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University of Groningen, The Netherlands

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Does Uncovered Interest Rate Parity Hold After All?[1]

Muhammad Omer[a], Jakob de Haan[a,b,c] and Bert Scholtens[a]

[a] University of Groningen, The Netherlands
[b] De Nederlandsche Bank, Amsterdam, The Netherlands
[c] CESifo, Munich, Germany

Abstract:

This paper tests Uncovered Interest Rate Parity (UIP) using LIBOR rates for the major international currencies for the period January 2001 to December 2008. We find that UIP generally holds over a short-term horizon for individual and groups of currencies. Our results suggest that it is important to take the cross correlation between currencies into account. We also find that 'state dependence' plays an important role for currencies with a negative interest differential vis-à-vis the US. This 'state dependence' could also be instrumental in explaining exchange rate overshooting.

JEL Classification: G12; G15; F31

Key words: UIP, LIBOR, system SUR, System DGLS, System DOLS

Corresponding author: Muhammad Omer, Faculty of Economics and Business, University of Groningen, PO Box 800, 9700 AV Groningen, The Netherlands. Tel. +31-50-336-8961; Fax +31-50-336-7356; email m.omer@rug.nl.

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1. Introduction

Uncovered interest rate parity (*henceforth* UIP) suggests that any arbitrage opportunity between interest-earning assets, of different economies but with similar characteristics, will disappear due to exchange rate movements. A positive shock to the domestic interest rate *vis-à-vis* the foreign interest rate will lead to the depreciation of the home currency and *vice versa*. UIP plays a critical role in most exchange rate determination theories, such as the monetary exchange rate model, Dornbusch's (1976) overshooting model and Krugman's (1991) target zone model. Also central banks frequently count on this relationship for anchoring exchange rate expectations in the economy (Kalyvitis and Skotida, 2010).

It is surprising that theorists continue to rely on UIP despite ambiguous empirical support. Several studies (Bekaert and Hodrick (1993), Engel (1996), Froot and Thaler (1990), Mark and Wu (1998), Weber (2011) and Tang (2011) to mention just a few) reject UIP. Only few studies report (some) support for UIP, including Flood and Rose (1996), Bekaert and Hodrick (2001), Baillie and Bollerslev (2000), Chaboud and Wright (2005) and Beyaert, García-Solanes and Pérez-Castejón (2007).

Given the crucial role played by UIP in exchange rate theories and exchange rate stabilization policies, this relationship warrants more detailed investigation. Unambiguous evidence supporting UIP will not only increase the confidence in the existing exchange rate models but may also enhance the quality of monetary policy decision-making. This research is an effort in this direction.

This paper extends the existing UIP literature by zooming in on important issues affecting this relationship. First, we use a multi-currency setup to exploit cross currency correlation. Some previous studies using Seemingly Unrelated Regression Equations (SURE), such as Flood and Rose (1996) and Mark and Wu (1998), have exploited cross currency correlations. However, most studies investigate UIP mostly bilaterally. In our view, bilateral studies implicitly impose restrictions on the third-country effect, which may play an important role in determining exchange rates. This is equally true for studies using a panel setup that ignores cross sectional dependence. In a globalized world, any shock to the US debt market say, will not only affect the Japanese debt market but also the Euro debt market. Therefore, an interest rate shock in the US will not only affect the US Dollar and the Japanese Yen exchange rate or the US Dollar and the Euro exchange rate, but also the Euro-Yen exchange rate. Studies on UIP have mostly ignored this cross currency correlation.

Second, we use data for industrial economies as the literature suggest that for these countries the problem of a forward premium puzzle is more prominent (see Alper, Ardic, and

Fendoglu (2009), Bansal (1997), Bansal and Dahlquist (2000)). For developing and emerging market economies, the empirical evidence provides more support for UIP (see, for example, Frankel and Poonawala (2006), Ferreira and Leon-Ledesma (2007), Flood and Rose (2002) and, Bansal and Dahlquist (2000)).

Third, instead of using domestic interest rates we use London Interbank Offered Rates (LIBOR). LIBOR is an indicative interbank rate for specific currencies based on the non-binding quotes in the London interbank market.[2] LIBOR rates are widely used as benchmarks in global financial transactions.[3] The statistical evaluation supports LIBOR as a substitute for domestic interest rates. Factor analysis shows that the LIBOR rates are defined by only one factor, the domestic interest rates.[4] Using LIBOR has several advantages. For instance, the currency specific LIBOR rates have similar transaction costs for the assets denominated in various currencies, while capital is perfectly mobile. Juselius and MacDonald (2004), Harvey (2005) and Ichiue and Koyama (2007) have used LIBOR as a proxy for Japanese domestic rates, arguing that the thin and heavily regulated Japanese money market in the eighties and nineties was less reflective of Japan's economic fundamentals.

Finally, following a suggestion of Moon and Perron (2005), we take as our null hypothesis that UIP holds (the slope coefficient is unity). Often the null hypothesis tested is that the slope coefficient is not different from zero, which on rejection provides support for the alternative hypothesis that the slope coefficient is different from zero. According to Moon and Perron (2005), such a test design has a strong bias towards the null hypothesis, which is rejected only when there is a strong support against it. Moreover, when the null of a zero slope coefficient cannot be rejected, it is difficult to conclude whether the theory is rejected or the power of the test is low.

