisation implies $\partial U/\partial\theta = 0$:

$$U_1 \frac{\partial E_t[w_{t+1}]}{\partial \theta} + U_2 \frac{\partial Var_t[w_{t+1}]}{\partial \theta} =$$

$$U_1 \frac{W_t}{S_t}\{r_t - r_t^* + (1+r_t)E_t[\eta_{t+1}]\} + U_2 2\theta(1+r_t)^2 \frac{W_t^2}{S_t^2} Var_t[\eta_{t+1}] = 0, \tag{A5}$$

from which it can easily be seen that:

$$\theta_t = \frac{r_t - r_t^* + (1+r_t)E_t[\eta_{t+1}]}{\phi(1+r_t)^2 Var_t[\eta_{t+1}]}, \quad where\ \phi = -2\frac{W_t}{S_t}\frac{U_2}{U_1}. \tag{A6}$$

Likewise:

$$E_t[w_{t+1}^*] = (1+r_t)\theta_t^* W_t^*(1+E_t[\eta_{t+1}]) + (1+r_t^*)(1-\theta_t^*)W_t^*, \tag{A7}$$

$$Var_t[w_{t+1}^*] = (1+r_t)^2 \theta_t^{*2} W_t^{*2} Var_t[\eta_{t+1}], \tag{A8}$$

$$\theta_t^* = \frac{r_t - r_t^* + (1+r_t)E_t[\eta_{t+1}]}{\phi^*(1+r_t)^2 Var_t[\eta_{t+1}]}, \quad where\ \phi^* = -2W_t^* \frac{U_2^*}{U_1^*}. \tag{A9}$$

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