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UNCOVERED INTEREST PARITY REVISITED

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Abstract

A standard empirical finding is that international interest rate differentials are negatively correlated with ex post exchange rate changes. However, tests of uncovered interest parity have almost exclusively relied on data on short interest rates. In this paper, a standard test of uncovered interest parity is performed using data on long interest rates in 14 OECD countries versus the United States. In contrast to the typical finding, the test results are rather favourable to UIP.

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1. Introduction

According to the stylised facts, the exchange rates of countries with high nominal interest rates tend to appreciate instead of depreciate as expected from uncovered interest parity (UIP). Not only does UIP not hold in empirical tests, but the coefficient β from regressing nominal exchange rate changes on interest differentials is negative, large and significant. This negative relationship has been confirmed by many authors. McCallum (1994) surveys the empirical literature and concludes that β typically equals minus three. According to Engel (1996), minus three to minus four is a representative result. Froot and Thaler (1990) calculate the average β over a large number of studies and find it to be -0.88.[1] Uncovered interest rate parity remains a key assumption in international macroeconomic modelling in spite of the massive body of empirical evidence against the hypothesis.

However, virtually every published test of UIP relies on data on short interest rates. Before elevating the empirical failure of UIP to a stylised fact, long interest rates should also be investigated. Arguably, a monetary policy aimed at affecting the nominal exchange rate by manipulating the short interest rate may produce the negative relationship discussed above (see Mc Callum, 1994). Exchange rate movements are also difficult to predict in the short run, while the presumed existence of long run equilibrium real exchange rates lend at least some predictability to long run movements in nominal exchange rates given inflation forecasts.

There appears to be only two papers studying UIP for data on long interest rates, both of which reject the hypothesis. DeGennaro et al. (1994) study cointegration

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[1] Recent evidence on UIP may not be quite as conclusive as these surveys suggest, see Ngama (1994), Solocha and Saidi (1995), Turtle and Abeysekara (1996), Ayuso and Restoy (1996), Guin and Maxwell (1996) and Moosa and Bhatti (1997) for more favourable results. All these papers study short interest rates.

among long interest rates for five major nations and reject UIP. De Haan et al. (1991) focus on Dutch and German long interest rate differentials and conclude that they are stationary with shifts in mean, which is inconsistent with UIP given the behaviour of exchange rates. Both papers use short sample periods and neglect the issue of coupon payments. A main reason for this striking neglect of UIP for long interest rates is probably found in the poor quality of data on long term government bond yields (especially when compared to data on short interest rates, where precision is literally minute). However, even poor quality data may contain interesting information that should not be neglected.

In this paper, UIP is tested for long term government bond yields and nominal exchange rates in 14 countries against the US. The sample period is 1957:1 to 1997:4 and the data are collected from the IFS data tape. There are two major problems with the data set. First, the maturity of the long interest rates is vaguely specified. When stated, it is ten years. However, the lack of exact maturities may not constitute quite as large a problem as it appears on first sight. The difference between long interest rates is usually small, i.e. the yield curve is rather flat at long maturities. For instance, the difference between an eight year bond and a ten year bond depends on expected events between eight years into the future and ten years into the future (as well as on a possible risk premium). It seems unlikely that the market has enough information about events from eight to ten years ahead for this to have a major impact on interest rates. Parallel shifts dominate over slope changes in this segment of the yield curve. It also turns out that the qualitative test results to not change as the assumed maturity is varied from eight to twelve years.

The second data problem is that the presence of coupon payments induce a measurement error between observed yields to maturity and true returns to investments. Various methods to deal with this problem, i.e. to capture the discount bond yield, are available (see Dalhquist and Svensson (1996)). Most techniques require more observations of the yield curve in each time period than what is available here. In most cases, the only known interest rate in addition to the ten year bond yield is a three month interest rate.[2] Better information about the yield curves can be found only from the mid to late 1980:s. The attempts to remove the effects of coupon payments are therefore necessarily rather crude.

In the first set of UIP tests, the original series on yields to maturity are used. Disregarding coupons is correct if the yield curve is flat and the bond is traded at par. Second, the sizes of the coupon payments are approximated and their present values are removed from the (also unobserved) coupon bond prices. A second order polynomial function for the term structure is approximated from the available points on the yield curve and the relevant interest rates are used as discount rates. A different approach is to let the coupon payments affect the length of the investment rather than the bond price. Macaulay's measure of duration is a weighted average of the maturities of the principal payments and the coupon payments. In effect, it moves the observed interest rate to the left on the yield curve. The final set of UIP tests is performed on duration data. While measurement errors may still be considerable, they do not appear to pose insurmountable obstacles to exploring this data set.

