EUROPEAN CENTRAL BANK

EUROSYSTEM

WORKING PAPER SERIES

NO 801 / AUGUST 2007

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EVIDENCE ON EMERGING ECONOMIES & NONLINEARITIES

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In 2007 all ECB publications feature a motif taken from the €20 banknote.

1 Earlier drafts of this paper have benefited from suggestions by Lucio Sarno as well as from comments by an anonymous referee, Geert Bekaert, Menzie D. Chinn, Marcel Fratzscher, Joe E. Gagnon, Georges Pineau, Michael Sager, Christian Thimann, participants in an ECB internal seminar and members of the International Relations Committee. In addition, the authors are grateful to Thierry Bracke, Matthieu Bussière, Ibrahim Chowdhury, Reint Gropp, Philipp Hartmann, Simone Manganelli, Kleopatra Nikolaou, Eoin O'Donovan and Chiara Osbat for useful discussions. The views expressed in the paper do not necessarily reflect those of the European Central Bank.

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ISSN 1561-0810 (print)
ISSN 1725-2806 (online)

CONTENTS

Abstract 4
Non-technical summary 5
1. Introduction 6
2. Methodology and data 8
3. Results 14
4. Conclusions 20
References 22
Figures and tables 25
European Central Bank Working Paper Series 37

ABSTRACT

This paper tests for uncovered interest parity (UIP) at distant horizons for the US and its main trading partners, including both mature and emerging market economies, also exploring the existence of nonlinearities. At long and medium horizons, it finds support in favour of the standard, linear, specification of UIP for dollar rates vis-à-vis major floating currencies, but not vis-à-vis emerging market currencies. Moreover, the paper finds evidence that, not only yield differentials widen, but that US bond yields do react in anticipation of exchange rate movements, notably when these take place vis-à-vis major floating currencies. Last, the paper detects signs of nonlinearities in UIP at the medium-term horizon for dollar rates vis-à-vis some of the major floating currencies, albeit surrounded by some uncertainty.

Key words: Uncovered interest parity, distant horizon, emerging economies, nonlinearities

JEL classification number: E43, F31, F41

Non-technical summary

What is the impact of expected exchange rate movements on bond yields? This question has been the natural focus of the literature on uncovered interest parity. Admittedly, a conventional view is that uncovered interest parity is appealing in theory but rejected empirically. But the conventional wisdom is starting to change, for two main reasons. First, recent research suggests that uncovered interest parity tends to hold for financial instruments of long maturities, with the evidence being restricted thus far to *mature* economy currency pairs. Second, other recent evidence indicates that the relationship is characterised by significant nonlinearities, i.e. regime changes, with findings being restricted thus far to *short horizons*.

Against this background, this paper aims at contributing to the literature by providing evidence on uncovered interest parity at distant horizons for a country coverage enlarged to emerging economies. In addition, the paper explores the existence of nonlinearities at the medium-term horizon, differently from previous literature which had focused on short-horizons. To this end, it tests for uncovered interest parity at long and medium horizons for the US and its main trading partners, including both mature and emerging market economies, to assess whether bond yield differentials react in anticipation of, and proportionately to, US dollar movements. Aside from yield differentials, the paper also decomposes the response of exchange rates to US and foreign yields separately to test whether expected exchange rate movements have similar effects on US and foreign bond markets. Finally, the paper considers the existence of nonlinearities in uncovered interest parity in the medium term.

Overall, the paper finds evidence for uncovered interest parity at long and medium horizons, with bond yield differentials often reacting in anticipation of – and proportionately to – future dollar movements over the next five to ten years, in particular vis-à-vis major floating currencies. By contrast, results for dollar rates vis-à-vis emerging market currencies are less supportive of uncovered interest parity. This suggests that for these currencies – which are, arguably, occasionally managed – the standard explanations put forth to explain the empirical failure of uncovered interest parity at short horizons, such as those relating to the existence of high and varying risk premia, are likely to be also relevant at these longer horizons. Moreover, the paper finds evidence that, not only yield differentials react, but that US bond yields do react in anticipation of exchange rate movements – notably when these take place vis-à-vis major floating currencies. Last, at the medium-term horizon, the paper detects signs of nonlinearities in uncovered interest parity for dollar rates vis-à-vis some of the major floating currencies, albeit surrounded by some uncertainty. These nonlinearities reflect, perhaps, variations in risk premia. The results vis-à-vis other currencies, including emerging economy ones, are less supportive of such nonlinearities, however.

1. Introduction

What is the impact of expected exchange rate movements on bond yields? This question has been the natural focus of the literature on uncovered interest parity. Admittedly, a conventional view is that uncovered interest parity is appealing in theory but rejected empirically. This is mostly true, however, for financial instruments with short maturities (one year or less), which are typically involved in carry trades.[1] But the conventional wisdom is starting to change, for two main reasons.

First, more recent research suggests that uncovered interest parity tends to hold for financial instruments of longer maturities, notably three years or more (Flood and Taylor, 1997; Cochrane, 1999; Alexius, 2001; Chinn and Meredith, 2004 and 2005; Chinn, 2006; Zhang, 2006), with the evidence being restricted thus far to *mature* economy currency pairs. Using bonds with maturities ranging from five to ten years, Chinn and Meredith (2004, 2005), Chinn (2006) and Zhang (2006) show that yield differentials explain almost perfectly future currency movements. Building on McCallum (1994), the former explain that fundamentals play more of a role in the long run, which ties down the behaviour of bond yields and exchange rates in line with uncovered interest parity.[2] In line with these results, Cheung et al. (2005) find that uncovered interest parity performs well in predicting exchange rate movements at long horizons, relative to other structural models of the exchange rate. Cochrane (1999) considers these findings as one of the "new facts in finance". A second reason underlying the change in the conventional wisdom that uncovered interest parity fails empirically is the recent evidence that the relationship is characterised by significant nonlinearities, i.e. regime changes, with this evidence being restricted thus far to *short horizons*. Allegedly, nonlinearities are due to limits to speculation or time-varying risk premia, inter alia (Lyons, 2001; Sarno et al., 2006; Baillie and Kiliç, 2006).[3]

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[1] Evidence that interest rate differentials tend to be negatively – rather than positively – correlated with future currency movements, thereby wrongly predicting their direction, dates back to Fama (1984). An early survey by Froot and Thaler (1990), for instance, reports an average correlation of about -0.9. Calling upon the existence of rational bubbles, learning about regime shifts or fundamentals, as well as so-called "peso problems", subsequent research has endeavoured to explain the overwhelming empirical rejection of UIP (see, for instance, Sarno, 2005, for a more recent survey). Chaboud and Wright (2005) find results in support of uncovered interest parity at the short horizon, but only over very short windows of data that span the time of the discrete interest payment.

[2] Conversely, in the short-run, monetary policy authorities tend to 'lean against the wind' in the face of an exchange rate depreciation, which explains why uncovered interest parity fails at short horizons empirically. Other explanations put forward include a possible segmentation between short-term and long-term debt security markets, in line with the "preferred habitat" hypothesis as well as differences in exchange rate expectations between the short and long horizon (see Chinn, 2006).

[3] The 'limits to speculation' hypothesis suggests that market participants select a trading strategy only if its expected Sharpe ratio (excess return per unit of risk) is larger than that of alternative strategies (Lyons, 2001; Sarno et al., 2006).

Against this background, this paper aims at contributing to the literature by providing evidence on uncovered interest parity at distant horizons for a country coverage enlarged to emerging economies. This extension is relevant against the background of earlier studies which had found that the relationship tends to hold at *short* horizons in these economies. Bansal and Dahlquist (2000) show indeed that the empirical evidence from emerging and lower-income developed economies is consistent with economic theory. A positive short-term interest rate differential (relative to US rates) explains a depreciation of the domestic currency, in line with uncovered interest parity. Similarly, Flood and Rose (2001) find considerable heterogeneity across countries and detect signs that uncovered interest parity at the short horizon holds better in crisis countries, where both exchange and interest rates display high volatility. Uncovered interest parity at longer horizons remains yet untested, notably due to lack of data. The paper takes advantage of the increasing availability of data on long-term domestic interest rates in emerging economies, whose local bond markets have significantly deepened in the last decade (Mehl and Reynaud, 2005; Jeanne and Guscina, 2006; BIS, 2007). Another contribution of the paper is to explore the existence of nonlinearities at the medium-term horizon (two years), differently from previous literature which had focused on short-horizons (three months at most). This extension is relevant given that the choice of maturity has proved essential in standard, linear, tests of uncovered interest parity.

To this end, the paper tests for uncovered interest parity at long and medium horizons for the US and its main trading partners, including both mature and emerging market economies, to assess whether bond yield differentials react in anticipation of, and proportionately to, exchange rate movements. The analysis is carried out with the most important bilateral dollar pairs (i.e. those used in the calculation of the effective exchange rate of the US dollar). Aside from yield differentials, the paper also decomposes the response of exchange rates to US and foreign yields separately to test whether expected exchange rate movements have similar effects on US and foreign bond markets. Finally, the paper considers the existence of nonlinearities in uncovered interest parity in the medium term.

