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Abstract

The forward premium anomaly refers to the situation where the slope coefficient in a regression of spot returns on the lagged interest rate differential is negative and significantly different to unity. This paper explores some of the asymmetries and non linearities present in the anomaly and the apparent rejection of Uncovered Interest Parity (UIP). The methodology is motivated by some recent economic theory literature on transactions costs, the limits to speculation and hysteresis. The paper estimates Logistic Smooth Transition Dynamic Regression (LSTR) models with the transition variable being the lagged forward premium for a range of currencies. An inner regime with foreign interest rates exceeding US rates is found to be consistent with the anomaly. While a third and outer regime with US interest rates exceeding foreign rates indicates convergence towards UIP. Detailed Monte Carlo experiments support the finding that an LSTR data generating process can indeed induce the forward premium anomaly. While the methodology appears promising in terms of uncovering important non linear and asymmetric behavior in the relationship, it should be noted that parameter estimation uncertainty indicates quite wide confidence intervals on the estimated transition functions. Hence, the accurate prediction of states, or regimes where UIP has a high probability of holding, is quite hard.

JEL Classification: C22, F31, F41.

Keywords: Forward premium anomaly, Uncovered Interest Parity, Non-linearity, LSTR models.

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1 Introduction

The theory of uncovered interest rate parity (UIP) states that the expected return, or rate of appreciation on a currency equals the interest rate differential, or equivalently the forward premium. One popular method for testing the theory has been to regress the rate of appreciation of the spot rate on the lagged forward premium. A test for UIP is then to test if the slope coefficient is unity, the intercept zero and the residuals serially uncorrelated. The forward premium anomaly is the widespread empirical finding of a negative slope coefficient in the above regression, so that the rate of appreciation of the spot exchange rate is negatively correlated with the lagged forward premium. This phenomenon has been consistently found for most freely floating currencies in the current float and appears robust to the choice of numeraire currency. Hence the forward premium anomaly implies that the country with the highest interest rate will have an appreciating currency, and not a depreciating currency, as implied by the theory of uncovered interest rate parity.

Many possible theories have been proposed to attempt to explain the anomaly. For example there has been an intensive study of modeling time dependent risk premia, e.g., Hodrick (1987, 1988) and Mark and Wu (1997); tests for the possible presence of a peso problem or bubble, e.g. Lewis (1995); and the modeling of irrational behavior of market participants and heterogenous trading behavior, e.g. Frankel and Froot (1987). Excellent surveys of the forward premium anomaly and suggested resolutions have been provided by Hodrick (1987) and Engel (1996). More recent work has emphasized the econometric issues involved in regressing spot returns, which are close to a martingale difference sequence with high volatility; on the lagged forward premium, which appears to be a highly autocorrelated, possibly long memory process, with relatively little volatility. These issues are analyzed by Baillie and Bollerslev (2000) and Maynard and Phillips (2001). However, the anomaly has yet to be resolved and is consequently a major puzzle in the field of international finance.

This paper takes a relatively different approach by considering some recent developments in non linear time series processes and specifically analyzes the form of nonlinearity and asymmetry in the relationship between spot returns and the lagged forward premium. This approach is partly motivated by some theoretical work on nonlinearities arising from the occurrence of significant transactions costs from closing arbitrage conditions in financial markets; see Baldwin (1990), Dumas (1992) and Hollifield and Uppal (1997). Similar issues concerning the presence of limits to speculation have been discussed by Lyons (2001); while Mark and Moh (2004) describe similar effects resulting from central bank intervention. These articles are discussed in more detail in the next section.

There has also been some applied research which has begun to address the nonlinear aspects of the forward premium anomaly. In particular, Bansal (1997) and Bansal and Dahlquist (2000) have found that the sign of the estimated slope coefficient in the forward premium regression tends to be related to the sign of the interest rate differential. The forward premium anomaly appears to occur when US interest rates are less than foreign rates. Baillie and Bollerslev (2000) also show that the magnitude and sign of the estimated slope coefficient in the forward premium regression appears to be slowly time varying, particularly in small sample sizes.