Our estimates using weekly data for the period January 2001 to December 2008 support UIP over the short-term horizon for currencies from advanced countries. Moreover, our currency specific estimates show that the null hypothesis of a unit coefficient can generally not be rejected at the 5% level of significance. However, for the Japanese Yen and the Swiss Franc, the slope coefficients are negative. This finding is consistent with the argument put forward by Bansal and Dahlquist (2000) and Ballie and Kalic (2006) that deviations from UIP appear only when the US interest rate exceeds the foreign interest rate

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[2] For details see Michaud and Upper (2008).

[3] Forbes Investopedia estimates that $360 trillion worth of international financial products are benchmarked with LIBOR. Additionally, one trillion dollars of sub-prime mortgages have rates adjustable to LIBOR.

[4] Factor analysis is widely used technique for summarizing a usually large number of variables with small number of factors. For the sake of brevity, we do not report the result of the factor analysis but they are available on request.

('state dependence'). Once we incorporate the negative interest rate differential, UIP cannot be rejected for the Japanese Yen and the Swiss Franc. Our results show that cross currency effects play an important role in determining the exchange rate between currencies. Finally, we also find some support for Dornbusch's (1976) overshooting hypothesis for exchange rates, specifically for the Japanese Yen and the Swiss Franc against the US Dollar, suggesting that 'state dependence' could also be instrumental in explaining exchange rate overshooting.

The rest of the paper is structured in the following way. Section 2 reviews the literature. Section 3 delves into data and methodology issues, while section 4 presents our results. Finally, section 5 offers our conclusions.

2. Literature review

According to the Covered Interest Rate Parity (CIP) hypothesis, under risk free arbitrage the ratio of the forward to the spot exchange rate will be equal to the ratio of the returns on two similar assets, measured in the local currencies. Using logarithmic transformation, CIP can be written as:

$$(f_{t,t+i} - s_t) = (r_{i,t} - r^*_{i,t}) \tag{1}$$

where $f_{t,t+i}$ is the forward rate for maturity *i*, s_t is the spot exchange rate, $r_{i,t}$ and $r^*{}_{i,t}$ are the gross return at any time *t* for maturity *i* on a domestic and foreign asset, respectively. However, if forward rates deviate from the expected future spot rate, a risk premium is required such that

$$[E(s_{t,t+i}) - s_t] = \alpha + (r_{i,t} - r^*_{i,t}) \tag{2}$$

where α is the risk premium and $E(s_{t,t+i})$ is the expected future exchange rate at time *t+i*. Under UIP, the risk premium is zero and the coefficient of the interest differential is one. Since the future spot exchange rates cannot be observed directly, UIP is generally tested jointly with the assumption of rational expectations in the exchange rate market (Chinn, 2007).

$$[RE(s_{t+i}) - s_t] = \alpha + \beta(r_{i,t} - r^*_{i,t}) + \varepsilon_{t+i} \tag{3}$$

Following studies such as Tang (2011), Bekaert, Wei and Xing (2007), Chinn and Meredith (2004), and Carvalho, Sachsida, Loureiro, and Moreira (2004), we assume that agents have perfect foresight so that exchange rate movements can be estimated using equation (4)

$$[s_{t+i} - s_t] = \alpha + \beta(r_{i,t} - r^*_{i,t}) + \varepsilon_{t+i} \tag{4}$$

To simplify the notations, the exchange and interest rate differentials are denoted by

y_{it} and x_{it}, respectively. Equation (4) then reduces to,

$$y_{it} = \alpha + \beta x_{it} + \varepsilon_{t+k} \tag{5}$$

Most studies on UIP report a negative point estimate for beta over the short-term horizon (see Froot and Thaler (1990), MacDonald and Taylor (1992), Isard (1996), McCallum (1994), and Engel (1996) and Chin and Meredith (2004)). A notable exception is Flood and Rose (1996), who report a slope coefficient close to one during the period with exchange rate alignments within Europe's Exchange Rate Mechanism (ERM). Other studies, like Bruggemann and Lutkepohl (2005), Huisman, Koedijk, Kool and Nissen (1998), and KrishnaKumar and Neto (2008) provide indirect support for UIP. More precisely, Huisman, Koedijk, Kool and Nissen (1998) have shown that the large forward premium provides an unbiased estimate of the future change in the spot rate while small forward premium fails to predict the same correctly. Bruggemann and Lutkepohl (2005), and KrishnaKumar and Neto (2008) have tested UIP jointly with the expectation hypothesis of the term structure (EHT) using interest rates of the respective economies. By assuming that exchange rates are generated by a stationary process they provided evidence in support of UIP using the stationarity of the interest rate differential.

Bansal (1997) reports that the failure of UIP is more severe for industrial economies compared to developing economies. In addition, Bansal and Dahlquist (2000) and Ballie and Kalic (2006) point to 'state dependence' in the UIP relationship, i.e. the exchange rate denominated in the US Dollar responds differently to the positive or negative interest rate differentials. More specifically, deviations from UIP appear only when the US interest rate exceeds the foreign interest rate. When the foreign interest rate exceeds the US interest rate, the expected depreciation and the interest rate differentials are positively related.

Several studies have tested UIP bilaterally, thereby implicitly imposing restrictions on the third economy's effect. Moreover, this restriction might have fostered non-linearities in the UIP relationship, a subject investigated by a different string of literature.[5] Studies using panel techniques and ignoring the cross currency effect, suffer from similar problems.