2. UIP-tests for long term government bond using yields to maturity

The data set includes quarterly data on nominal exchange rates and long term government bond yields from 1957:1 to 1997:4 for 15 OECD countries from the IFS data tape. The sample spans over different degrees of integration of the international capital market as well as different exchange rate regimes. However, a long sample period is essential when testing UIP on long term bond yields.

[2] For The United States and the United Kingdom, one year and three year interest rates are available since 1957 in addition to the three month interest rate and the ten year yield to maturity.

What can be done is to check whether the results for the period up to the mid 1980:s differs from the results for the period thereafter.

UIP states that the expected percentage change in nominal exchange rates (defined as units of domestic currency per unit of foreign currency) between t and $t+\tau$ equals the interest rate differential $r_{t,t+\tau} - r^*_{t,t+\tau}$. The relevant measure of $r_{t,t+\tau}$ is the holding period return from between t to $t+\tau$. Letting y_t denote the yield to maturity, $r_{t,t+\tau}$ in this section equals $(1+y_t)^\tau - 1$.[3] Assuming rational expectations, expected exchange rate changes equals ex post changes plus an error term. The standard test of UIP is whether $[\alpha,\beta]$ in (1) equals $[0,1]$. Alternatively, a constant risk premium is allowed and only the hypothesis that β equals one is tested.

$$\frac{s_{t+\tau} - s_t}{s_t} = \alpha + \beta\left(r_{t,t+\tau} - r^*_{t,t+\tau}\right) + \varepsilon_t . \qquad (1)$$

First, UIP is tested without taking the coupon payments into consideration. This is correct under the assumptions that (i) all bonds are traded at par and (ii) the yield curve is flat. While these assumptions are certainly not strictly fulfilled, it remains to be seen whether the errors are large enough to matter for the results.

When constructing the variables $\frac{s_{t+\tau} - s_t}{s_t}$ and $\left(r_{t,t+\tau} - r^*_{t,t+\tau}\right)$ from quarterly data with τ equal to ten years, there will be 40 overlapping observations. To correct for the induced autocorrelation, Newey and West (1987) robust standard errors are used. Table 1 shows the OLS estimation results[4].

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[3] The yield to maturity is defined from the coupon bond price as $p_t \equiv \sum_{k=1}^{K} \frac{C}{(1+y_t)^{t+k}} + \frac{1}{(1+y_t)^{t+K}}$, where C is the coupon payment and K is the number of coupons.

[4] A number of authors have tested UIP using cointegration techniques under the assumption that the time series on interest differentials and exchange rate depreciations contain unit roots. In this data set, the ADF test for unit roots indicate that the time series are generally I(0).

Table 1: Tests of UIP without correcting for coupon payments.

Regression: $\frac{s_{t+\tau}-s_t}{s_t}=\alpha+\beta\left(r_{t,t+\tau}-r^*_{t,t+\tau}\right)+\varepsilon_t$.

	α	β	$F(\alpha=0,\beta=1)$	$F(\beta=1)$	R^2
Austria	**-0.182**	**0.076**	129.092 (0.000)	120.214 (0.000)	0.185
	[-3.550]	**[2.276]**			
Belgium	-0.088	0.085	31.836 (0.000)	17.302 (0.000)	0.003
	[-0.859]	[0.387]			
Canada	0.098	**-0.178**	868.322 (0.000)	359.450 (0.000)	0.060
	[2.011]	**[-2.863]**			
Denmark	-0.104	0.132	101.691 (0.000)	66.119 (0.000)	0.094
	[-1.249]	[1.236]			
France	0.046	0.232	**5.133 (0.077)**	4.812 (0.028)	0.025
	[0.433]	[0.661]			
Germany	-0.229	0.757	181.444 (0.000)	174.507 (0.000)	0.004
	[-3.937]	[1.081]			
Italy	0.460	-0.085	232.047 (0.000)	115.608 (0.000)	0.017
	[1.868]	[-0.845]			
Japan	-0.323	**0.121**	633.714 (0.000)	625.448 (0.000)	0.254
	[-12.542]	**[3.437]**			
Netherlands	**-0.165**	**0.197**	82.089 (0.000)	62.937 (0.000)	0.135
	[-2.876]	**[1.951]**			
Norway	0.016	-0.122	402.672 (0.000)	116.269 (0.000)	0.037
	[0.181]	[-1.174]			
Spain	0.179	-0.048	393.274 (0.000)	178.716 (0.000)	0.004
	[0.869]	[0.539]			
Sweden	0.144	-0.061	107.310 (0.000)	87.143 (0.000)	0.002
	[1.128]	[-0.533]			
Switzerland	-0.269	0.012	566.310 (0.000)	184.957 (0.000)	0.001
	[-2.893]	[0.162]			
UK	0.084	**0.245**	125.685 (0.000)	101.169 (0.000)	0.306
	[1.551]	**[3.259]**			