To anticipate on the paper's main results, we find indeed evidence for uncovered interest parity at long and medium horizons, with bond yield differentials often reacting in anticipation of – and proportionately to – future dollar movements over the next five to ten years vis-à-vis major floating currencies. By contrast, results for dollar rates vis-à-vis emerging market currencies are less supportive of uncovered interest parity. Moreover, we find evidence that, not only yield differentials react, but that US bond yields do react in anticipation of exchange rate movements – notably when these take place vis-à-vis major floating currencies. Last, at the medium-term horizon, we detect signs of nonlinearities in uncovered interest parity for dollar rates vis-à-vis some of the major floating currencies, albeit surrounded by some uncertainty. These nonlinearities reflect, perhaps, variations in risk premia. The results vis-à-vis other currencies, including emerging market ones, are less supportive of such nonlinearities, however.

The rest of the paper is organised as follows. Section 2 describes the methodology and the data. Section 3 presents the results. Section 4 concludes.

2. Methodology and data

Methodology

(i) Linear specification

Similarly to Flood and Taylor (1997), Alexius (2001), Chinn and Meredith (2004, 2005), Chinn (2006) and Zhang (2006) we take the standard Fama (1984) equation as a starting point and estimate it over selected long horizons *k*, namely ten, five and two years:

$$E_t(s_{t+k}) - s_t = i_t^k - i_t^{k*} \Rightarrow \Delta_k s_{t+k} = \alpha + \beta(i_t^k - i_t^{k*}) + \varepsilon_{t+k} \tag{1}$$

where $E_t(.)$ is the expectation operator conditional on the information set available at time *t,* Ω_t; *s* the logarithm of the exchange rate (dollar price per unit of foreign currency); *i* the domestic (US) bond yield of maturity *k*; i^* the foreign (US trade partner) bond yield of equal maturity; Δ_k the *k*-period difference operator; and ε_{t+k} the residual (forecast error). UIP holds if $\hat{\beta} = 1, \hat{\alpha} = 0$.

Observations are overlapping by construction, implying moving average (MA) terms in the residuals of order *k*-1, so that standard errors have to be corrected for autocorrelation. Therefore, in line with Chinn and Meredith (2004, 2005) and Chinn (2006), we use GMM to correct the standard errors of the parameter estimates for MA serial correlation. We also report panel regression results, where all currencies are pooled and a fixed-effect estimator is used, which increases the efficiency of the estimation. This allows to account for the fewer (non-overlapping) observations left in long-horizon regressions relative to short-horizon ones (Bekaert et al., 2007).

We assume that investors hold their assets until maturity. The constant term α may reflect a constant foreign exchange risk premium as well as default risk or liquidity risk. Clearly, this is especially important for emerging economies. If $\hat{\beta} > 0.5$, then the expected currency change is more variable than the risk premium (the converse of the so-called ‘Fama-Hodrick-Srivastava’ hypothesis; see Hodrick and Srivastava, 1986; Froot and Frankel, 1989; Chinn, 2006); therefore, the risk premium does not play an important role as explanatory variable in this case.

The key estimate is the slope of the regression $\hat{\beta}$:

a) If $\hat{\beta}$ = 1, bond yield differentials explain perfectly (one-to-one) future currency movements;

b) $\hat{\beta}$ = 0 suggests that future currency movements are unrelated to bond yield differentials today;

c) $\hat{\beta}$ < 0, an estimate common in standard uncovered interest parity regressions at short horizons, suggests that bond yield differentials explain future currency movements systematically in the “wrong” direction.

An intuition of the regression is provided for in Figure 1 which plots the yield differential between US and euro area bonds (US Treasuries minus German Bund) at the ten-year maturity and the change in the dollar-euro over the subsequent ten years. It is apparent that the former is a good predictor of the latter. In other words, when the US dollar is expected to depreciate relative to the euro, US bond yields rise relative to euro area yields.

(ii) Decomposition of the exchange rate response to US and foreign yields

To assess whether expected exchange rate movements have similar effects on US and foreign bond markets, we further decompose the predicted response of exchange rates to both US and foreign yields separately. For instance, if the foreign economy is small or poorly integrated with the US financially, it cannot be excluded that the foreign bond market bears the brunt of the adjustment upon an expected depreciation of the US dollar relative to the foreign currency, with only foreign bond yields falling and US yields remaining unaltered. However, the standard regression aforementioned is ill-suited to capture potential dissimilarities as it constrains the elasticity of the exchange rate change with respect to US bond yields to be equal, but of opposite sign, to the elasticity with respect to foreign yields (i.e. β and -β, respectively):

$$\Delta_k s_{t+k} = \alpha + \beta(i_t^k - i_t^{k*}) + \varepsilon_{t+k} \Rightarrow \Delta_k s_{t+k} = \alpha + \beta i_t^k - \beta i_t^{k*} + \varepsilon_{t+k}$$

This assumption is relaxed in order to estimate the two elasticities separately, i.e one for the US bond yield i_t^k (denoted β_1) and another one for the foreign bond yield i_t^{k*} (denoted β_2)

$$\Delta_k s_{t+k} = \alpha + \beta_1 i_t^k + \beta_2 i_t^{k*} + \varepsilon_{t+k} \tag{2}$$

and a Wald statistic is used to test whether $\hat{\beta}_2 = -\hat{\beta}_1$.

(iii) Nonlinearities

Finally, we consider the existence of nonlinearities in uncovered interest parity at the medium-term horizon (two years). In particular, we modify the standard regression to capture potential regime changes in the relationship between exchange rates and yields that are driven by *large*, *unexpected* dollar movements.

The latter are proxied with the magnitude and sign of the *"surprise"* component in the exchange rate change relative to prior expectations (actual outcome minus market forecast from survey data of professional forecasters):

$$s_t - E_{t-k}(s_t)$$

To introduce nonlinearities in the Fama regression, we use a specification which draws from the smooth transition regression (STR) class of models. This specification allows parameters to change smoothly depending on the values taken by a driver of change (referred to as a 'transition variable'). STR models were initially introduced by Granger and Teräsvirta (1993) and recently applied by Baillie and Kiliç (2006) as well as Sarno et al. (2006) to UIP, although at the short-term horizon only. This involves reformulating (1) as follows

$$\Delta_k s_{t+k} = \alpha' + \beta'(i_t^k - i_t^{k*}) + [\alpha'' + \beta''(i_t^k - i_t^{k*})]\Phi[\gamma; z_t] + \varepsilon'_{t+k} \tag{3}$$

where the transition function $\Phi(.)$ is bounded between 0 and 1. Φ allows parameters to change smoothly, the speed of these changes being determined by the coefficient γ and the transition variable z_t. Here, z_t is the magnitude of the "surprise" component in the exchange rate change.

A key implication of (3) is that the nonlinear equivalent of β, denoted β^{NL}, i.e. the elasticity of the future exchange rate movement with respect to today's interest rate differentials, depends on z_t:

$$\beta^{NL} = \beta' + \beta''\Phi[\gamma; z_t]$$

Put it differently, it is a nonlinear regression which captures changes in β, the elasticity of the expected exchange rate change with respect to today's yield differential, as a function of "surprise" exchange rate movements. This helps assess whether the relationship between bond yield differentials today and future currency movements is characterised by regime changes, depending on whether the dollar has been depreciating or appreciating unexpectedly.

A potential challenge in using the "surprise" component in the exchange rate change as a transition variable, however, is that it should contain no additional information relative to that already embodied in the expected exchange rate change at time *t*, if the assumption of rational expectations holds. To test whether this assumption holds empirically – and check that the "surprise" component in the exchange rate change can be used validly as a transition variable – we regress the residuals of equation (1) on the latter. By the law of iterated expectations, the rational expectations hypothesis holds only if these two are orthogonal (the "surprise" component cannot explain what is not already explained by the expected exchange rate change). Rejection of this hypothesis indicates that the "surprise" component in the exchange rate change does contain information which is not embodied in the expected exchange rate change and, thereby, can be used as a transition variable.

The specification of the transition function can accommodate different nonlinear patterns. Two popular specifications are the exponential and logistic functions (Granger and Teräsvirta, 1993). The logistic function, used by Baillie and Kiliç (2006), is S-shaped and asymmetric and has the following properties: Φ: $\Re \rightarrow [0,1]$; $\lim_{z_t \rightarrow -\infty} = 0$; $\Phi(0)=1/2$ and $\lim_{z_t \rightarrow +\infty} = 1$[4]

$$\Phi[\gamma; z_t)] = \{1 + \exp[-\gamma(z_t)]\}^{-1}$$

Baillie and Kiliç (2006) use this specification to consider several transition variables which, as they put it, are linked to time varying risk premia. In our context, this function implies that

[4] Ibid.

the change in β is asymmetric. For instance, with large unexpected appreciations, $\Phi(.)$ tends to zero, so that (3) collapses to a standard, linear, Fama regression. Conversely, with large unexpected depreciations, $\Phi(.)$ tends to one, so that (3) becomes a different Fama regression.