Chinn and Meredith (2004) note that UIP models by construction have cross-equation correlation of the error terms and therefore techniques incorporating cross-currency correlations, such as seemingly unrelated regression estimation (SURE), are appropriate. Two studies, Flood and Rose (1996) and Mark and Wu (1998), have employed SURE to control for cross currency correlations. However, the outcomes of both studies are very different.

[5] Studies discussing non-linearities in UIP include Baldwin (1990), Dumas (1992), Sercu and Wu (2000), Lyons (2001), Kilian and Taylor (2003), and Carlson and Osler (1999).

While Flood and Rose (1996) report a slope coefficient close to one during the period with exchange rate alignments within Europe's Exchange Rate Mechanism (ERM), Mark and Wu (1998) do not find enough support for UIP. To control for the cross equation correlation, both studies employ SURE based on OLS but using the contemporaneous covariance matrix. A contemporaneous covariance matrix uses current information only, ignoring long-run relationships which may be misleading if there exists such a long-run relationship.

When regressors are integrated, indicating a long-run relationship between them, Moon and Perron (2005) have shown that the limiting distributions of OLS estimators are not normal. To solve this problem, they propose augmenting the regressors with their leads and lags to capture the long-run correlation. In addition, they argue for using the long-run covariance matrix instead of the contemporaneous covariance matrix, which enhances the efficiency gain of the long-run estimators. This study therefore uses SURE with integrated regressors as proposed by Moon and Perron (2005).

3. Data and methodology

3.1 Data

Our sample period is January 2001 - December 2008. We use the following currencies: the Euro, the Japanese Yen, the British Pound, the Australian Dollar, the Canadian Dollar, and the Swiss Franc against the US Dollar. We have acquired daily data on the exchange rates from the International Monetary Fund (IMF).[6] For the interest rates, we use daily LIBOR rates for the above currencies, with short maturities. The LIBOR interest rates data can be accessed from the British Bankers Association (BBA) Website.[7] Exchange rate differentials are calculated assuming that economic agents have perfect foresight. So the 6-month exchange rate differential series, for example, is calculated by subtracting current spot rates from spot rate after six months. Similarly, interest rate differentials are generated by subtracting currency- and maturity-specific LIBOR from US Dollar LIBOR with similar maturity. We use maturities ranging from 6 to 12 months. Consistency of the estimates over the consecutive maturities enhances confidence in our results.

From daily data we have calculated weekly and monthly.[8] Figure 1 shows the 6-months interest rate differential series for all currencies. Other maturities show more or less similar variation. Figure 1 shows that these series follow similar pattern, and hence are highly

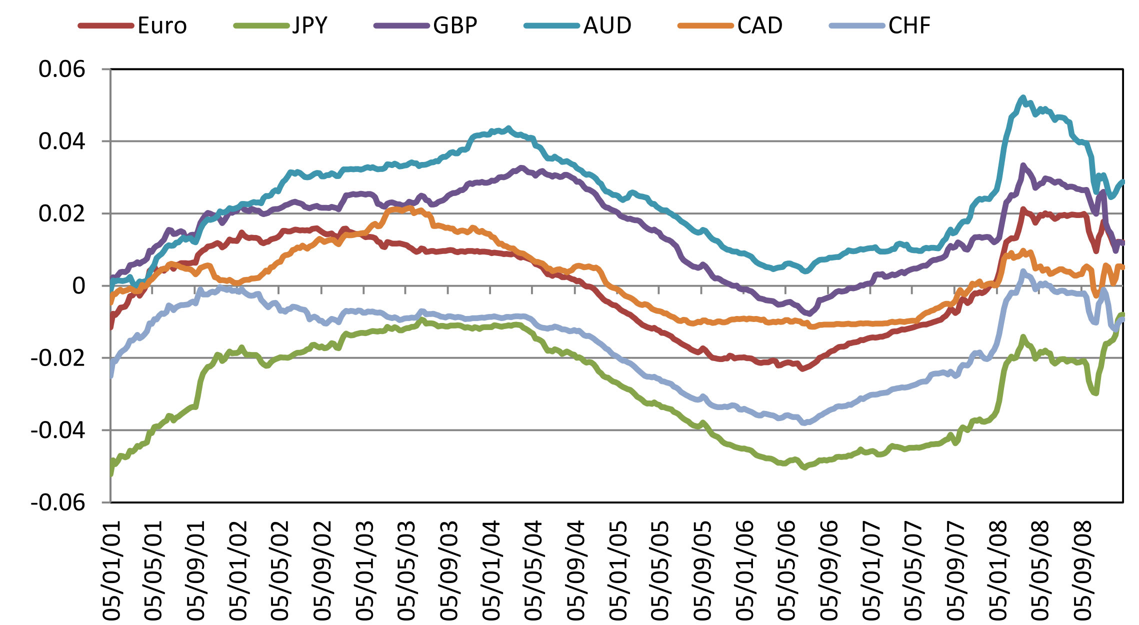
[6] http://www.imf.org/external/np/fin/data/param_rms_mth.aspx.
[7] http://www.bbalibor.com/rates/historical.
[8] Weekly averages are calculated using five working days.

positively correlated (see Panel A of Table A1 in the appendix). Importantly, both the Japanese Yen and the Swiss Franc have negative interest rate differentials since the US Dollar LIBOR rates are higher than these currency specific rates.