This paper attempts to formalize the presence of nonlinearities and asymmetries in the forward premium anomaly, by means of implementing some of the recent developments in non-linear time series analysis. In particular, this study considers a logistic smooth transition dynamic regression (LSTR) model. In our set up the speed of adjustment towards parity of the UIP condition is allowed to depend on the size of the forward premium. The resulting estimated LSTR models are shown to imply the existence of three regimes. The lowest regime occurs when the forward premium is less than the threshold level and gives rise to persistent deviations from UIP and is thus a domain that is consistent with the forward premium anomaly. This pattern is found to be broadly consistent across currencies, with the lower regime occurring most frequently in the early to late 1980s and late 1990s. The second or middle regime is interpreted as the transition between the lower and upper regimes. This state is characterized by relatively small forward premiums and with deviations from UIP being less extreme and less persistent. The third, or outer regime occurs when the US interest rate is substantially larger than the corresponding foreign interest rate, and is interpreted as being a regime with a relatively high probability of the UIP condition being satisfied. This outer regime corresponds to the late 1980s to the early 1990s. These results tend to explain why the forward premium anomaly is so pronounced in studies which utilized data from the 1970s and 1980s; and also why studies using data from the late 1980s and early 1990s do not find such strong evidence of the forward premium anomaly.

An important issue is the extent to which the LSTR model resolves the forward premium anomaly. Simulation experiments conducted in section 5 of the paper indicate that a data generating process (d.g.p.) with the LSTR model are entirely consistent with generating the anomaly from the sample sizes used in practice. Hence the inherent non linearity in the spot returns, forward premium relationship, when represented as an LSTR model would lead to downward biased slope coefficients in the forward premium regression. However, close examination of the estimated LSTR models over nine currencies reveals that inherent sampling variability from parameter estimation gives rise to considerable variability in the estimated transition functions between regimes. While the estimated LSTR models certainly explain some of the non linear aspects of the anomaly, they do not tell the whole story. Prediction of regime changes and hence the situations when UIP is most likely to hold is problematic. Several possibilities emerge as to the most useful direction for future research. One possibility is to use theoretical models of risk premia to specify non linear transition functions. Another possibility is to use more complex econometric procedures to statistically identify both the transition functions and also the types of non linearity in the relationship. However, the reported LSTR models in this study seem an important first step in the research towards incorporating non linear explanations in the forward premium anomaly.

The rest of the paper is organized as follows; section 2 briefly discusses some of the more important background literature. Section 3 discusses the econometric model and section 4 presents the empirical results. Section 5 then describes a detailed simulation study in which the LSTR model is found to generate data that is indeed consistent with some of the important stylized facts of the forward premium anomaly. The final section provides a brief conclusion.

2 Uncovered Interest Rate Parity and Nonlinearity

The theory of Uncovered Interest Rate Parity (UIP) is of central importance in international finance and states that,

$$E_t(\Delta s_{t+k}) = (i_{t,k} - i^*_{t,k}), \tag{1}$$

where $E_t(.)$ denotes the conditional expectation based on a sigma field of all relevant information at time t. The variable s_t is the logarithm of the spot exchange rate and is measured in terms of the number of dollars in terms of a unit of foreign currency at time t; $i_{t,k}$ and $i^*_{t,k}$ are the k periods to maturity nominal interest rates available on similar domestic and foreign assets respectively, while $\Delta s_{t+k} \equiv s_{t+k} - s_t$. Since covered interest rate parity is known to hold virtuously continuously, equation (1) can also be written as