t-values using Newey and West standard errors within brackets. Bold numbers denote significance at the five percent level.

p-values of the *F*-tests within parenthesis. Bold numbers in indicate that the null hypothesis is not rejected at the five percent level.

Five of the estimated β:s in Table 1 differ significantly from zero. Four of these are positive and one is negative (Canada). The hypothesis that β equals one can be rejected in all cases. The main observation from Table 1 is however the lack of significant relationships between interest rate differentials and ex post exchange rate changes. Since the presence of measurement errors in the dependent variable tend to bias the regression coefficient toward zero, attempts to reduce the errors by taking coupon payments into consideration may well yield different results.

3. Approximating the present value of the coupon payments

Given that yield curves are usually not flat and may have different slopes in different countries, it is necessary to consider what can be done to remove the effects of coupons payments from the data. The time series contain yields to maturity. The size of the K coupons C is determined as to make

$$\sum_{k=1}^{K} \frac{C}{\left(1+r_{t,t+k}\right)^{t+k}} + \frac{1}{\left(1+r_{t,t+K}\right)^{t+K}} \approx 1 \tag{2}$$

i.e. bonds are traded approximately at par when they are issued. Now, if a complete term structure for discount bonds , i.e. $r_{t,t+k}$ for k=1,...,K is known, the zero coupon bond price p^{nc} is easily found by calculating the discounted value of each coupon and subtracting it from the coupon bond price p^{c}:

$$p^{nc} = p^{c} - \sum_{k=1}^{K} \frac{C}{\left(1+r_{t,t+k}\right)^{t+k}} \tag{3}$$

Given that the bond is traded at par, the per period return on holding the bond to maturity is then defined by

$$\frac{1-p_{nc}}{p_{nc}} = \frac{1}{\left(1+r_{t,t+K}\right)^{t+K}} - 1 \tag{4}$$

A problem here is that a complete term structure for zero coupon bonds is not available. Typically, the only observed points on the yield curve are a 10 year yield to maturity and a three month interest rate. A simple and feasible way to deal with this problem is to fit a curve $r_{t,t+\tau}(\tau)$ as a function of maturity τ to the available observations. Since yield curves are generally not linear, a quadratic function has been chosen.

$$r_{t,t+\tau}(\tau) = a + b\tau + c\tau^2 \tag{5}$$

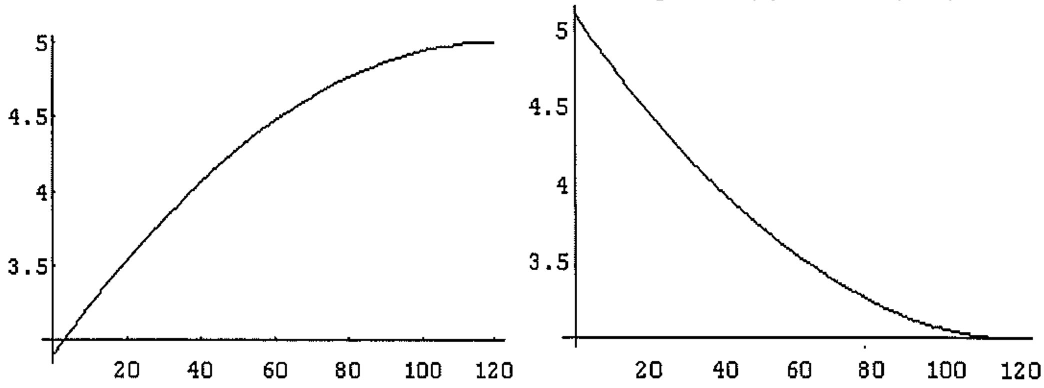
To identify the three parameters in (5), three restrictions are needed. Two are given by letting the curve pass through the two interest rates. Assuming that the yield curve is flat at long maturities yields a third restriction:

$$\begin{aligned} &r(3) = r^{3m}, \\ &r(120) = r^{10y} \text{ and} \\ &\frac{\partial r^{c}(120)}{\partial(\tau)} = 0 \end{aligned} \tag{6}$$

Two examples of the resulting yield curves are shown in Figures 1 a and b. Compared to typical actual yield curves, they are too linear and "flatten out" too far to the right. However, note that this has the opposite effect on the resulting discount bond price compared to assuming a completely flat yield curve as was done in the previous section. In the cases where more than three points on the yield curve are observed, (5) is estimated using OLS.

Figure 1a: An upward sloping yield curve given $r(3)=3$ and $r(120)=5$

Figure 1b: A downward sloping yield curve given $r(3)=5$ and $r(120)=3$



The series on discount bond yields are constructed as follows. Given the assumptions that all bonds are traded at par, the coupon payments are equal to the yield to maturity. The present value of each payments is found by discounting it using the relevant interest rate from (5), i.e. a coupon payments due in t plus three years is discounted using the three year interest rate. The values of the coupon payments are then subtracted from the coupon bond price according to (3) to yield the discount bond price. The holding period return can then be calculated from (4).

Table 2 shows the results from running the regressions after adjusting the returns for coupons as described above. In one case (Norway), the time series on three month interest rates are so short that less than 30 observation remain after adjusting for end points.

Table 2: Tests of UIP when the values of the coupon payments have been approximated and removed from the data.

Regression: $\frac{s_{t+\tau} - s_t}{s_t} = \alpha + \beta\left(r_{t,t+\tau} - r^{*}_{t,t+\tau}\right) + \varepsilon_t$.

	α	β	$F(\alpha = 0, \beta = 1)$	$F(\beta = 1)$	R^2
Austria	**-0.167**	**0.206**	559.392 (0.000)	326.963 (0.000)	0.303
	[-6.593	**[4.702]**			
Belgium	-0.097	0.192	16.997 (0.000)	9.243 (0.002)	0.012
	[-0.991]	[0.722]			
Canada	**0.092**	**-0.207**	173.610 (0.000)	131.735 (0.000)	0.048
	[1.931]	**[-1.972]**			
Denmark	-0.041	0.006	212.2 (0.000)	159.998 (0.000)	0.002
	[-0.214]	[0.215]			
France	0.108	0.232	**3.725 (0.155)**	**3.618 (0.057)**	0.016
	[0.563]	[0.573]			
Germany	**-0.070**	**0.301**	450.596 (0.000)	431.993 (0.000)	0.399
	[-2.380]	**[8.944]**			
Italy	0.213	-0.081	409.053 (0.000)	217.476 (0.000)	0.001
	[1.235]	[-0.118]			
Japan	**-0.323**	**0.131**	586.099 (0.000)	581.463 (0.000)	0.268
	[-12.732]	**[3.640]**			
Netherlands	-0.050	**0.431**	433.058 (0.000)	88.891 (0.000)	0.509
	[-0.933]	**[7.132]**			
Spain	0.125	-0.040	389.494 (0.000)	266.586 (0.000)	0.002
	[0.779]	[-0.627]			
Sweden	0.152	0.036	43.033 (0.000)	39.238 (0.000)	0.001
	[1.042]	[0.232]			
Switzerland	-0.1745	**0.0967**	765.265 (0.000)	234.601 (0.000)	0.079
	[-2.200]	**[1.640]**			
UK	0.092	**0.243**	113.330 (0.000)	96.363 (0.000)	0.326
	[1.784]	**[3.327]**			

t-values using Newey and West standard errors within brackets. Bold numbers denote significance at the five percent level.

p-values of the *F*-tests within parenthesis. Bold numbers in indicate that the null hypothesis is not rejected at the five percent level.