The exponential function, used by Sarno et al. (2006), is bell-shaped, symmetric and has the following properties: Φ: $\Re \rightarrow [0,1]$; $\Phi(0)=0$ and $\lim_{z_t \rightarrow \pm\infty} = 1$[5]

$$\Phi[\gamma; z_t)] = \left\{1 - \exp[-\gamma(z_t)^2]\right\}$$

Sarno et al. (2006) use this specification and the Sharpe ratio as a transition variable to test the 'limits to speculation' hypothesis. In our context, this function implies that the change in β is symmetric. For instance, with *both* small unexpected depreciations and appreciations, $\Phi(.)$ tends to zero, so that (3) collapses to a standard, linear, Fama regression. Conversely, with *both* large unexpected depreciations and appreciations, $\Phi(.)$ tends to one, so that (3) becomes a different Fama regression.

To select the appropriate model specification, we follow the approach suggested by Granger and Teräsvirta (1993) and Teräsvirta (1998). We first test the general hypothesis of linearity against the alternative of nonlinearity. Upon rejection of linearity, we then discriminate between the logistic and exponential functions. The models are estimated with nonlinear least squares and robust standard errors. We follow Granger and Teräsvirta (1993) and Teräsvirta (1998) and divide the transition variable by its sample standard deviation to use an initial guess of unity for γ and the linear estimates for the remaining parameters.

Data[6]

As observed in Chinn and Meredith (2004, 2005) and Chinn (2006), short-horizon tests of uncovered interest parity have benefited from the availability of interest rate series that match closely theoretical requirements. Comparable data for the long horizon are trickier to obtain. This is particularly the case of long-term rates in offshore markets on liquid instruments of a known fixed maturity. Likewise, onshore instruments are often not immediately comparable due to differences in tax regime or capital controls. Data are sometimes available for shorter time spans, notably for emerging economies (Mehl and Reynaud, 2005; Jeanne and Guscina, 2006). Other challenges are that interest rate series are often for debt instruments with maturities that only proxy the posited horizon, and not the zero-coupon yields that would be

[5] Note that the values taken here by the transition variable are in practice bounded from below, given that the maximum depreciation of the foreign currency vis-à-vis the US dollar is -100%.
[6] Further details on the data can be found in Table 2.

exactly consistent with equations (1), (2) and (3).[7] Having said that, although the data tend not to be as "clean" as those used for short-horizon tests of UIP, we would expect, along with Chinn and Meredith (2004, 2005) and Chinn (2006), the coefficient on the interest differential in long-horizon regressions to be biased towards zero, and away from its hypothesised value of unity. Hence, the results we obtain should be conservative in nature.

The empirical analysis focuses on the main currencies included in the nominal effective exchange rate (NEER) of the US dollar, as available from Bloomberg. International trade linkages – together with liquidity, financial linkages and exchange rate policy – can be thought of being indeed among the main determinants of the distribution across currencies of a potential US dollar depreciation. Therefore, the paper considers bilateral rates of the US dollar vis-à-vis both mature economy currencies and emerging market ones which, taken as two groups, receive almost equal weights in the nominal effective exchange rate of the US dollar. The currency weights used to calculate the US's NEER are reported *pro memoria* in Table 1.

The data on bilateral exchange rates are taken from Bloomberg and were sampled at the monthly frequency. They are available from the early 1970's to mid-2006 for US dollar rates vis-à-vis mature economy currencies while, for the rates vis-à-vis the ten emerging market economies for which we have bond yield series, they are mostly available from the 1980's. To proxy market expectations for the transition variable of the nonlinear specification (3), we use survey data available from Consensus Economics.[8] They were available for all US dollar exchange rates vis-à-vis mature economy currencies as well as for six of the US dollar rates vis-à-vis emerging market currencies.[9]

The bond yield series are taken from Global Financial Data. They refer to benchmark government issues. The maturity of the benchmark bond is occasionally not strictly constant, although it is always that closest to the reference maturity. Bond yield series for mature economies are available for the two-year, five-year and ten-year maturities from the early 1970's (occasionally later) to mid-2006. As regards emerging economies, data are available for the ten-year maturity for three countries (Malaysia, Taiwan and Thailand), against seven

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[7] Admittedly, zero-coupon, constant maturity yields would be more appropriate. Unfortunately these data are not readily available on a cross-country basis. Alexius (2001) applies a correction to account for the absence of zero-coupon yields and obtains better results relative to those based on unadjusted data. Presumably using adjusted data in our context would have a similar effect, as noted in Chinn and Meredith (2004, 2005) and Chinn (2006).

[8] These data are the result of a monthly survey of between 120 to 240 prominent forecasters, with time series starting in the mid-1990s. They refer to the average of those forecasts (the median was not available).

[9] We also had survey data for US dollar rates vis-à-vis the Hong Kong dollar and the Saudi riyal. However, we discarded them from the final estimations, given the exchange rate peg maintained in these countries throughout the estimation period.

for the five-year maturity (Malaysia again, as well as Hong Kong, India, Korea, the Philippines, Saudi Arabia and Singapore), typically from the mid-1990's and occasionally before.[10] Some data are also available for Mexico for the five-year maturity, also since the mid-1990's.[11]

The period of generalised floating started in 1973. After allowing for a ten-year lag on the yield differential, the available estimation period is early 1983 to mid-2006. For the sake of comparability, the same estimation period is used for the five-year horizon. For both horizons, we also use a shorter estimation period (January 1983 to December 2004) to compare our estimates with those in Chinn (2006). For the two-year horizon, there are two sample periods: one that resorts to the full sample of observations and another which is restricted to the mid-1990s onwards. The latter matches the shorter sample used in the nonlinear estimations (as survey data of market expectations are available from the mid-1990s, at best). In addition, to account for the instability brought about by the string of crises affecting emerging markets in the 1990s (such as those in Latin America, emerging Asia and Russia), the estimation starts in 1999 for US dollar rates vis-à-vis emerging market currencies.[12] Last, our empirical analysis clearly distinguishes between those countries which have maintained a strict peg to the US dollar throughout the estimation period (for which uncovered interest parity is bound to fail) and the remaining ones.

3. Results

Linear estimates

Table 3 and 4 report our estimation results for the Fama equation (1) for the 10 year and 5-year horizon, respectively.

At the ten-year horizon, our estimates for US dollar rates vis-à-vis mature economy currencies are close to those of Chinn and Meredith (2004, 2005) and Chinn (2006).[13] In almost all cases, the estimated slope coefficient β is positive and significant, which stands in sharp contrast with the negative estimates that are typical of short-run uncovered interest parity regressions. Interestingly, the share of future exchange rate movements explained is also often higher than in short-horizon regressions, with e.g. a R^2 of one-half for the dollar-mark (euro) and one-third

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[10] Data for the long-term interest rate for the Philippines pertain to the primary market (unlike for others, which all refer to secondary market prices).

[11] Data for Saudi Arabia are available at the quarterly frequency only and are interpolated linearly to monthly frequency.

[12] We also have estimations using the full sample of available observations, which are not reported here to save space but are available upon request. Overall, the results remain comparable, barring some differences for some currency pairs.

[13] This also shows that they are robust to slightly higher frequency data (we use monthly observations while they used quarterly observations).

for the dollar-sterling. This echoes similar evidence reported in Cochrane (1999) for equity returns, whose predictability is shown to rise with the forecast horizon. Moreover, the point estimates for the dollar-Canadian dollar, dollar-mark (euro) and the dollar-sterling are not significantly different from unity, which suggests that UIP holds for these pairs at this horizon. This suggests that expected currency changes are more variable than the risk premium (as implied by the 'Fama-Hodrick-Srivastava' hypothesis).[14] The estimated slope for the dollar-yen is positive, but slightly lower in magnitude, at around 0.3, which may partly reflect the inclusion of the early 2000s in the sample, when the zero bound to nominal interest rates was binding and preventing a downward adjustment in Japanese bond yields. The dollar-kronor stands in contrast with the previous currency pairs, with a negative – albeit insignificant – estimate for β. Turning to the US dollar rates vis-à-vis the 3 emerging market currencies for which we have data at the 10-year horizon (the Malaysian ringgit, Thai bath and Taiwanese dollar), we also obtain point estimates for β that are positive and significant. They are smaller in magnitude than those for mature economy currencies, at about 0.2-0.3, however (barring a 0.7 estimate for the dollar-bath on the full sample of available observations). Table 3 also reports panel estimates for various currency groups. The estimated β stands at 0.5 when US dollar rates vis-à-vis all currencies are pooled but rises to 0.7 when the estimation is restricted to the major floating currencies (i.e. vis-à-vis the Canadian dollar, Deutsche mark (euro), Japanese yen and Pound sterling). Conversely, it decreases to 0.2 when the estimation is restricted to rates vis-à-vis emerging market currencies.