$$E_t(\Delta s_{t+k}) = (f_{t,k} - s_t) = (i_{t,k} - i^*_{t,k}). \tag{2}$$

Hence the expected future k period rate of appreciation, (depreciation) is equal to the current forward premium (discount), denoted by $(f_{t,k} - s_t)$. Following Fama (1984), and assuming $k = 1$ for simplicity, it has been common to test the UIP hypothesis by embedding (2) into the regression framework of

$$\Delta s_{t+1} = \alpha + \beta(f_{t,1} - s_t) + u_{t+1}. \tag{3}$$

where UIP implies that u_{t+1} is a disturbance. The null hypothesis of UIP being valid, implies $\alpha = 0$ and $\beta = 1$ and u_{t+1} being serially uncorrelated. The forward premium anomaly refers to the fact that estimation of (3) has generally led to a significant and negatively valued slope coefficient. This finding is widespread regardless of choice of numeraire currency and sample period, although the finding is most extreme with data from 1973 through to the mid 1980s. In one survey, Froot and Thaler (1990) report that the mean value of $\hat{\beta}$ across 75 published studies is -0.88. Hence the forward premium anomaly implies that the country with the higher rate of interest has an appreciating currency, rather than a depreciating currency, as implied by the theory of UIP.

The standard explanations of time dependent risk premium, peso problems, and the learning by agents, have all been investigated; but no one theory has entirely resolved the anomaly. In contrast, this paper explores the asymmetry and non linear aspects of the anomaly. Some theoretical justification for considering non linearities are available from the work of Baldwin (1990), Dumas (1992), Sercu and Wu (2000), Lyons (2001) and others. In particular, Baldwin (1990) developed a partial equilibrium model with two assets denominated in terms of domestic and foreign currency. The model considers homogenous, risk neutral currency market traders who have small transaction costs of transferring between the two assets. These transaction costs, together with uncertainty, imply that the optimal level of cross currency interest rate speculation is marked by a first order hysteresis band. Spot returns are only influenced by interest rate differentials outside a band; so that small differentials have no effect. Hence it is only when the return on a currency is high enough, that forward looking behavior will induce investors to transact currency. In a different approach, Dumas (1992) developed a general equilibrium model of exchange rate determination in spatially separated markets with significant costs of international trade. The model implies that the nominal exchange rate will depend nonlinearly on the fundamentals, with the speed of adjustment to parity increasing in proportion to the deviation from parity. Hollified and Uppal (1997) also derive nonlinear relationships resulting from agents closing arbitrage conditions in financial markets.

More recently, Sercu and Wu (2000) proposed a model where transactions cause a bias in the forward premium anomaly regression, regardless of the possible existence of a risk premium. While Kilian and Taylor (2003) consider a non linear relationship between the level of a nominal exchange rate and its fundamental value. The form of non-linearity is driven by the presence of heterogenous traders and the strength of reversion to the fundamental level. Finally, Mark and Moh (2004) consider a continuous time stochastic process for the exchange rate which has a solution where the spot rate is a nonlinear function of the interest rate differential, which is represented by a jump-diffusion process regulated by occasional central bank intervention.

A further rationalization for the introduction of nonlinear dynamics in the UIP condition comes from the limits to speculation hypothesis of Lyons (2001). This model emphasizes the importance of the Sharpe Ratio in determining whether certain investment strategies are followed.[1] Lyons (2001) argues that if UIP holds exactly, so that $\alpha = 0$, and $\beta = 1$ in (3), then the associated Sharpe Ratio is zero. As the slope coefficient departs from unity, the Sharpe Ratio becomes positive. It is only when the Sharpe Ratio is higher than a threshold level, that the deviation from UIP will be high enough to be viewed as an arbitrage opportunity. Hence UIP is not expected to hold when the Sharpe Ratio on a currency strategy is higher than a threshold level. One implication is that limits to speculation creates a band of Sharpe Ratios, and hence a band of forward premia, where UIP does not hold.