The qualitative test results in Table 2 are rather similar to those in Table 1. The β-coefficients for Germany and Switzerland are now significant. Six of seven significant estimates are positive. However, UIP is still rejected as the parameters are significantly smaller than one. Hence, the relationship between interest rate differentials and exchange rate changes is typically positive, but the exchange rates do not move enough to compensate fully for international differences in interest rates. However, there may still be considerable measurement errors in the dependent variable since the assumptions about the unobservable variable are not strictly fulfilled.

4. Calculating durations

A different approach to the issue of coupon payments is to let them affect the duration of the bond instead of its price. The maturity of a coupon bond is not a good measure of the length of the investment period since a large part of the value of the bond stems from coupon payments that are made before maturity. A coupon bond can be thought of as a package of discount bonds. Macaulay's measure of duration D is a weighted average of the maturities of these discount bonds, where the weights consist of the present value of the payments using the yield to maturity as discount rate:

$$D_t = \left(\frac{C}{(1+y_t)} + \frac{2C}{(1+y_t)^2} + \frac{3C}{(1+y_t)^3} + \ldots + \frac{K(1+C)}{(1+y_t)^K} \right) \frac{1}{p^c} \qquad (7)$$

The duration of the coupon bond is shorter than the maturity, i.e. the observed yield y_t is moved to the left on the yield curve. The effect is larger the larger are the coupon payments. Compared to the previous methods of removing the effects of coupon payments from the data, this method is less dependent on the assumptions about the unobserved bond price. Moving the observation to the left on the yield curve will not imply drastic changes in the yield, especially not as the yield curve is assumed to be rather flat at long maturities.

It is still necessary to approximate yield curves from the existing observations, usually only a three month interest rate and the yield to maturity on a ten year bond. However, the latter is now interpreted as an observation of a bond with a maturity equal to the duration in (7), which is shorter than ten years. The parameters of the constructed yield curve $r_{t,t+\tau}(\tau) = a + b\tau + c\tau^2$ are now identified by the conditions

$$
\begin{aligned}
r(3) &= r^{3m}, \\
r(D_t) &= y_t \text{ and} \\
\frac{\partial r_{t+\tau}(120)}{\partial(\tau)} &= 0
\end{aligned}
\tag{8}
$$

instead of (6). Again, when more points on the yield curve are observed, the parameters are estimated using OLS.

After transforming the maturities of the ten year bonds $(\tau = 10)$ for all observations to time varying durations $\tau_t = D_t$ according to (7), matching data on exchange rate changes are needed. For each observation, there is one duration for the home country and another duration for the foreign country. The change of the exchange rate between t and τ can obviously only be calculated for a single τ. Here, the integer closest to the *average* of the two durations have been used to construct a matching exchange rate change. Since this procedure uses home and foreign interest rates different from the observed yields, it is still necessary to construct coupon yield curves along which to move.

Each observation is constructed by i) calculating home and foreign durations from (7); ii) taking the average of the two durations and iii) calculating home and foreign bond yields for this average duration by moving along the yield curves given by (8).