The estimates at the five-year horizon broadly confirm those at the ten-year horizon. Barring the dollar-Swiss franc and the dollar-yen, the estimated slope coefficient β is significantly positive for all mature currency pairs. Moreover, the point estimates for the dollar-Canadian dollar, the dollar-mark (euro) and the dollar-sterling are not significantly different from unity, which suggests that uncovered interest parity holds for these pairs at this horizon. Conversely, the evidence for the US dollar rates vis-à-vis the seven emerging market currencies for which we have data is mixed. For the dollar-won, β is insignificantly different from unity, which suggests that uncovered interest parity holds. As for the dollar-Indian rupee, β is significantly *above* unity, although this may be partly due to the small sample of available observations. For the remaining currencies, we obtain results that echo those of standard uncovered interest parity tests at short horizons, with estimated βs that are either insignificant (see the results for the dollar rates vis-à-vis the Hong Kong dollar, Malaysian ringgit, Mexican peso, Saudi riyal) or significantly negative (vis-à-vis the Philippines peso and Singapore dollar). This indicates that uncovered interest parity fails at this longer horizon. Clearly, for the dollar-Hong Kong dollar and dollar-riyal, the empirical failure of uncovered interest parity, even at long

[14] By contrast, the point estimates for the remaining currencies are not found to be significantly above 0.5.

horizons, is due to Hong Kong and Saudi Arabia's stringent peg to the US dollar. While Bansal and Dahlquist (2000) and Flood and Rose (2001) found evidence supporting uncovered interest parity at the short horizon for emerging market currencies, our results suggest that for these currencies – which are, arguably, occasionally managed – the standard explanations put forth to explain the empirical failure of uncovered interest parity, such as those relating to the existence of high and varying risk premia, are likely to be relevant at these longer horizons. Table 4 also reports panel estimates for various currency groups. The estimated β stands at 0.3 when US dollar rates vis-à-vis all currencies are pooled. In line with single equation results, β rises to 0.7 when the estimation is restricted to the major floating currencies and turns insignificant when the estimation is restricted to rates vis-à-vis emerging market currencies.

Decomposition of the exchange rate response to US and foreign yields

Not only do bond yield *differentials* widen, but there is also some evidence that *US yields* genuinely react in anticipation of a US dollar movements, albeit not systematically. Tables 5 and 6 report estimates based on equation (2), where the uncovered interest parity assumption of equal in magnitude, but opposite in sign, elasticities of the exchange rate change with respect to US and foreign bond yields is relaxed. The results confirm that, for a given US dollar movement, these two elasticities may differ, possibly reflecting dissimilarities in terms of relative economic size or financial integration with the US.

In this respect, at the ten-year horizon, a depreciation of the US dollar is preceded by an increase in US bond yields, albeit not vis-à-vis all currencies (see Table 5). US bond yields rise while foreign bond yields fall proportionately in anticipation of future depreciations of the US dollar vis-à-vis the Deutsche mark (euro), the Australian dollar and the Swiss franc. This suggests that an expected depreciation of the US dollar vis-à-vis these currencies may have ex ante implications for valuations in US bond markets. Conversely, for the remaining currencies, the local bond market bears the brunt of the adjustment. Higher US bond yields today explain a future depreciation of the US dollar vis-à-vis the Pound sterling, but to a lesser extent than do lower British bond yields. Moreover, higher Thai and Taiwanese bond yields today predict a depreciation of the domestic currency vis-à-vis the US dollar, in line with uncovered interest parity. However, the US dollar remains in this case insensitive to movements in US bond yields.

At the five-year horizon, US bond yields are found to rise ahead of a US dollar depreciation but, again, not vis-à-vis all currencies, notably emerging market ones. US bond yields rise while foreign bond yields fall proportionately in anticipation of a depreciation of the US dollar vis-à-vis the Canadian dollar, the Deutsche mark (euro) and the Swiss franc.[15] This confirms that a future depreciation of the US dollar vis-à-vis these currencies may have ex ante implications for valuations in US bond markets. As for the Australian dollar and the Swedish kronor, the results suggest that the impact of a rise in US bond yields is greater than proportional than the fall in foreign bond yields. Conversely, for most other currencies, the local bond market bears the brunt of the adjustment. Barring the Korean won, the estimated slopes for the remaining emerging market currencies are found to be "wrongly" signed with higher US bond yields predicting a stronger US dollar (vis-à-vis the Mexican peso, Malaysian ringgit, Philippines peso or Singapore dollar), which further confirms the failure of uncovered interest parity for these currencies at this horizon.

Interpretation

Overall, the results provide evidence that bond yield differentials may react ahead of larger US dollar movements. From a policy perspective, understanding the impact of expected exchange rate movements on bond yields is of relevance to the discussions on the large and persisting imbalances in current account positions globally. The potential role played by exchange rates as an adjustment mechanism is often central in these discussions. In this respect, while there is consensus that in emerging economies with large and growing current account surpluses, especially China, it is desirable that effective exchange rates move so that necessary adjustments will occur, the debate on the role of major floating currencies in the adjustment, and on the possible implications thereof, including on other financial asset prices, is more open. In this context, an aspect which has been particularly discussed in both policy and academic circles is whether a potential adjustment in the US dollar would have benign implications.

More specifically, one question debated is whether a potentially large depreciation in the US dollar might be associated with a large rise in US bond yields and adverse consequences on financial markets and growth (Volcker, 2005; The Economist, 2005a, or Roubini and Setser, 2005). The rise in US bond yields might be due to, inter alia, higher imported inflation, monetary policy tightening or a higher risk premium, Evidence on this remains scant. A recent contribution by Gagnon (2005) suggests, however, that concerns might be misplaced. In reviewing historical developments, Gagnon (2005) finds no evidence that bond yields rise after large and abrupt currency depreciations. On the contrary, they tend to decline. These results have received some attention (The Economist, 2005b). The relationship between bond yields and exchange rates is complex, however. It can be driven by third variables and causality is likely bidirectional. In particular, bond yields might rise not only in response to a

[15] However, the decline in German bond yields is significant at the 17% level of confidence only; the elasticities for the Swiss franc are not significant, when estimated separately.

currency depreciation but also *in anticipation* thereof, as uncovered interest parity theory would suggest.

In this respect, it is straightforward to compute the implied widening in yield differentials upon an expected large, protracted, US dollar depreciation from our estimates.[16] Figure 2 plots the yield differential today between US and euro area (German) bond yields at the ten-year maturity which is consistent with a 0.9 β-estimate and an array of cumulated changes in the dollar-mark (euro) over the subsequent ten years.[17] For instance, if the US dollar is expected to depreciate by ten percent in the next ten years (which is close to the historical average taken over the last three decades), the estimate implies that ten-year bond yields in the US have to be higher than the corresponding euro area bond yields by around 110 basis points per annum today.[18] On the other hand, if the expected depreciation over the next ten years is much larger, at 60%, which is close to the depreciation observed between 1985 and 1995, i.e. that following the Plaza agreement, the estimate implies that ten-year bond yields in the US should be today higher than the corresponding German bond yields by around 550 basis points per annum. This is very close to the actual yield differential observed in the mid-1980s, which peaked at around 560 basis points prior to the agreement.

Another perspective is provided by Figure 3 which plots the differential today between US and foreign yields consistent with (i) the estimated β for each currency pair and (ii) an expected 10% US dollar depreciation. This differential is inversely proportional to β. The lower is β (the more tenuous is the link between future currency movements and yield differentials), the wider the yield differential needed to predict a 10% depreciation of the US dollar vis-à-vis the corresponding currency. Note that specification (1) is symmetric: a sharp *appreciation* of the US dollar in the future and a positive β imply lower US bond yield and higher foreign bond yields today. For this reason, the large yield differential implied for low-β currency pairs (such as the dollar-ringgit in Figure 2) probably mirrors – aside from the risk premium – the periods of high foreign yields relative to US yields, notably those ahead of currency crises in emerging markets.

Nonlinear estimates

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[16] Arguably, an analysis restricted to uncovered interest parity cannot encompass all the forces driving bond and foreign exchange markets. Yet, it provides a well-established framework to analyse the relationship between interest rates and exchange rates. The paper's focus on testing for uncovered interest parity at distant horizons is all the more relevant in such a policy context as short-term movements in nominal exchange rates and bond yields might fail to produce real effects if they are insufficiently sustained, notably if they are volatile and subsequently reversed.

[17] McKinnon (2005) and Obstfeld (2006) carry out a similar exercise, but with the dollar-renminbi.

[18] Assuming that euro area bond yields stand at 3.8% (like in late 2006, taking German Bund yields as a proxy for the euro area), this is calculated as $\ln[(1+0.10)]=0.9\times 10[\ln(1+x)-\ln(1+0.038)] \Rightarrow x=0.049$ and x–0.038=0.110.

Considering a shorter, two-year, horizon, uncovered interest parity is overwhelmingly rejected using the standard, linear, regression framework, with negative or even insignificant β estimates (see Table 7).[19] This is in line with standard estimates at very short horizons, with interest rate differentials being either unable to predict future currency movements, or predicting them in the "wrong" direction systematically.