3.2 Methodology

Since our dataset involves long time series, it is essential to ascertain the nature of the data-generating process of the regressors. Therefore we have applied unit root tests.

Figure1: Movement in 6m- interest rate differential



Previous studies generally adopted ‘time series’ unit root tests, such as the Augmented Dickey Fuller (ADF) or Phillip and Perron (PP) tests. As these ‘time series’ tests impose restrictions on cross correlation effects. Therefore, we apply the Cross-sectional Dependence Robust Block Bootstrap (CDRBB) panel unit roots test proposed by Palm, Smeekes, and Urbain (2010).

The CDRBB unit root test does not require modeling the temporal or cross sectional correlation (dependence) structure between the currency-specific interest rates. Moreover, it uses block bootstrap techniques, a time series version of a standard bootstrap where the dependence structure of the time series is preserved by dividing data into blocks and then re-sampling the blocks. However, the block length selected can have a large effect on the performance of any designed block bootstrap test. Inferences from the CDRBB test are valid under a wide range of possible data-generating processes, which makes it an appropriate tool for dealing with the fixed number of correlated cross-sections and large time series asymptotics.

Although this CDRBB test provides both 'pooled' (τ_p) and 'group-mean' (τ_{gm}) test statistics, we are interested in the 'group mean' statistic only (see equation 6). This statistic incorporates the member-specific information, which is more relevant for the one-on-one currency based analysis of UIP. The null hypothesis assumes that the series is non-stationary while under the alternative hypothesis a portion of the series is stationary. Rejection of the null hypothesis for the first difference of a series and non-rejection for the level of the same series indicates that the series concerned has a unit root.

$$\tau_{gm} = \frac{1}{N}\sum_{i=1}^{N} T \frac{\sum_{t=2}^{T} y_{i,t-1}\Delta y_{i,t}}{\sum_{t=2}^{T} y_{i,t-1}^{2}} \tag{6}$$

In equation (6), y_t is the series tested for unit roots, N is the number of currencies and T is the sample period.

Next, we apply individual time series specific Johansen's cointegration test as well as Westerlund's (2007) ECM based panel cointegration test. The former, being the 'individual time series' test, imposes restrictions on cross correlation effects, while the latter takes those effects into account. For brevity, we will only report the results of Westerlund's (2007) ECM based cointegration tests.

$$\Delta y_{it} = \delta_i d_t + \alpha_i (y_{i,t-1} - \beta_i' x_{i,t-1}) + \sum_{j=1}^{p_i} \gamma_{1ij} \Delta y_{i,t-j} + \sum_{j=0}^{p_i} \gamma_{2ij} \Delta x_{i,t-j} + u_{it} \tag{7}$$

Westerlund (2007) suggests a panel cointegration test based on the error correction mechanism (ECM) as indicated by equation (7). Here, d_i is the currency specific deterministic component, δ_i is the associated parameter, α_i is the speed of adjustment for the error correction term, β_i is the cointegrating vector while x_{it} and y_{it} are interest and exchange rate differentials series, respectively. The choice of the appropriate number of leads and lags, given by p_i, using information selection criteria, such as Akaike's Information Criterion (AIC), transforms u_{it} into white noise.

The null hypothesis of the cointegration test is $\alpha_i = 0$, which indicates no cointegration of the variables. For the alternative hypothesis Westerlund (2007) provides four test statistics, two for the pooled test and two for the group mean test. Pooled test statistics assumes common slope parameter across the cross sections, while the group mean test statistics aggregates the estimated individual slope coefficients. We use the group mean test statistics (G_α and G_τ) only. These statistics differ in composition. Whereas G_α is calculated by aggregating the individual slope coefficients with the help of conventional standard errors, G_τ is designed by aggregating the individual slope coefficients using Newey and West (1994) long-run standard errors. The alternative hypothesis for the group mean test is that at least one member of the panel is cointegrated. Simulation results of Westerlund (2007) show that G_α has a higher power compared to G_τ in samples where T is substantially larger than N. Asymptotically, both statistics have a limiting normal distribution and they are consistent. Moreover, Westerlund's (2007) procedure provides robust critical values for the test statistics by applying bootstrapping which accounts for the cross sectional dependence.

For drawing inference on long-run relationships, we use Moon and Perron's (2005) efficient estimation method of a system of Seemingly Unrelated Regression (SURE) equations with integrated regressors. This method provides more efficient estimates by exploiting the correlations among the multiple currencies while allowing for individual currency-specific inferences. Conventional system estimation methods, such as GLS, with integrated regressors have a nonstandard limiting distribution that is skewed and shifts away from the true parameters. This renders inference difficult. Moon and Perron (2005) suggest a method for obtaining efficient estimators with a mixed normal limiting distribution. By adding the leads and lags of the first differences of the regressors, they suggest applying GLS on this augmented dynamic regression model using information on the long-run covariance matrix, hence its name: System Dynamic GLS (SDGLS).