Figure 1: Yields and average duration for two countries

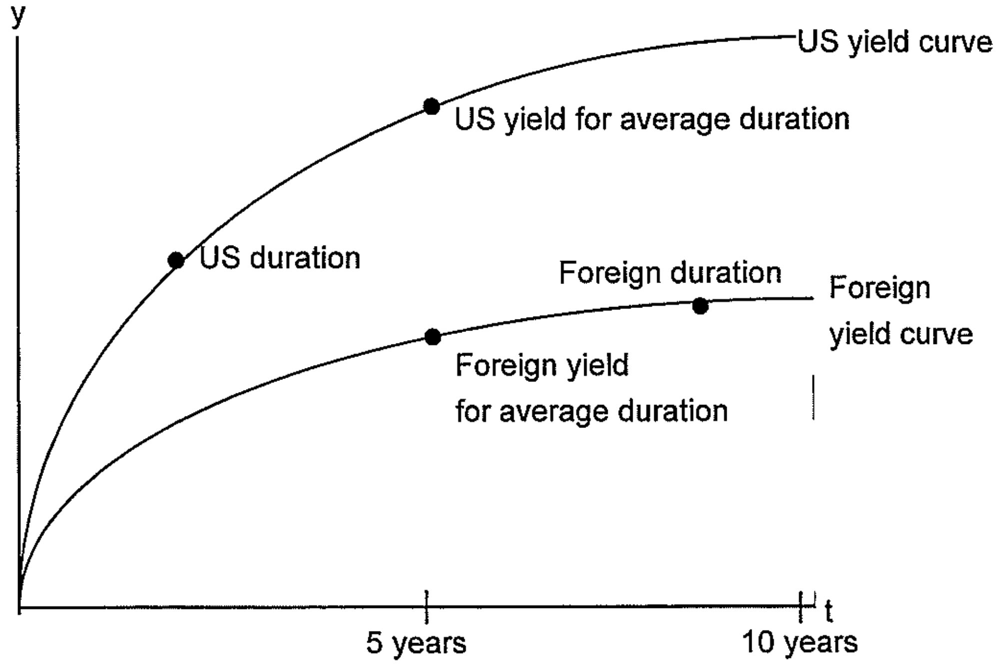


Figure 1 illustrates how the average duration and the two interest rates at this average duration are derived. It shows the constructed US yield and Foreign yield curves as given by (8). The two observed interest rates are yields to maturity on ten year coupon bonds. Depending on how large the coupons are, the durations of the bond are located a certain distance to the left along the yield curves. The change of the exchange rate can only be matched to a single point on the yield curve, taken to be the average of the home and foreign durations. The two interest rates at this average duration D^* are found by moving to the relevant maturity along the constructed yield curves. Holding period returns from t to D^* are given by as $r_{D^*} = (1 + y_{10})^{D^*} - 1$ and exchange rate changes are calculated as $(s_{t+D^*} - s_t)/s_t$, where D^* differs for each observation

Table 3: Tests of UIP using duration data.

Regression: $\frac{s_{t+\tau}-s_t}{s_t}=\alpha+\beta\left(r_{t,t+\tau}-r^*_{t,t+\tau}\right)+\varepsilon_t$.

	α	β	$F(\alpha=0,\beta=1)$	$F(\beta=1)$	R^2
Austria	0.041	**0.852**	**3.079 (0.214)**	**1.925 (0.165)**	0.475
	[0.658]	**[8.014]**			
Belgium	-0.046	0.067	15.458 (0.000)	6.646 (0.009)	0.001
	[-0.394]	[0.184]			
Canada	0.066	0.070	15.248 (0.000)	13.202 (0.000)	0.242
	[1.697]	[0.275]			
Denmark	**-0.210**	**1.126**	10.584 (0.005)	**0.240 (0.624)**	0.478
	[-2.991]	**[4.375]**			
France	0.050	**0.748**	**0.494 (0.781)**	**0.438 (0.508)**	0.069
	[0.334]	**[1.969]**			
Germany	0.105	**0.820**	**3.890 (0.143)**	**0.865 (0.352)**	0.312
	[1.575]	**[4.240]**			
Italy	0.058	**0.998**	**0.120 (0.905)**	**$3*10^{-5}$ (0.996)**	0.181
	[0.447]	**[2.606]**			
Japan	**-0.235**	**0.209**	180.447 (0.000)	80.841 (0.000)	0.183
	[-9.577]	**[2.370]**			
Netherlands	0.142	**1.273**	**3.427 (0.180)**	**0.574 (0.449)**	0.436
	[1.021]	**[3.530]**			
Spain	0.107	**0.336**	15.393 (0.000)	15.379 (0.000)	0.021
	[0.547]	**[1.988]**			
Sweden	0.111	0.157	**5.364 (0.068)**	**5.234 (0.022)**	0.003
	[0.864]	[0.424]			
Switzerland	-0.037	**0.371**	103.181 (0.000)	17.832 (0.000)	0.277
	[-0.448]	**[2.489]**			
UK	0.093	**0.278**	16.884 (0.000)	14.383 (0.000)	0.096
	[1.496]	**[1.859]**			