To investigate whether these results are due to a misspecification of equation (1) we explore the existence of possible nonlinearities driven by large "surprise" US dollar movements. To this end, we first test whether the assumption of rational expectations holds and regress the residuals on the "surprise" component in the exchange rate change. By the law of iterated expectations, the rational expectations hypothesis holds only if these two are orthogonal (the "surprise" component cannot explain what is not already explained by the expected exchange rate change). The results, reported in Table 8, indicate that the hypothesis of rational expectations is rejected in all cases, with the residuals being significantly correlated with the "surprise" component. This suggests that the latter adds information which is not embodied in the expected exchange rate change and can be used as a transition variable.

The existence of significant regime changes, driven by the sign and size of "surprise" exchange rate movements is further confirmed by the results of linearity tests à la Granger and Teräsvirta (1993) and Teräsvirta (1998), reported in Table 9 (see first column under the heading F_L). The hypothesis of linearity is strongly rejected. Moreover, there is evidence in favour of the LSTR specification for all major currency pairs and the dollar-Mexican peso (see second, third and fourth columns under the heading F_3, F_2, F_1). This suggests that large unexpected currency movements change the relationship between exchange rates and bond yields, with the change in β being asymmetric. This differs from Sarno et al. (2006) who find evidence for symmetric change, but is consistent with Baillie and Kilic (2006) who find support for asymmetries. As referred to in Baillie and Kilic (2006), one interpretation for the existence of such asymmetries is that they reflect changes in the risk premium, with risk-averse investors prone to be more sensitive to losses than to gains. Seen from the perspective of a US investor investing in foreign currency denominated assets indeed, a large "surprise" appreciation of the US dollar increases the currency risk premium, which leads uncovered interest parity to fail. Conversely, a large "surprise" depreciation of the US dollar decreases the currency risk premium, which leads uncovered interest parity to hold.[20]

The estimation results are reported in Table 10. The estimated transition parameter γ, which determines the speed of adjustment in the parameters, is significantly different from zero for

[19] Note that survey data on market expectations were often unavailable, which explains why the sample of currencies had to be restricted.

[20] Conversely, there is evidence of nonlinearities of ESTR form for the dollar-won and the dollar-Philippines peso, with *the conditional change in* β being symmetric.

all major currency pairs. Considering first the estimation results with the Canadian dollar and the euro, the currencies of the US's two main trade partners, we find a negative estimate for β' but a positive, larger, estimate for β''. This suggests that, when the US dollar has depreciated more than expected, β^{NL} (the conditional sum of β' and β'') tends to become positive and closer to 1, in line with UIP. The results vis-à-vis other currencies are less supportive of similar nonlinearities, however. Those for the US dollar rates vis-à-vis the Japanese yen and the Pound sterling underscore the existence of significant nonlinearities in the constant terms, but not in the slopes, with the estimated β'' being not significantly different from zero. As for the emerging market currencies, the changes in either the constant terms or the slope coefficients are not found to be significant. All the estimates are surrounded by some uncertainty, however, as we find in almost all cases evidence of significant autocorrelation and ARCH effects remaining in the residuals. This could suggest the existence of other forms of nonlinearities that are not captured by the models.

To illustrate how the estimated nonlinearities play out, Figure 4 and 5 plot the conditional elasticity estimated for the dollar-euro and the dollar-Canadian dollar against a range of "surprise" changes. The estimated transition parameters imply well-behaved, S-shaped, transition functions. In an environment when the US dollar has been appreciating more than expected vis-à-vis the euro, the elasticity is negative. Similarly with the standard, linear, estimate, this means that US bond yields would fall, relative to foreign yields, if the dollar were now expected to depreciate. Conversely, in an environment when the US dollar has been depreciating more than expected vis-à-vis the euro (above 15% in the last two years), the elasticity turns positive to reach the theoretical value of unity with even larger (above 25%) "surprise" depreciations. In this case, US bond yields would rise, relative to foreign yields, if the dollar were expected to continue to depreciate.

4. Conclusions

The paper has explored in detail the impact of expected exchange rate movements on bond yields from the perspective of uncovered interest parity. In so doing, the paper has aimed at extending the literature on uncovered interest parity at distant horizons, by considering a country coverage extended to emerging economies and by exploring nonlinearities at longer maturities than considered in previous research.

At long- and medium-term horizons, the paper has found support in favour of the standard, linear, specification of UIP for dollar rates vis-à-vis major floating currencies, but not vis-à-vis emerging market currencies. While Bansal and Dahlquist (2000) and Flood and Rose (2001) found evidence supporting uncovered interest parity at the short horizon for emerging market currencies, our results suggest that for these currencies – which are, arguably, occasionally managed – the standard explanations put forth to explain the empirical failure of uncovered interest parity, such as those relating to the existence of high and varying risk premia, are likely to be relevant at these longer horizons. Moreover, the paper has found evidence that, not only yield differentials react, but that US bond yields do react in anticipation of exchange rate movements – notably when these take place vis-à-vis major floating currencies, possibly reflecting similarities in terms of relative economies' size or financial integration with the US. At the medium-term horizon, the paper has detected signs of nonlinearities in UIP for dollar rates vis-à-vis some of the major floating currencies, but these are surrounded by some uncertainty. These nonlinearities reflect, perhaps, variations in risk premia. The results vis-à-vis other currencies, including emerging market ones, are less supportive of such nonlinearities, however.

Looking ahead, an aspect which might deserve more attention is the role of real variables. One question which remains, in particular, is whether real interest rate differentials anticipate movements in real exchange rates. As this aspect pertains to the literature on real interest rate parity, rather than on the standard uncovered interest rate parity, we will take this up in future research.

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Figure 1: Yield differentials as predictors of future currency movements

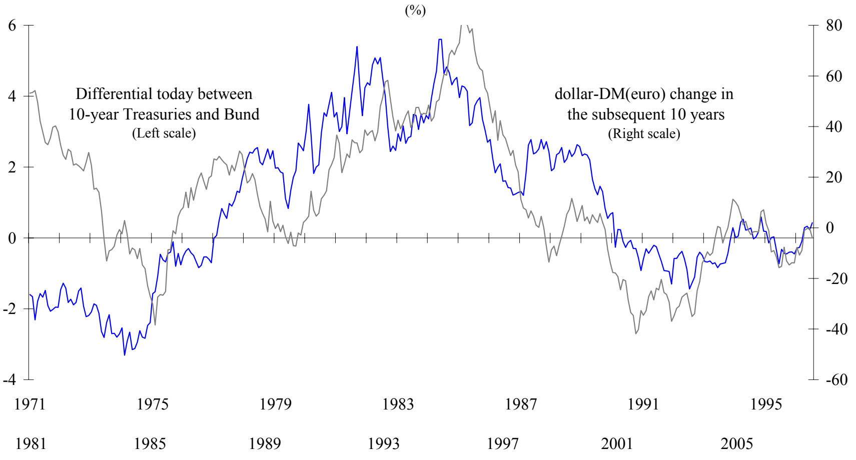
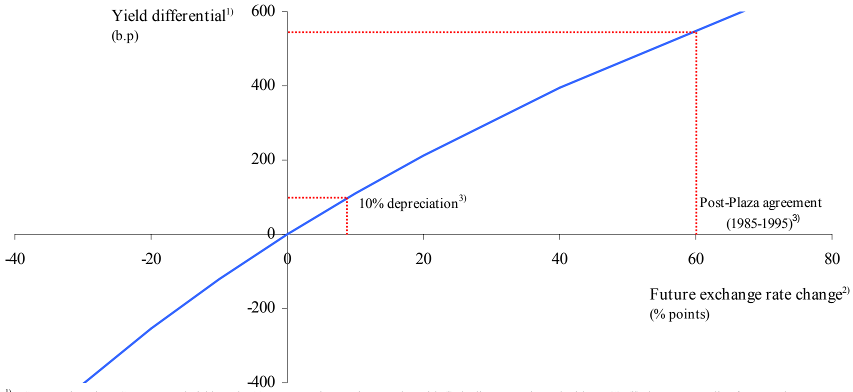
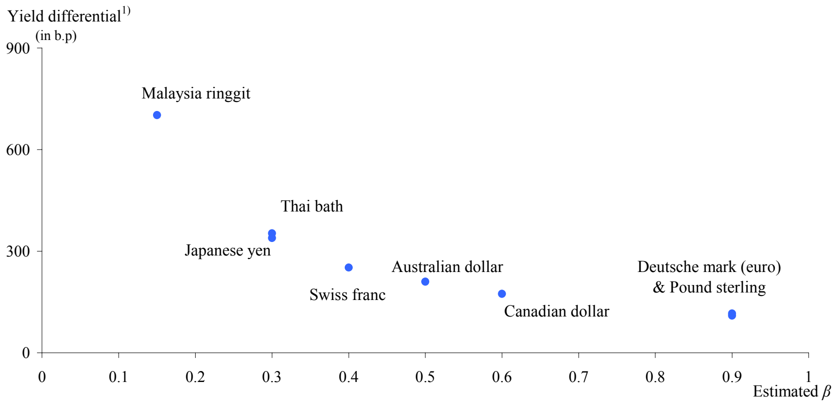