While the forward premium has appeared to be a widespread phenomenon across the literature; e.g. see Froot and Thaler (1990), several studies have noted apparent empirical non linearities in the adjustment process for UIP. Bilson (1981), Flood and Rose (1994) and Huisman, Koedijik, Kool and Nissen (1998) all consider examples where extreme observations of the forward premium have proportionately more influence on forcing UIP to hold. A related study by Wu and Zhang (1996) finds evidence that the forward premium anomaly is asymmetric; while Zhou (2002) shows that UIP does not hold between 1980 and 1987. Bansal (1997) and Bansal and Dahlquist (2000) provide evidence that indicates that the spot rate appreciation when dollar denominated, responds differently to positive and negative interest rate differentials. The work suggests that the forward premium anomaly is far more likely to occur during periods where US interest rates are less than foreign interest rates. These stylized facts then motivate the development of the econometric approach to the anomaly in the next three sections of this study.

3 Dynamic Logistic UIP Regression

This section considers the search for a parsimonious, non linear model which adequately characterizes the dynamics between spot returns and the lagged forward premium. The main focus is to find a specification which has predictive power in understanding the reasons for the failure of UIP and the occurrence of the forward premium anomaly. Conversely, it is also desirable for the model to indicate regions where there is a high probability of UIP holding. Given the nature of the theoretical adjustment mechanisms based on hysteresis and the limits to speculation it appears appropriate to use a modeling approach that is based on smooth asymmetric adjustment, rather than discrete adjustment. For this reason a dynamic logistic smooth transition regression (LSTR) modeling approach is implemented in this study. The LSTR model is related to the Logistic Smooth Transition Auto-Regressive (LSTAR) models introduced by Granger and Teräsvirta (1993) and by Teräsvirta (1994). An excellent survey of STAR models is available in van Dijk et al. (2002). Similar to STAR models, the adjustment process in the LSTR model occurs in every period and the speed of adjustment is governed by the values of a transition variable. The use of a logistic function specification allows for sharp asymmetries in the adjustment processes. The LSTR model for the forward premium anomaly is

$$\Delta s_{t+1} = [\alpha_1 + \beta_1(f_{t,1} - s_t)] + [\alpha_2 + \beta_2(f_{t,1} - s_t)]F(z_t, \gamma, c) + u_{t+1}, \tag{4}$$

[1]The Sharpe ratio is defined as $\frac{E[R_s - R_{rf}]}{\sigma_s}$, where $E[R_s]$ is the expected return on the strategy, R_{rf} is the risk-free interest rate, and σ_s is the standard deviation of the returns to the strategy.

where u_{t+1} is a zero mean, stationary I(0) disturbance term, and $F(.)$ is the transition function which determines the speed of reversion. In this study $F(.)$ is chosen to be the logistic function,

$$F(z_t; \gamma, c) = (1 + \exp(-\gamma(z_t - c)/\sigma_{z_t}))^{-1} \quad \text{with } \gamma > 0, \tag{5}$$

where z_t is the transition variable, σ_{z_t} is the standard deviation of z_t, while γ is a slope parameter and c is a location parameter. The parameter restriction $\gamma > 0$ is an identifying restriction.

The logistic function (5), is bounded between 0 and 1, and depends on the transition variable z_t so that $F(z_t; \gamma, c) \to 0$ as $z_t \to -\infty$, $F(z_t; \gamma, c) = 0.5$ for $z_t = c$, and $F(z_t; \gamma; c) \to 1$ as $z_t \to +\infty$. When $\gamma \to \infty$, $F(z_t; \gamma, c)$ becomes a step function, such that the LSTR model becomes effectively a threshold model. Therefore, the LSTR model nests a two-regime threshold model. For $\gamma = 0$, $F(z_t; \gamma, c) = 0.5$ for all z_t, in which case the model reduces to a linear regression model with parameters $\alpha = \alpha_1 + 0.5\alpha_2$, and $\beta = \beta_1 + 0.5\beta_2$. The exponent in (5) is normalized by dividing by σ_{z_t}, which allows the parameter γ to be approximately scale-free. This is particularly useful for the initial estimates for the nonlinear optimization used to estimate the parameters in (4).