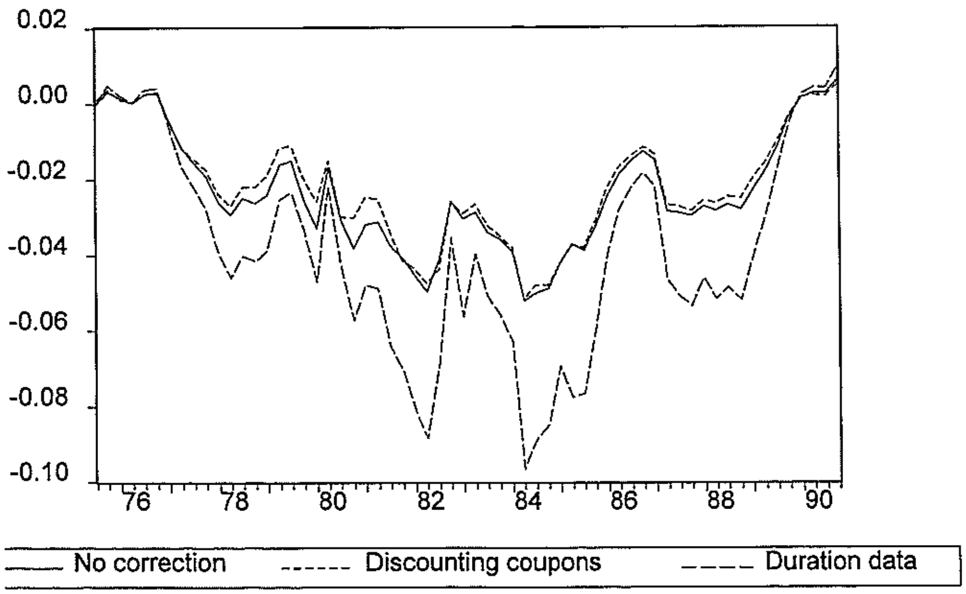
t-values using Newey and West standard errors within brackets. Bold numbers denote significance at the five percent level.

p-values of the *F*-tests within parenthesis. Bold numbers in indicate that the null hypothesis is not rejected at the five percent level.

The results in Table 3 are much more favourable to UIP then the previous ones. All estimates of β are positive and ten of thirteen are significant. They are also larger than previous estimates - six of them are above 0.5 and two are larger than one. The UIP-hypothesis that β equals one is not rejected in more than half of the cases and the joint hypothesis $[\alpha,\beta]$=[0,1] is not rejected in five of fifteen cases. Furthermore, interest differentials now have considerable explanatory power for exchange rate changes as evident from the (centered) R^2 values in the final column.

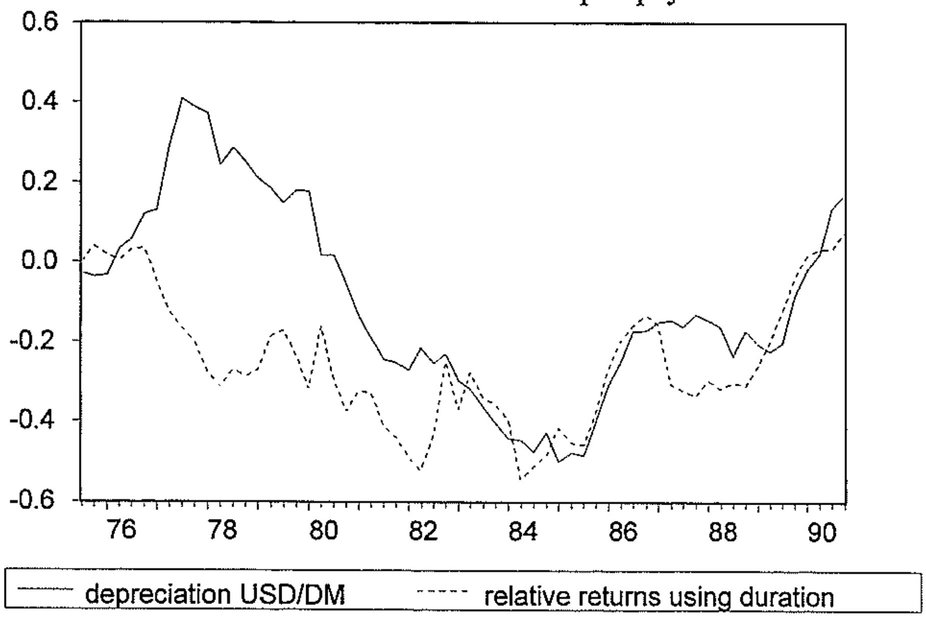
Figure 2 shows the three series on relative annual effective interest rates for Germany versus the US, i.e. the yearly DM return to holding a German bond over the period minus the yearly USD return to holding a US bond. The solid line is simply the difference in yields to maturity as they appear on the IFS data tape. The other two lines show the German effective interest rate minus the US effective interest rate after the effects of coupon payments has been removed as described in Sections 2 and 3. For instance, the dashed "Duration data"-line is calculated as $\left(1+r_{t,t+D^*}^{GE}\right)^{\frac{1}{D^*}} - \left(1+r_{t,t+D^*}^{US}\right)^{\frac{1}{D^*}}$.