The Monte Carlo simulation results of Moon and Perron (2005) show that SDGLS performs better compared to other estimators.[9] Moreover, the efficiency gain of the SDGLS

[9] Using their proposed method based on SURE technique Moon and Perron (2005) have suggested number of estimators such as system dynamic OLS (SDOLS) or fully modified OLS (FMOLS), besides dynamic GLS estimator. The system dynamic OLS (SDOLS) given by (8A).

$$\hat{b}_{SDOLS} = \left(\sum_{t=k+1}^{T-k} z_t z_t^{'}\right)^{-1} \left(\sum_{t=t+k}^{T-k} z_t y_t\right) = b + \left(\sum_{t=k+1}^{T-k} z_t z_t^{'}\right)^{-1} \left(\sum_{t=t+k}^{T-k} z_t \xi_t^{*}\right) \tag{8A}$$

estimates is greater compared to other estimates obtained in similar fashion. Furthermore, the SDGLS estimator suffers least from distortion due to a small sample. Based on its superior performance, we utilize the SDGLS estimator.

$$\hat{b}_{DGLS} = \left(\sum_{t=k+1}^{T-k} z_t \hat{\Omega}_{uu.v}^{-1} z_t^{'} \right)^{-1} \left(\sum_{t=t+k}^{T-k} z_t \hat{\Omega}_{uu.v}^{-1} y_t \right) = b + \left(\sum_{t=k+1}^{T-k} z_t \hat{\Omega}_{uu.v}^{-1} z_t^{'} \right)^{-1} \left(\sum_{t=t+k}^{T-k} z_t \hat{\Omega}_{uu.v}^{-1} \xi_t^{*} \right) \quad (8)$$

Equation (8) shows the SDGLS estimator using the multivariate format of SURE. Here, *b* is matrix of coefficients of regressors and the leads and lags of the first difference of the regressors, $z_t = (\tilde{x}_t^{'}, \Delta x_{t-k}^{'} \otimes I_N, ..., \Delta x_{t+k}^{'} \otimes I_N)^{'}$, $\tilde{x}_t^{'} = diag(\tilde{x}_{1t}, \ldots \tilde{x}_{Nt})$, $\tilde{x}_{it} = (1, x_{it}^{'})$, $x_{it} = (x_{1t}^{'}, \ldots, x_{Nt}^{'})^{'}$, ξ_t^{*} is error term with non-estimable part of regressors beyond *k*. The null hypothesis tests whether the individual slope coefficient (*b*) is unity, or in other words whether UIP holds on a currency-specific basis.

This direct test of UIP differs from the usual testing methodology in which the null hypothesis is that the coefficient is not different from zero. According to Moon and Perron (2005), such test design has a strong bias towards the null hypothesis which also affects the interpretation of the test results in an undesirable way. When the null hypothesis cannot be rejected in such tests, it is hard to determine whether the theory is rejected or the power of the test is low. Another advantage of the Moon and Perron test design is that it does not require testing cointegration separately. If the error term is non-stationary for any of the model coefficients, the test statistics diverge to infinity thereby rejecting the null hypothesis that UIP holds. This alternative test for cointegration based on coefficient of the cointegrating vector is more powerful than simple cointegration tests (Cheung and Lai, 1993).

4. Results

Table 1 reports the group mean CDRBB panel unit root tests. For both the interest and the exchange rate differential series, at all maturities, the null hypothesis of a unit root cannot be rejected at the 5% level of significance indicating that the level of these series are non-stationary. A test on the first differences of these series confirms that all maturities are following an I(1) process (not reported for brevity).

Notations have the same meaning as in equation (8). Both estimator $\hat{b}_{SDOLS}$ and $\hat{b}_{DGLS}$ uses the long-run correlation information of the system.

Table 1. Block bootstrap panel unit root tests

	Exchange Rate Differential Series			Interest Rate Differential Series		
	Coeff.	**5% CV**	**P-value**	**Coeff.**	**5% CV**	**P-value**
6-month	-6.9340	-11.4470	0.3370	-2.1540	-6.7890	0.6620
7-month	-6.5470	-12.1160	0.4220	-2.1300	-6.8150	0.6820
8-month	-5.7180	-12.1480	0.5700	-2.0930	-6.9130	0.7060
9-month	-5.9400	-12.4110	0.6010	-2.0780	-7.0330	0.7270
10-month	-5.7460	-13.2370	0.6740	-2.0730	-7.1480	0.7420
11-month	-5.5980	-13.2420	0.6330	-2.1030	-7.3280	0.7520
12-month	-5.8910	-13.0600	0.5870	-2.1220	-7.5120	0.7690

Estimated test statistics for equation (6) at level of exchange rate and interest rate differential series. 5% CV indicates robust critical values calculated at 5% level of significance. P-values indicate the corresponding probability values of the calculated test statistics.

Next, we apply the Johansen (1995) cointegration tests on individual currency-specific time series. The results do not provide any evidence for a cointegration relationship between interest and exchange rate series (results available on request). In contrast, the Westerlund (2007) ECM based panel cointegration tests as shown in Table 2 indicate that the null hypothesis of no cointegration is rejected for maturities ranging between 6 and 9 months at the 5% level of significance. The results indicate that at least one member of the panel is cointegrated for these maturities. For the other maturities, the evidence for 'no cointegration' is rather weak as the rejection probabilities (p-values) are very low. So our results suggest that inferences regarding financial market variables based on the Johansen cointegration test can be misleading if cross correlation effects are ignored.