Figure 2: Ten-year yield differential consistent today with future changes in the USD/DEM(EUR)



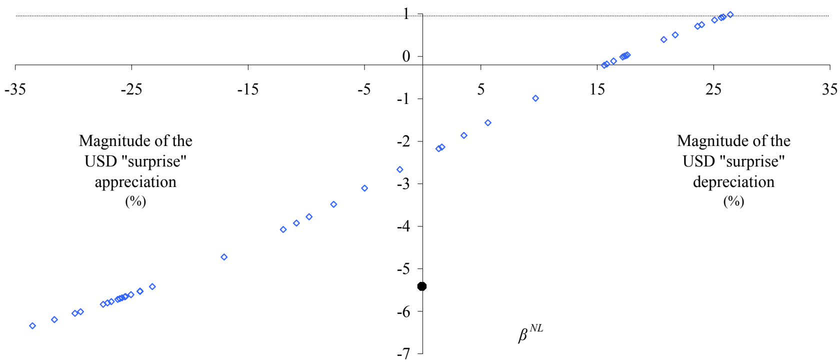
[1] US Treasuries minus German Bund yields at the 10 year maturity consistent today with (i) the linear β estimated with eq. (1); (ii) the corresponding future exchange rate change and (iii) assuming that 10-year Bund yields stand at 3.8% per annum.
[2] Cumulated exchange rate change over the subsequent 10-years; + = US dollar depreciation. [3] For illustration only.

Figure 3: Ten-year yield differential today consistent with a 10% US dollar depreciation in the future vis-à-vis selected currencies



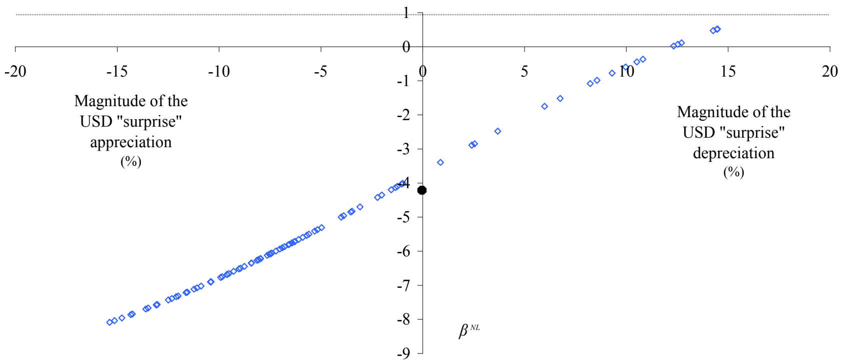
[1] US Treasuries minus foreign yields at the 10 year maturity consistent today with (i) the corresponding, linear, β estimated with eq. (1); (ii) a 10% US dollar depreciation and (iii) assuming that 10-year foreign yields stand at their corresponding historical average.

Figure 4: Estimated β^{NL} vs. "surprise" changes in the USD/EUR[1]



[1] Nonlinear β estimate using two-year bond yields and as transition variable the unexpected (i.e "surprise") change in the exchange rate change today relative to prior expectations two years ago, i.e. s_t - $E_{t-2y}(s_t)$. The black dot is the linear β estimate.

Figure 5: Estimated β^{NL} vs. "surprise" change in the USD/CAN rate[1]



[1] Nonlinear β estimate using 2 year bond yields and as transition variable the unexpected (i.e "surprise") component in the exchange rate change today relative to prior expectations 2 years ago, i.e. s_t - $E_{t-2y}(s_t)$. The black dot is the linear β estimate.

Table 1: Weights in the US dollar's nominal effective exchange rate (%)

Euro	18.1	Chinese renminbi*	13.4
Canadian dollar	16.3	Mexican peso	9.8
Japanese yen	10.0	Korean won	4.0
Pound sterling	4.8	Taiwanese dollar	2.8
Swiss franc	1.4	Malaysian ringgit	2.1
Australian dollar	1.2	Singaporean dollar	2.1
Swedish crown	1.2	Hong Kong dollar	2.0
		Brazilian real*	2.0
		Thai bath	1.4
		Indian rupee	1.1
		Israeli shekel*	1.0
		Russian rouble*	0.9
		Indonesian rupiah	0.9
		Philippines peso	0.8
		Saudi Arabian riyal	0.7
		Chilean peso*	0.6
		Other	0.6
Mature currencies	53.0	Emerging market currencies	46.2

Source: Federal Reserve.

Note: (*) currencies for which sufficiently long bond yield series were not available for estimation .

Table 2: Data overview

	Start of sample*		Start of sample*
Spot US dollar exchange rate vs.[1]			
(Industrial country currencies)		*(Emerging market currencies)*	
Australian dollar	January 1971	Hong Kong dollar	April 1974
Canadian dollar	January 1971	Indian rupee	January 1973
German mark	January 1971	Korean won	April 1981
Japanese yen	January 1971	Malaysian ringgit	January 1971
Pound sterling	January 1971	Mexican peso	January 1971
Swedish kronor	January 1971	Philippines peso	November 1991
Swiss franc	January 1971	Saudi riyal	December 1988
Euro	January 1999	Singapore dollar	January 1981
		Taiwanese dollar	October 1983
2-year ahead expected US dollar rate vs.[2]			
(Industrial country currencies)		*(Emerging market currencies)*	
Australian dollar	January 1995	Hong Kong dollar	November 2000
Canadian dollar	January 1995	Indian rupee	November 2000
German mark	January 1995	Korean won	November 2000
Japanese yen	January 1995	Malaysian ringgit	November 2000
Pound sterling	January 1995	Mexican peso	October 1995
Swedish kronor	January 1995	Philippines peso	November 2000
Swiss franc	January 1995	Saudi riyal	January 1995
Euro	January 1999	Singapore dollar	November 2000
		Taiwanese dollar	November 2000
		Thai bath	November 2000
10-year domestic bond yields[3]			
(Industrial sovereign issuer)		*(Emerging market sovereign issuer)*	
Australia	January 1970	Malaysia	January 1970
Canada	June 1982	Taiwan	January 1995
Germany	January 1970	Thailand	January 1980
Japan	January 1972		
Swedish kronor	January 1970		
Switzerland	January 1991		
United Kingdom	January 1970		
United States	January 1970		
5-year domestic bond yields[3]			
(Industrial sovereign issuer)		*(Emerging market sovereign issuer)*	
Australia	November 1970	Hong Kong	September 1994
Canada	June 1982	India	November 1994
Germany	January 1970	Korea	January 1970
Japan	January 1980	Malaysia	January 1992
Swedish kronor	January 1984	Philippines	January 1996
Switzerland	January 1991	Saudi Arabia	March 1992
United Kingdom	January 1970	Singapore	January 1988
United States	January 1970		
2-year domestic bond yields[3]			
(Industrial sovereign issuer)		*(Emerging market sovereign issuer)*	
Australia	January 1970	Hong Kong	November 1991
Canada	June 1982	Korea	January 1970
Germany	January 1970	Malaysia	January 1992
Japan	January 1980	Mexico	January 1995
Swedish kronor	January 1987	Philippines	January 1996
Switzerland	September 1996	Saudi Arabia	March 1992
United Kingdom	January 1979		
United States	January 1970		

Note: *End of sample is always July 2006. All the time series are sampled at the monthly frequency.
Sources: [1] Bloomberg.
[2] Consensus Economics.
[3] Global Financial Data.

Table 3: Long-horizon (10 years) Fama regressions (US dollar rates vs. respective currencies)