Figure 2: German bond return minus US bond return



As shown in Figure 2, the series used in the first two sections do not differ very much, while the duration method has a larger effect on relative returns. The sample period for the two coupon adjusted series is only 1975:3 to 1997:4 due to the short series on German three month interest rates. Since the ex post change of the nominal exchange from t to t+τ is known only up to 1997:4 minus τ, the sample for which UIP may be tested is even shorter (1975:3 to 1990:4 in this case). Figure 3 shows the ex post depreciation of German Marks per US dollar and the relative returns in respective currencies for Germany against the US for the duration data. As evident from Table 3, UIP is not rejected for this data set.

Figure 3: Ex post depreciation of USD/DEM and relative returns using the duration method to remove the effects of coupon payments



The results of Tables 1 to 3 are robust to changes in the maturity of the long term government bond yield. Varying τ between eight and twelve years does not effect the qualitative results. Since the duration method of removing the effects of coupon payments also affects the maturity, the results from reducing the assumed maturity of the raw time series from ten to eight years and matching the exchange rate change could be expected to mimic Table 3 rather than Table 1, especially since the average duration is actually between seven and eight years.

The fact that the coefficients are much more significant when durations are used for the dependent variable than with a similar *average* correction of the maturity supports the interpretation that the duration method of removing coupons reduces the measurement errors. The results from running the regression for the samples 1973:1 to 1997:4 or 1985:1 to 1997:4 are also qualitatively similar, although no formal tests for breaks in the relationhip has been performed.[5]

The results in Table 3 are also robust with respect to the assumption used when constructing common durations for the two countries. Turning back to Figure 2 for a moment, it is clear that calculating the average of the two durations is not the only way to proceed. The shortest or the longest durations could be used, as well as the foreign or the US duration. None of these alternative ways of constructing the data matter for the results (i.e. the same parameters are still significant and the UIP tests yield the same results).

5. Conclusions

Uncovered interest rate parity is normally rejected in empirical tests. Ex post exchange rate changes are even negatively related to interest rate differentials - the regression coefficient is typically minus three to minus four instead of plus one as expected from UIP. However, previous tests have almost exclusively relied on data on short interest rates and short run exchange rate changes. It can be argued that the negative relationship is due to the fact that short interest rates are used in monetary policy to affect the nominal exchange rate. Before establishing that UIP does not hold in empirical tests, the hypothesis should also be tested on long term interest rates.

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[5] Due to the overlapping data on ex post exchange rate changes, there are too few independent observations for formal tests for break points to be meaningful.

In this paper, UIP is tested on a data set of long interest rates and matching exchange rate changes for 14 OECD countries between 1957 and 1997. The interest rates are yields to maturity for long term government bonds with coupon payments. The presence of coupon payments induces a measurement error between yields to maturity and actual returns to investments. This is handled in three different ways. First, the problem is simply neglected and the UIP tests are run using the raw time series on yields to maturity. Here, only five of fourteen regression coefficients are significantly different from zero. However, measurement errors in the dependent variable bias the coefficient toward zero. Attempts to reduce the measurement errors by removing the effect of coupon payments from the data may well yield results that are more favourable to UIP.

Second, the discount bond price is approximated by removing the present values of the coupon payments from the coupon bond price. The results from testing UIP on this data set are slightly more favourable to UIP in that more regression coefficients are significant. They are still smaller than one, implying that exchange rates do not move enough to offset international differences in interest rates.

In the final set of UIP tests, the coupon payments are allowed to affect the length of the investment period rather than the bond price. The duration is calculated for each observation and the matching exchange rate change is found. Now, UIP cannot be rejected in more than half of the cases. All estimated betas are positive and ten out of thirteen are significant. The parameter estimates are also much larger than in the previous tests. Since this method of removing the effects of coupon payments is clearly preferable to the other two (the measurement errors are smaller since the assumptions about the unobservable variables relating to coupon payments are less crucial), it may be motivated to give more weight to these results than to the previous ones.

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