Table 2. Results for the Westerlund cointegration test

	G_α			G_τ		
	Value	**Z-value**	**Rob. P-value**	**Value**	**Z-value**	**Rob. P-value**
6-month	-12.2080	-4.5270	0.0000	-2.2560	-3.0120	0.0020
7-month	-9.8120	-3.2370	0.0000	-1.9260	-2.2370	0.0200
8-month	-8.3590	-2.4540	0.0200	-1.7430	-1.8050	0.0540
9-month	-7.6540	-2.0740	0.0360	-1.7510	-1.8240	0.0640
10-month	-6.6340	-1.5250	0.0560	-1.6230	-1.5220	0.0620
11-month	-5.4670	-0.8960	0.1240	-1.4310	-1.0710	0.1440
12-month	-5.3260	-0.8210	0.1220	-1.4430	-1.0990	0.1240

Estimates of ECM coefficient based on equation (7). The alternative hypothesis of these test statistics are the cointegration relationship exists when the panel taken as whole. 5 and 21 are the maximum number of leads and lags considered for estimation. Values give the estimated values of the coefficients and Z-values are their standardized values. Rob. P-values are the robust probability values calculated using the bootstrap technique. The corresponding values show the level of significance.

As pointed out, the methodology we have adopted here for making inference does not require testing cointegration separately. Therefore, our cointegration results as reported in

Table 2 should be considered as a robustness check of the system SURE estimates to which we turn now.

We have applied SURE on interest and exchange rate differential series for each maturity separately using a maximum of 12 leads or lags. Table 3 shows the estimation results using system DGLS, which includes the individual slope coefficient for each currency *vis-à-vis* the US Dollar.

The Wald test aggregates the individual currency specific slope coefficient and tests the null hypothesis whether the joint slope coefficient is unity. In other words, it tests whether UIP holds for the system of currencies taken together. The reported p-values for Wald test statistics shows that the null hypothesis cannot be rejected for maturities ranging between 10 and 12-months. Hence, UIP holds for these maturities when all six currencies are taken together.[10]

Table 3. Estimation results using System DGLS (SDGLS)

	6-m	7-m	8-m	9-m	10-m	11-m	12-m
Euro	3.0261	2.3765	3.8135	5.1693**	2.9493	2.9231	3.2848
	1.7716	*1.5452*	*1.8149*	*2.2376*	*2.3597*	*2.8520*	*3.6336*
JPY	-1.2921**	-1.1077**	-1.296**	-1.5944*	-1.0286**	-1.6118**	-1.3214
	1.2585	*1.2611*	*1.1870*	*1.0759*	*1.0482*	*1.0551*	*1.5769*
GBP	2.1321	0.4204	0.4771	-0.1292	-0.4108	0.1099	3.1353
	1.6566	*1.3567*	*1.5417*	*1.9650*	*2.3640*	*2.0442*	*2.1757*
AUD	0.5314	-0.4379	0.6683	-0.1183	1.3554	1.9217	1.0261
	1.7308	*1.6050*	*1.9285*	*2.3469*	*3.1521*	*2.2673*	*2.5794*
CAD	-0.1784	1.1095	0.0519	-1.1833	-0.4472	0.0817	-1.4897
	1.6382	*1.9127*	*1.7832*	*1.7276*	*1.6198*	*1.9262*	*2.8642*
CHF	-5.6004*	-3.3885*	-1.8798**	-1.6504**	-1.3616**	-1.1008	-1.5929
	2.4140	*1.7692*	*1.5111*	*1.5988*	*1.3264*	*1.7211*	*1.7633*
Wald Stats	17.3979	16.0100	12.6966	19.2999	9.3929	7.5384	7.8070
Wald p	0.0079	0.0137	0.0481	0.0037	0.1527	0.2739	0.2526

Estimates of System DGLS coefficient based on equation (8) using average weekly data with maximum leads and lags of 12 weeks. The optimal lag length selected using Bayesian Information Criteria (BIC). The null hypothesis is individual coefficient is unity. The figure in the italics shows the standard errors. The null hypothesis for the Wald test is the joint beta coefficient is unity. Wald P shows the P-values of the Wald test statistics. The symbols indicates *, < 5 % and ** < 10 % level of significance, respectively.

For the individual currency-specific results the conclusion is similar. The null hypothesis of unit slope coefficients cannot be rejected for almost all maturities at the 5% level of significance. Only for the 9-months Japanese yen and the 6- and 7-months Swiss franc the null is rejected. The slope coefficient of the Japanese yen and the Swiss franc are persistently negative. However, as pointed out in section 3.1, both currencies have negative

[10] Estimates from monthly data, reported in Table A2, also fails to reject null hypothesis of the Wald tests for all maturities.

interest rate differential *vis-à-vis* US interest rate. Ballie and Kalic (2006), Bansal and Dahlquist (2000) and Bansal (1997) provide evidence that the exchange rate vis-à-vis the US Dollar responds differently to positive and negative interest rates differentials. Specifically, Bansal and Dahlquist (2000) argue that the forward premium puzzle is present only when the US interest rate exceeds the foreign interest rate.

Interestingly, for the negative interest rate differential series, any increase in the domestic (Japanese/Swiss) interest rates *vis-à-vis* US interest rate means a decrease in the 'differential'. Some studies have used the US Dollar as domestic currency, instead of the foreign currency, to avoid the negative interest rate differential. In a bilateral environment, the flipping of the exchange rate may work, but it is less likely to work in our multi-currency setup. Panel B of Table A1 shows that the correlation structure between the interest rate differential of various currencies when the Japanese Yen and Swiss franc are taken as numeraire currencies against the US Dollar. This flipping of currencies solves the problem of the negative interest rate differential since the US Dollar become home currency. However, the correlation structure between the interest rate differential of the various currencies gets significantly distorted. Our estimation with this modified Japanese Yen and Swiss Franc interest rate setup gives a similar distorted picture of the slope coefficients (results are available on request).