The values taken by the transition variable and the transition parameter γ will determine the speed of reversion to UIP. For any given value of z_t, the transition parameter γ determines the slope of the transition function and hence the speed of transition between extreme regimes, with low values of γ implying slower transition. An obvious choice for the transition variable is the lagged forward premium. This choice seems consistent with the previously discussed theoretical work of Dumas (1990) and others, which implies that the speed of adjustment to equilibrium is a function of the size of the deviation from equilibrium. Also, some empirical work such as Bansal (1997), noted how the extent of the forward premium anomaly appears to be related to the interest rate differential. Hence throughout this study $z_t = (f_{t,1} - s_t)/\sigma_{z_t}$; so that the transition variable is the Risk Adjusted Forward Premium (RAFP). [2] The parameter c can be interpreted as the *threshold* between the two regimes corresponding to $F(z_t; \gamma, c) = 0$ and $F(z_t; \gamma, c) = 1$, in the sense that the logistic function changes monotonically from 0 to 1 as z_t increases, while $F(c; \gamma, c) = 0.5$. Note that the inner regime corresponds to $z_t = c$, where $F(z_t = 0; \gamma, c) = \frac{1}{2}$ and equation (4) becomes a UIP regression of the form

$$\Delta s_{t+1} = [(\alpha_1 + 0.5\alpha_2) + (\beta_1 + 0.5\beta_2)(f_{t,1} - s_t)] + u_{t+1}. \tag{6}$$

The lower regime corresponds to for given γ and c to $\lim_{z_t \to -\infty} F(z_t; \gamma, c)$ where (4) becomes a standard linear UIP regression

$$\Delta s_{t+1} = [\alpha_1 + \beta_1(f_{t,1} - s_t)] + u_{t+1}, \tag{7}$$

while upper regime corresponds to $\lim_{z_t \to +\infty} F(z_t; \gamma, c)$ where (4) becomes a different UIP regression

$$\Delta s_{t+1} = [(\alpha_1 + \alpha_2) + (\beta_1 + \beta_2)(f_{t,1} - s_t)] + u_{t+1}. \tag{8}$$

[2] A further specification attempted in our work was to use the deviations from UIP; i.e. $z_t = (s_{t+1} - f_{t,1})$. Although the results were qualitatively quite similar to the forward premium being used, diagnostic tests suggested the deviations from UIP were statistically inferior to the forward premium. This is partly due to the variability of $z_{t+1} = (s_{t+1} - f_{t,1}) = (s_{t+1} - s_t) - (f_{t,1} - s_t)$ being dominated by the spot rate appreciation (noise) compared to the forward premium (signal). Full details of the results from using deviations from UIP can be obtained from the authors on request.

Hence the model in (4) is very general in the sense that it nests three regimes with quite different dynamics. Under the restrictions of $\alpha_1 + \alpha_2 = 0$ and $\beta_1 + \beta_2 = 1$, the upper regime corresponds to a domain where UIP has a high probability of holding. As the forward premium increases, so does the probability of the UIP condition being valid. Conversely, small or negative values of the forward premium are consistent with a regime with insufficient arbitrage incentives for agents to trade for UIP equilibrium to be attained. This set up is in the spirit of the theories concerning hysteresis effects resulting from transactions costs and the presence of limits to speculation in the currency market. The conditions for the validity of the UIP regime are tested later from the estimated models.