		Authors' estimates				Pro memoria: Chinn (2006)'s estimates		
	Sample	α	β	H_0: $\beta = 1$	R^2	α	β	R^2
Mature economy currencies								
Canadian dollar	01/83-06/06	-0.05 *	0.61 *	0.23	0.03			
		(0.03)	(0.32)					
	01/83-12/04	-0.04	0.86 ***	0.65	0.08	0.00	0.67 ***	0.09
		(0.03)	(0.29)			(0.00)	(0.13)	
Deutsche mark (euro post-1999)	01/83-06/06	-0.03	0.86 ***	0.29	0.48			
		(0.03)	(0.12)					
	01/83-12/04	-0.03	0.86 ***	0.25	0.47	0.00	1.02 ***	0.51
		(0.02)	(0.12)			(0.00)	(0.22)	
Japanese yen	01/83-06/06	0.20 ***	0.22	0.00	0.03			
		(0.04)	(0.17)					
	01/83-12/04	0.21 ***	0.30 **	0.00	0.07	0.02 ***	0.46 **	0.10
		(0.04)	(0.14)			(0.01)	(0.20)	
Pound sterling	01/83-06/06	0.05	0.85 ***	0.54	0.35			
		(0.03)	(0.23)					
	01/83-12/04	0.04	0.89 ***	0.64	0.36	0.00	0.76 ***	0.45
		(0.03)	(0.22)			(0.00)	(0.20)	
Other mature economy currencies								
Australian dollar	01/83-06/06	0.32 ***	0.53 ***	0.00	0.21			
		(0.03)	(0.13)					
Swedish kronor	01/83-06/06	-0.40 ***	-0.29	0.00	0.01			
		(0.09)	(0.36)					
Swiss franc	01/83-06/06	0.00	0.40 ***	0.00	0.21			
		(0.04)	(0.09)					
Emerging market currencies								
Malaysian ringgit	01/87-06/06	-0.30 ***	0.21 *	0.00	0.04			
		(0.03)	(0.13)					
	01/99-06/06	-0.46 ***	0.15 **	0.00	0.09			
		(0.01)	(0.06)					
Thai bath	01/83-06/06	-0.29 ***	0.73 ***	0.16	0.20			
		(0.05)	(0.19)					
	01/99-06/06	0.56 ***	0.29 ***	0.00	0.18			
		(0.03)	(0.08)					
Taiwanese dollar	01/83-06/06		...					
	03/05-06/06	-0.19 ***	0.25 *	0.00	0.22			
		(0.00)	(0.16)					
Pooled estimates								
All currencies	01/83-06/06	-0.03 ***	0.50 ***		0.20			
		(0.00)	(0.03)					
Only mature economy currencies	01/83-06/06	0.00 ***	0.52 ***		0.19			
		(0.00)	(0.03)					
Only major currencies[1]	01/83-06/06	0.01 ***	0.75 ***		0.44			
		(0.00)	(0.04)					
Only emerging market currencies	01/99-06/06	-0.49 ***	0.24 ***		0.42			
		(0.00)	(0.03)					

Notes: Estimation of equation (1) by GMM with lags used as instruments for the single currency pair equations. Fixed-effect estimates for the pooled regressions. Standard errors are robust to heteroskedasticity and autocorrelation and reported in parentheses. (***), (**), (*) indicate statistical significance at the 1%, 5% and 10% level of confidence, respectively. 1) Canadian dollar, German mark (euro), Japanese yen and Pound sterling. 2) *p* -value of the Wald statistic.

Table 4: Long-horizon (5 years) Fama regressions (US dollar rates vs. respective currencies)

		Authors' estimates				Pro memoria: Chinn (2006)'s estimates		
	Sample	α	β	H_0: $\beta=1$ [1)]	R^2	α	β	R^2
Mature economy currencies								
Canadian dollar	06/86-07/06	0.01	1.22 **	0.61	0.12			
		(0.02)	(0.45)					
	06/86-12/04	-0.01	0.92 **	0.85	0.08	0.00	0.51 *	0.02
		(0.02)	(0.42)			(0.00)	(0.33)	
Deutsche mark (euro post-1999)	01/83-07/06	-0.06	0.67 *	0.43	0.03			
		(0.04)	(0.41)					
	01/83-12/04	-0.09 **	0.83 **	0.68	0.04	0.00	0.90 *	0.07
		(0.04)	(0.39)			(0.01)	(0.53)	
Japanese yen	07/85-07/06	0.11 *	0.08	0.00	0.00			
		(0.06)	(0.33)					
	07/85-12/04	0.10 ***	0.19 ***	0.01	0.00			
		(0.06)	(0.33)					
Pound sterling	01/83-07/06	0.00	0.98 *	0.97	0.06			
		(0.04)	(0.56)					
	01/83-12/04	-0.02	0.89 *	0.84	0.05	0.00	0.51 *	0.02
		(0.04)	(0.55)			(0.00)	(0.31)	
Other mature economy currencies								
Australian dollar	01/83-07/06	0.13 ***	0.51 **	0.06	0.05			
		(0.04)	(0.26)					
Swedish kronor	03/89-07/06	0.00	0.78 ***	0.49	0.06			
		(0.05)	(0.31)					
Swiss franc	03/93-07/06	0.00	0.22	0.00	0.00			
		(0.04)	(0.28)					
Emerging market currencies								
(Constant peggers)								
Hong Kong dollar	01/83-07/06		…					
	11/99-07/06	0.00 **	0.01	0.00	0.02			
		(0.00)	(0.01)					
Saudi riyal	01/83-07/06		…					
	01/99-07/06	0.00	0.00	0.00	0.00			
		(0.00)	(0.00)					

(Others)					
Indian rupee	01/83-07/06		…		
	01/00-07/06	0.25 ***	1.57 ***	0.09	0.38
		(0.09)	(0.33)		
Korean won	01/83-07/06	0.10	0.96 ***	0.94	0.14
		(0.06)	(0.30)		
	01/99-07/06	0.02	1.03	0.96	0.08
		(0.17)	(0.84)		
Mexican peso	01/83-07/06		…		
	01/99-07/06	-0.16 ***	-0.06	0.00	0.01
		(0.04)	(0.11)		
Malaysian ringgit	01/83-07/06		…		
	01/99-07/06	-0.24 ***	-0.66	0.04	0.02
		(0.05)	(0.82)		
Philippines peso	01/83-07/06		…		
	03/01-07/06	-0.97 ***	-1.07 ***	0.00	0.12
		(0.20)	(0.39)		
Singapore dollar	01/93-07/06	0.09	-0.63	0.00	0.02
		(0.07)	(0.66)		
	01/99-07/06	0.15 **	-2.01 ***	0.00	0.44
		(0.06)	(0.48)		
Pooled estimates					
All currencies	01/83-06/06	-0.02 ***	0.30 ***		0.09
		(0.00)	(0.04)		
Only mature economy currencies	01/83-06/06	0.00	0.60 ***		0.05
		(0.00)	(0.06)		
Only major currencies[2)]	01/83-06/06	-0.02 ***	0.68 ***		0.10
		(0.00)	(0.10)		
Only emerging market currencies[3)]	01/99-06/06	-0.21 ***	0.04		0.01
		(0.01)	(0.04)		

Notes: Estimation of equation (1) by GMM with lags used as instruments for the single currency pair equations. Fixed-effect estimates for the pooled regressions. Standard errors are robust to heteroskedasticity and autocorrelation and reported in parentheses. (***), (**), (*) indicate statistical significance at the 1%, 5% and 10% level of confidence, respectively. [1)] *p*- value of the Wald test. [2)] Canadian dollar, German mark, Japanese yen & Pound sterling. [3)] Excluding constant peggers.

Table 5: Test of equality of the response of exchange rates (cumulated over a 10-year horizon) to US and foreign interest rates today

	Sample	β_1	β_2	H_0: β_2=-β_1 [1)]
Mature economy currencies				
Canadian dollar	01/83-06/06	0.41	-0.19	0.01
		(0.32)	(0.34)	
Deutsche mark (euro post-1999)	01/83-06/06	0.86 ***	-0.84 ***	0.88
		(0.11)	(0.22)	
Japanese yen	01/83-06/06	0.42 ***	0.78 ***	0.00
		(0.13)	(0.14)	
Pound sterling	01/83-06/06	0.68 ***	-1.21 ***	0.00
		(0.20)	(0.28)	
Other mature economy currencies				
Australian dollar	01/83-06/06	0.54 ***	-0.51 ***	0.80
		(0.13)	(0.14)	
Swedish kronor	01/83-06/06	-0.02	1.07 **	0.00
		(0.27)	(0.47)	
Swiss franc	01/83-06/06	0.35 ***	-0.50 *	0.58
		(0.09)	(0.27)	
Emerging market currencies				
Malaysian ringgit	01/99-06/06	0.35 ***	-0.02	0.00
		(0.04)	(0.09)	
Thai bath	01/99-06/06	-0.16	-0.79 ***	0.01
		(0.23)	(0.16)	
Taiwanese dollar	03/05-06/06	0.21	-0.48 ***	0.02
		(0.13)	(0.09)	

Note: Estimation of equation (2) by GMM with lags used as instruments. Standard errors are robust to heteroskedasticity and autocorrelation and reported in parentheses. (***), (**), (*) indicate statistical significance at the 1%, 5% and 10% level of confidence, respectively.

[1)] p-value of the Wald statistic.