Interestingly, whenever the null hypothesis is rejected in our setup it implies overshooting/undershooting of exchange rates, consistent with Dornbusch's (1976) exchange rate overshooting hypothesis. According to Frenkel and Rodriquez (1982), the exchange rate overshoots when capital is highly mobile while it undershoots when capital is highly immobile. With LIBOR market rates, our setup replicates perfect capital mobility environment. If we restrict the confidence level of our interval estimation to 90%, we find some evidence of persistent overshooting in line with the view of Frenkel and Rodriquez (1982). For both the Japanese Yen and the Swiss Franc the null hypothesis of a unit slope coefficient is rejected at the 10% level of significance. However, we find little evidence of overshooting for the other currencies which leads us to suspect that overshooting could be a state dependent phenomenon as well. In other words, when currencies have low interest rates compared to US interest rates, overshooting of the exchange rate become a possibility. However, more research is needed to draw strong conclusions.

As a robustness check, Table A3 provides the results for the SDOLS estimator.[11] This

[11] For the SDOLS estimator: see footnote 9.

estimator is the most efficient next to the DGLS estimator and suffers least from size distortion compared to fully modified estimators. It turns out that the SDOLS estimates are very similar to those of reported in Table 3.

Finally, a caveat that has to be made is the high variance of the individual slope coefficients. Fully modified estimators, such as FM-GLS, show relatively low estimated variances (results are shown in Table A4) but these estimators are less efficient compared to the system DGLS or DOLS estimators. Moreover, the simulation results of Moon and Perron (2005) show that these fully modified estimators suffer more from size distortion than DGLS or DOLS estimators.

5. Conclusions

In this study we have tested UIP over short-term horizons using the major international currencies. By taking into account the cross correlation, we find strong support for UIP. It turns out that cross currency effects play an important role in the determination of the exchange rate. We also find that state dependence in the interest rate differential series significantly affects the point estimate of the slope coefficients. Once the negativity of the interest rate differential is accounted for, UIP is validated. Changing the numeraire does not help in fixing the state dependence problem in a multi-currency environment. Finally, we find some support for exchange rate overshooting notably for currencies with a negative interest differential vis-à-vis the US Dollar suggesting that this phenomenon may also be 'state dependent'.

We have applied several robustness checks. First, we use maturities ranging between 6 and 12 months. Second, we provide cointegration tests separately, which is a necessary but not sufficient condition, for UIP. Our cointegration results support the presence of a long- run relationship between interest rates and the exchange rate. Third, we provide additional estimates using SDGLS (based on monthly data) and SDOLS (based on weekly data). All robustness checks suggest that our conclusions hold.

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Appendix.

Table A1. Correlation between the interest rate differential series

	Euro	JPY	GBP	AUD	CAD	CHF
Panel A: Full Sample differential vis-à-vis US interest rate						
Euro	1.00	0.85	0.85	0.77	0.84	0.97
JPY	0.85	1.00	0.92	0.88	0.89	0.82
GBP	0.85	0.92	1.00	0.91	0.81	0.86
AUD	0.77	0.88	0.91	1.00	0.74	0.72
CAD	0.84	0.89	0.81	0.74	1.00	0.80
CHF	0.97	0.82	0.86	0.72	0.80	1.00
Panel B: Full Sample Japanese and Swiss interest rates differential vis-à-vis US interest rate						
Euro	1.00	-0.85	0.85	0.77	0.84	-0.97
JPY	-0.85	1.00	-0.92	-0.88	-0.88	0.82
GBP	0.85	-0.92	1.00	0.91	0.81	-0.85
AUD	0.77	-0.88	0.91	1.00	0.74	-0.72
CAD	0.84	-0.88	0.81	0.74	1.00	-0.80
CHF	-0.97	0.82	-0.85	-0.72	-0.80	1.00

This table shows the correlation structure between the currency specific 6-months interest rate differential series. In Panel A, 6-months interest rate differential series are calculated by subtracting the US Dollar interest rate from other currency interest rate. In Panel B, similar procedure applied for all currencies specific interest rates except for the Japanese Yen and the Swiss Franc. For these two interest rates, the home currency interest rate is subtracted from the US Dollar interest rate.

Table A2. Estimation Results for System DGLS (monthly data)

	6-m	**7-m**	**8-m**	**9-m**	**10-m**	**11-m**	**12-m**
Euro	-0.2891	0.1103	2.8782	0.8791	-0.7631	17.0376*	21.8353*
	2.3187	*3.0917*	*2.8813*	*2.9739*	*6.5479*	*4.3569*	*4.8998*
JPY	-3.4596*	-0.2399	1.1653	1.8085	-2.8062**	-1.4989	0.8416
	2.0816	*2.2002*	*1.6906*	*1.6871*	*2.1245*	*2.0517*	*2.5742*
GBP	-0.2934	-0.1913	1.9738	-1.1612	-6.5941*	-3.1139**	1.5832
	1.6146	*1.8934*	*1.568*	*1.7073*	*3.1399*	*2.3495*	*2.7434*
AUD	-0.7196	-0.1402	0.9967	2.8111	0.8597	2.5259	1.5688
	1.8027	*1.4448*	*1.5513*	*1.4736*	*1.9878*	*1.7369*	*2.3275*
CAD	1.5326	-0.0836	-4.0469*	-1.9124**	5.3586**	0.7959	-2.4556
	1.8467	*2.066*	*2.0445*	*1.5508*	*2.3028*	*1.7798*	*2.3412*
CHF	-3.3382	-4.6011	-8.2813*	-5.3422*	-6.0788*	0.6033	6.1701*
	3.8133	*4.3311*	*2.8784*	*2.2252*	*2.485*	*1.6825*	*2.0713*
Wald Stats	11.4885	4.1313	10.0425	7.8104	12.738	10.1588	6.5247
Wald p	0.0744	0.6589	0.1229	0.2523	0.0474	0.1181	0.3670