If the non linear model in (4) approximates the true d.g.p. of the UIP relationship, then the slope coefficient in the forward premium anomaly regression (3) can be expected to be different to unity. Also, if the majority of realizations of the forward premium are far from the outer regime, then the forward premium anomaly seems more likely to occur. This conjecture turns out to be fully supported by the simulation evidence presented in section 5 of this paper. A further point about (4) is that it has the interpretation of being an error correction model with the lagged forward premium being the deviations from a cointegrating relationship, with the transition function $F(.)$ playing the role of a non linear filter applied to the forward premium.[3]

A particularly important consideration in subsequent analysis turns out to be the choice of parametric transition function. The logistic transition function of the LSTR and LSTAR models appears considerably more general and flexible in this situation, than the ESTAR model, $G(z_t; \gamma) = 1 - \exp(-\gamma(z_t^2))$, with z_t again being the transition variable. The ESTR model inevitably imposes strong restrictions of symmetry that are not consistent with the data employed in this study.[4] A far simpler and less appealing set up would be to allow discrete switching from one period to another. This can be expressed as a regression of spot returns on an intercept and two separate variables involving the positive lagged forward premia and the negative lagged forward premia. The regression is then

$$\Delta s_{t+1} = \alpha + \beta^{+}(f_{t,1} - s_t)^{+} + \beta^{-}(f_{t,1} - s_t)^{-} + u_{t+1}, \tag{9}$$

where

$$(f_{t,1} - s_t)^{+} = \begin{cases} (f_{t,1} - s_t), & \text{if } (f_{t,1} - s_t) > 0 \\ 0, & \text{if } (f_{t,1} - s_t) < 0 \end{cases}$$

and

$$(f_{t,1} - s_t)^{-} = \begin{cases} 0, & \text{if } (f_{t,1} - s_t) > 0 \\ (f_{t,1} - s_t), & \text{if } (f_{t,1} - s_t) \leq 0. \end{cases}$$

where the variables $(f_t - s_t)^{+}$ and $(f_t - s_t)^{-}$ represent positive and negative forward premium respectively. This approach is essentially equivalent to separating the periods into those with positive and negative interest rate differentials, and was implemented by Bansal (1997) and Bansal and Dahlquist (2000).

[3] As discussed by Baillie (1989) there are several formulations of the spot and forward rate cointegrating relationship; and voluminous evidence that spot and forward rates are cointegrated with a coefficient of unity. Baillie and Bollerslev (1994) and Maynard and Phillips (2001) argue that the forward premium is well approximated as a long memory process, which suggests a form of fractional cointegration as developed by Granger (1986). The model in (4) implies a yet more complex form of cointegration.

[4] Since the first draft of this paper was written, the authors became aware of a working paper by Sarno et al. (2004), which uses ESTR models and deviations from UIP for a transition variable.

4 Empirical Results

This study uses monthly data on spot and one month forward exchange rates for the Belgian Franc (BF), Canadian Dollar (CD), Dutch Guilder (DG), French Franc (FF), German Mark (GM), Italian Lira (IL), Japanese Yen (JY), Swiss Franc (SF) and UK Pound (UKP) vis a vis the US Dollar. The data are provided by the Bank of International Settlements and are end of month mid rates, from December 1978 through December 1998 for the Eurozone currencies of the BF, DG, FF, GM and IL; and are from December 1978 through January 2002 for the other currencies of CD, JY, SF and UK.

The first panel of Table 1 reports the results from the conventional forward premium regression (3), and realizes results that are consistent with the forward premium anomaly. In particular, the estimates of the intercept α are close and generally not significantly different from zero, while the estimated slope coefficient β are negative in all cases except for the FF and the IL. Also, robust t statistics of the null hypothesis that $\beta = 1$ denoted by $t_{\beta=1}$ indicates rejection at conventional significance levels, except in the case of the FF.