Table 6: Test of equality of the response of exchange rates (cumulated over a 5-year horizon) to US and foreign interest rates today

	Sample	β_1	β_2	H_0: β_2=-β_1 [1)]
Mature economy currencies				
Canadian dollar	01/86-06/06	1.12 **	-1.19 *	0.57
		(0.45)	(0.45)	
Deutsche mark (euro post-1999)	01/83-06/06	0.71 *	-0.63	0.84
		(0.39)	(0.51)	
Japanese yen	07/85-06/06	0.74 ***	0.71 ***	0.00
		(0.24)	(0.21)	
Pound sterling	01/83-06/06	0.68	-1.48 ***	0.00
		(0.48)	(0.56)	
Other mature economy currencies				
Australian dollar	01/83-06/06	1.20 ***	-0.37 *	0.00
		(0.19)	(0.22)	
Swedish kronor	03/89-06/06	1.57 ***	-0.91 ***	0.02
		(0.40)	(0.29)	
Swiss franc	03/93-06/06	0.18	-0.12	0.89
		(0.38)	(0.33)	
Emerging market currencies				
(Constant peggers)				
Hong Kong dollar	11/99-06/06	-0.02	-0.02 ***	0.00
		(0.01)	(0.01)	
Saudi riyal	01/99-06/06	0.00	0.00	0.89
		(0.00)	(0.00)	
(Others)				
Indian rupee	01/00-06/06	-1.06 ***	-1.83 ***	0.00
		(0.37)	(0.25)	
Korean won	01/90-06/06	1.65 ***	-0.70 *	0.10
		(0.69)	(0.41)	
Mexican peso	01/00-06/06	-1.20 ***	0.13 *	0.00
		(0.37)	(0.08)	
Malaysian ringgit	01/99-06/06	-3.79 ***	-1.29 *	0.00
		(0.66)	(0.85)	
Philippines peso	03/01-06/06	-8.69 ***	-0.81	0.00
		(2.38)	(0.63)	
Singapore dollar	01/93-06/06	-1.74 ***	3.33 ***	0.05
		(0.69)	(0.75)	

Note: Estimation of equation (2) by GMM with lags used as instruments. Standard errors are robust to heteroskedasticity and autocorrelation and reported in parentheses. (***), (**), (*) indicate statistical significance at the 1%, 5% and 10% level of confidence, respectively. [1)] p-value of the Wald statistic.

Table 7: Medium-horizon (2 years) Fama regressions (USD vs. respective currencies)

	Sample	α	β	R^2
Mature economy currencies				
Canadian dollar	01/83-07/06	-0.01	-1.13 **	0.04
		(0.01)	(0.49)	
	01/97-07/06	0.00	-4.01 ***	0.52
		(0.01)	(0.52)	
Deutsche Mark (euro)	01/83-07/06	0.02	-0.42	0.00
		(0.02)	(0.55)	
	01/97-07/06	0.04 *	-5.12 ***	0.50
		(0.03)	(0.80)	
Japanese yen	01/83-07/06	0.18 **	-2.10 *	0.16
		(0.03)	(0.52)	
	01/97-07/06	0.12 **	-2.01 *	0.10
		(0.06)	(0.81)	
Pound sterling	01/83-07/06	-0.03	-1.12	0.05
		(0.02)	(0.82)	
	01/97-07/06	-0.02	-2.56 **	0.17
		(0.03)	(1.20)	
Emerging market currencies				
Korean won	01/83-07/06	-0.05	0.04	0.00
		(0.04)	(0.33)	
	01/99-07/06	-0.01	-1.03 **	0.02
		(0.04)	(0.52)	
Mexican peso	01/83-07/06	-0.05	0.11	0.04
		(0.03)	(0.10)	
	01/99-07/06	-0.10 **	-0.14	0.03
		(0.04)	(0.13)	
Philippines peso	01/83-07/06		…	
	01/99-07/06	-0.27 ***	-0.93 *	0.08
		(0.09)	(0.49)	

Note: Estimation of equation (1) by GMM. Standard errors are robust to heteroskedasticity and autocorrelation and reported in parentheses. (***), (**), (*) indicate statistical significance at the 1%, 5% and 10% level of confidence, respectively.

Table 8: Test of rational expectations

	μ	φ		μ	φ
Canadian dollar	0.00	0.37 ***	Korean won	0.02 ***	0.93 ***
	(0.00)	(0.04)		(0.00)	(0.03)
Euro	0.00	0.15 ***	Mexican peso	-0.02 **	0.55 ***
	(0.01)	(0.06)		(0.03)	(0.08)
Japanese yen	0.04 ***	0.94 ***	Philippines peso	0.11 ***	1.02 ***
	(0.00)	(0.06)		(0.00)	(0.05)
Pound sterling	-0.02 ***	0.69 ***			
	(0.00)	(0.05)			

Note: The table reports the estimated parameters and corresponding standard errors of the following OLS regression:

$\hat{\varepsilon}_t = \mu + \varphi[s_t - E_t(s_{t-k})] + \nu_t$

where $\hat{\varepsilon}_t$ is the residual of the linear equation (1), which is estimated under the assumption of rational expectations. By the law of iterated expectations, the rational expectations hypothesis holds only if the latter is orthogonal to the "surprise" exchange rate change relative to prior expectations $[s_t - E_t(s_{t-k})]$. Rejection of this hypothesis, i.e. $\varphi \neq 0$ indicates that the "surprise" exchange rate change can add information which is not embodied in the expected exchange rate change of the linear specification (1).

Table 9: Linearity tests on the 2-year horizon Fama regressions

	F_L	F_3	F_2	F_1
Transition variable: $s_t - E_{t-2y}(s_t)$				
Mature economy currencies				
Canadian dollar	0.00	0.23	0.00	0.00
Euro	0.00	0.00	0.05	0.00
Japanese yen	0.00	0.24	0.03	0.00
Pound sterling	0.00	0.08	0.00	0.00
Emerging market currencies				
Korean won	0.00	0.00	0.00	0.00
Mexican peso	0.00	0.01	0.75	0.00
Philippines peso	0.00	0.87	0.00	0.10

Note: The table reports the *p*-values from the linearity tests reviewed in Teräsvirta (1998). Given a transition variable z_t, one estimates the following auxiliary regression:

$$\hat{\varepsilon}_{t+k} = \boldsymbol{\beta}_0 \mathbf{A}_t + \boldsymbol{\beta}_1 \mathbf{A}_t z_t + \boldsymbol{\beta}_2 \mathbf{A}_t z_t^2 + \boldsymbol{\beta}_3 \mathbf{A}_t z_t^3$$

where the $\boldsymbol{\beta}$**s** are vectors of parameters, ε the residual from the corresponding Fama regressions reported in Table 7 and **A** the vector of explanatory variables in these regressions. A general test for linearity against nonlinearity of the STR form is the *F*-test of the null hypothesis: H_{0L}: $\boldsymbol{\beta}_1 = \boldsymbol{\beta}_2 = \boldsymbol{\beta}_3 = \mathbf{0}$. The choice between a LSTR and an ESTR model is based on a sequence of nested tests conditional on the rejection of H_{0L}, namely: H_{03}: $\boldsymbol{\beta}_3 = \mathbf{0}$; H_{02}: $\boldsymbol{\beta}_2 = \mathbf{0} \mid \boldsymbol{\beta}_3 = \mathbf{0}$; H_{01}: $\boldsymbol{\beta}_1 = \mathbf{0} \mid \boldsymbol{\beta}_2 = \boldsymbol{\beta}_3 = \mathbf{0}$. Again, an *F*-test is used, with the corresponding test statistics denoted F_3, F_2, and F_1, respectively. The decision rule is as follows: if the test of H_{02} has the smallest *p*-value, an ESTR specification is chosen, otherwise an LSTR specification is selected.

Table 10: Nonlinear estimates of the 2-year horizon Fama regression

	α'	α''	β'	β''	γ	ARCH(1)[1)]	AR(1)[1)]
Transition variable: $s_t - E_{t-2y}(s_t)$							
Mature economy currencies							
Canadian dollar	-0.05 **	0.21 ***	-10.95 ***	14.54 ***	1.01 ***	0.00	0.00
	(0.02)	(0.06)	(1.24)	(2.91)	(0.15)		
Euro	0.24 ***	-0.27 ***	-7.84 ***	11.33 ***	0.97 ***	0.21	0.00
	(0.02)	(0.03)	(1.87)	(3.45)	(0.07)		
Japanese yen	0.31 ***	-0.43 ***	-1.10 **	-0.91	0.96 ***	0.00	0.00
	(0.03)	(0.06)	(0.49)	(0.95)	(0.07)		
Pound sterling	0.18 ***	-0.48 ***	-4.13 ***	-1.39	0.97 ***	0.00	0.00
	(0.02)	(0.04)	(0.90)	(1.62)	(0.07)		
Emerging market currencies							
Korean won	-0.05 *	0.23	-1.34 ***	4.46	0.60 **	0.00	0.00
	(0.03)	(0.20)	(0.26)	(4.20)	(0.24)		
Mexican peso	0.39	-0.88	0.64	-1.65	0.45	0.10	0.00
	(0.51)	(1.02)	(0.95)	(1.88)	(0.60)		
Philippines peso†	-0.59 ***	0.27	-3.81 **	2.83	4.59 *	0.00	0.00
	(0.24)	(0.26)	(1.99)	(2.08)	(2.89)		

Notes: Estimation of equation (3) by nonlinear least squares. Standard errors are robust to heteroskedasticity and autocorrelation and reported in parentheses. (***), (**), (*) indicate statistical significance at the 1%, 5% and 10% level of confidence, respectively. All models have an LSTR specification except those signalled with a (†). 1) *p*- value of the test for ARCH effects of order 1 and autocorrelation of order 1, respectively. The AR test is constructed as in Eithreim and Teräsvirta (1996).

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ISSN 1561081-0