Estimates of System DGLS coefficient based on equation (8) using average monthly data with maximum leads and lags of 4months. The optimal lag length selected using Bayesian Information Criteria (BIC). The null hypothesis is individual coefficient is unity. The figure in the italics shows the standard errors. The Null hypothesis for the Wald test is the joint beta coefficient is unity. Wald P shows the P-values of the Wald test statistics. The symbols indicates *, < 5 % and ** < 10 % level of significance, respectively.

Table A3. Estimation results for System DOLS (SDOLS)

	6-m	7-m	8-m	9-m	10-m	11-m	12-m
				System DOLS			
Euro	-0.278	2.4236	3.5999	4.7121	6.152**	8.3515*	9.7004**
	2.0146	*2.1105*	*2.7217*	*2.8512*	*2.7927*	*3.7077*	*4.5603*
JPY	-0.7075	-1.048**	-1.1609	-1.1378	-0.7956	-0.6515	-1.0312
	1.2539	*1.1603*	*1.3721*	*1.3674*	*1.4037*	*1.7217*	*1.9135*
GBP	-0.6000	-0.1023	-0.6930	-0.4170	-0.6122	0.8814	3.7240
	1.7103	*1.5082*	*1.9771*	*2.0987*	*1.8826*	*2.4351*	*2.7847*
AUD	-0.9933	-0.9490	-0.5355	-0.5590	-0.6936	-0.7358	0.9697
	1.2610	*1.6110*	*2.2190*	*2.0494*	*1.8770*	*2.2867*	*2.6355*
CAD	0.0444	-0.3957	-0.9880	-2.2319**	-2.2188**	-3.0872**	-3.9464**
	1.4682	*1.7028*	*2.0284*	*1.8667*	*1.7915*	*2.1988*	*2.7878*
CHF	-1.4158	-4.0798*	-3.2265*	-2.1787	-1.8461	-1.2688	-0.1325
	2.6703	*2.2338*	*2.5293*	*2.4684*	*2.0253*	*2.3883*	*2.5069*
Wald Stats	12.2551	13.6237	11.5634	14.5290	13.6779	9.6855	7.6626
Wald p	0.0565	0.0341	0.0724	0.0243	0.0334	0.1385	0.2639

Estimates of System DOLS coefficient based on equation (8A) using average weekly data with maximum leads and lags of 12 weeks. The optimal lag length selected using Bayesian Information Criteria (BIC). The null hypothesis is individual coefficient is unity. The figure in the italics shows the standard errors. The Null hypothesis for the Wald test is the joint beta coefficient is unity. Wald P shows the P-values of the Wald test statistics. The symbols indicates *, < 5 % and ** < 10 % level of significance, respectively.

Table A4. Estimation results for Fully Modified GLS (FMGLS)

	6-m	7-m	8-m	9-m	10-m	11-m	12-m
Euro	-0.1343*	-0.1009*	-0.2798*	-0.2166**	0.9372	0.9682	2.0269
	0.3910	*0.4133*	*0.5771*	*0.6671*	*0.6569*	*0.7542*	*0.8333*
JPY	-2.5297*	-1.9882*	-2.6277*	-2.9759*	-3.1833*	-3.018*	-3.3997*
	0.7423	*0.6328*	*0.7704*	*0.7519*	*0.7710*	*0.9044*	*1.0604*
GBP	-2.7716*	-2.9399*	-3.4442*	-2.9375*	-1.594*	-1.1135*	-0.6979
	0.8325	*0.8597*	*0.9831*	*0.9786*	*1.0246*	*1.0299*	*1.1889*
AUD	-1.6014*	-1.957*	-2.1516*	-2.059*	-1.1999*	-1.4672*	-0.2616
	0.5065	*0.5582*	*0.7257*	*0.7719*	*0.7692*	*0.8853*	*0.9032*
CAD	0.6367	1.2165	0.5112	0.6486	-0.8737*	-0.0695	-0.3639
	0.7251	*0.8215*	*0.8769*	*0.9564*	*0.9486*	*1.0073*	*1.1083*
CHF	-0.0669*	0.0167**	-0.0223	0.1068	-1.5258*	-1.9338*	-2.948*
	0.5270	*0.5661*	*0.6754*	*0.8305*	*0.8261*	*0.9489*	*0.8932*

System fully Modified GLS (FMGLS) estimates on average weekly data with maximum leads and lags of 12 weeks. The optimal lag length selected using Bayesian Information Criteria (BIC). The null hypothesis is individual coefficient is unity. The figure in the italics shows the standard errors. The symbols indicates *, < 5 % and ** < 10 % level of significance, respectively.