The results in the second panel of Table 1 pertain to the estimation of the simple discrete switching model in equation (9). When the premium on the US dollar is positive, then the estimated slope coefficients are all positive, except for the CD and the UKP. When the forward premium is negative, the slope coefficients are all negative, and significantly different from unity in all cases except in the cases for the CD and the IL. The t test for the hypothesis that the slope coefficient is unity is rejected only in three cases (CD, JY, and UKP) at the .05 significance level in the state where forward premium is positive. While the same null hypothesis is rejected in seven out of nine cases in the sate when the forward premium is negative. Hence the overall evidence is that the forward premium anomaly is strongly related to situations when the US interest rate exceeds the foreign interest rate (i.e. when US dollar is quoted at a premium). Table 1 also reports the robust Wald test to test the equality of the slope coefficient for $(f_{t,1} - s_t)^+$ and $(f_{t,1} - s_t)^-$. This hypothesis is rejected for five out of the nine currencies. Hence this simple discrete switching model provides some, albeit informal evidence of non linearities in the relationship between spot returns and the forward premium. These results are broadly consistent with those of Wu and Zhang (1996), Bansal (1997) and Bansal and Dahlquist (2000).

The above results suggest the presence of nonlinearities and asymmetries in the relationship between spot rates and the lagged forward premium. The relationship is now more formally analyzed by utilizing the LSTR model for the UIP regression (4). Following Teräsvirta (1994) the model is estimated by nonlinear least squares, with the starting values obtained from a grid search over γ and c. Also, the transition variable is standardized by dividing it by the sample standard error of the transition variable, $\hat{\sigma}_{z_t}$. In the context of the UIP regression, the transition variable then has the interpretation of being the Risk Adjusted Forward Premium (RAFP).[5]

The dynamic LSTR for the UIP model (4) is estimated unrestrictedly in Table 2 and then subject to the restrictions, $\alpha_1 + \alpha_2 = 0$ and $\beta_1 + \beta_2 = 1$ in Table 3. Table 2 also reports the robust Wald test for testing the null of $\alpha_1 + \alpha_2 = 0$ and $\beta_1 + \beta_2 = 1$. The robust Wald test fails to reject these restrictions in all cases except for the CD at the .05 significance level.[6] The

[5]A further issue concerns the most appropriate definition of the sample standard deviation of the RAFP. One possibility with higher frequency data is to allow for time dependency and to use the conditional variance from a GARCH type model. Since monthly forward premium appear to have relatively little ARCH effects, the time invariant sample standard deviation has been used.

[6]The robust Wald test also fails to reject the null hypothesis for the CD at the 0.01 significance level. The restricted model is also chosen by the AIC and SIC criteria for all the currencies.

results in Table 2 are for the unrestricted estimation and indicate that except for the CD, all the estimates of the transition parameter γ, are significantly different from zero. Also, with the exception of the CD, all estimates of β_1 are negative and generally significantly different from zero, while estimates of β_2 are positive. The results in Table 2 generally indicate that the anomaly tends to occur for small and/or negative forward premium on the US dollar, while large forward premia are generally consistent with the UIP condition being less likely to be rejected.

The results of estimating the restricted LSTR UIP regression are presented in Table 3. Although the estimated transition parameter is generally highly significantly different from zero in all cases except for the BF, it has at least as much sampling variation as the corresponding results in the previous table. For the four currencies of the CD, FF, IL, and UKP, there is evidence of nonzero threshold levels ($\hat{c}$), and three of which are found to be statistically different from zero. For all cases the estimated value of β_1 is negative and that of β_2 is large and positive indicating that the UIP condition is more likely to hold as the transition function $F(.)$ converges to its asymptote of unity. Hence high values of the RAFP tend to push $F(.)$ towards the neighborhood of unity and towards the UIP condition holding. Conversely, when the premium on the US dollar is low, and sometimes negative, the transition function takes values in the neighborhood of zero, which tends to be associated with the forward premium anomaly.

Further insight into the nonlinear dynamics can be obtained from Figure 1, which displays the estimated transition function against the transition variable, together with the approximate values of the forward premia that correspond to $0 \leq F(.) \leq 1$. Clearly the estimated transition functions reveal well defined nonlinear dynamics with the spot returns and forward premium relationship being characterized by three relatively distinct regimes.

While adjustment toward UIP takes place once the forward premium on the US dollar reaches a certain threshold level, the speed of adjustment is determined by the size of the RAFP. There is a degree of uniformity across currencies in terms of adjustment speeds. However, given the relatively small sample sizes with monthly data of less than $T = 241$ for Eurozone currencies, the degree of uncertainty introduced from parameter estimation can be considerable in some cases. The difficulties involved in precisely estimating the transition parameter have been well documented by a number of authors, e..g. van Dijk et al (2002). These issues are carried over in terms of the estimation of the transition functions. The estimated transition functions are plotted against time for each currency in Figure 2 with the transition variable, i.e. the lagged forward premium, plotted in the corresponding panel below. Each estimated transition variable has its 95% confidence intervals plotted as broken lines around the solid line of the transition function. For the FF the sampling variability of the estimated transition function is particularly acute, while it is also substantial for the BF, CD, IL and UKP. Hence while the estimated LSTR models appear to describe the spot returns and forward premium relationship quite well in terms of very satisfactory diagnostic test statistics; the implied transition functions can be relatively poorly estimated. Hence the precise regimes where UIP can be expected to hold and conversely the regimes where the anomaly is dominant are hard to definitively evaluate.[7]

It is interesting to determine from the LSTR models the level of the risk premium that is required for the upper regime of the transition function to be reached. This information can be obtained from the panels of Figure 1 and the estimation results in Table (2). The forward

[7] Current research being conducted by the authors is focused on estimating the models for a wider range of information for the transition variables. More precision with measuring the transition variable and the threshold levels may well lead to improved estimation of the transition function.

premium is required to be approximately 0.4% for the BF and the CD; 0.1% for the DG; 0.15% for the GM; 0.2% for the JY; 0.5% for the FF; 0.55% for the UKP; and 1.0% for the IL.[8] Conversely, the forward premium anomaly is likely to exist when the forward premium on the US dollar is approximately less than -0.2% for the DG, GM, JY and SF; -0.4% for the BF; and -0.5% for the FF. Analogously the anomaly is likely to occur when the US dollar is quoted at a premium of approximately less than 0.2% for the CD; 0.6% for the IL; and 0.45% for the UKP. These results are consistent with either the transactions cost as well as the limits to speculation arguments.

The graphs in Figure 1 indicate that for many of the currencies the transition function is hovering just above zero. Hence most of these observations on the forward premium are generally less than the necessary threshold to induce reversion to UIP. This phenomenon can also be observed from the graphs of the transition function against time in Figure 2. For most currencies the transition function attains values closer to unity between 1989 and 1994 and closer to zero between 1979 and 1989 and also after 1994. This finding is interesting since US interest rates were generally less than foreign equivalents in the period 1989 through 1994. This finding may also explain why conventional UIP regressions tend to reject the UIP hypothesis less severely, when data from the 1990s used to test it; see the discussions in Baillie and Bollerslev (2000) and Flood and Rose (2002).

5 Simulation Evidence on the Anomaly

Given the above empirical findings, it is interesting to consider if the stylized facts of the forward premium anomaly can be obtained from calibrating a d.g.p. obtained from the estimated LSTR models in (4). The calibration is implemented by using estimated parameter values reported in Table (2), with independent and identically distributed Gaussian innovations superimposed on each model with the appropriate scaled innovation variance being used. The simulated data were generated for 50,000 replications with 241 observations in each replication (or sample), for the BF, DG, FF, GM, and IL; and 50,000 replications with 278 observations for the CD, JY, SF, and UKP. In each replication 100 additional observations were generated and then discarded to reduce the effects of initialization. Then for each replication the standard linear UIP regression (3) was estimated along with the predictability regression,

$$s_{t+1} - f_{t,1} = \alpha + \delta(f_{t,1} - s_t) + v_{t+1}, \tag{10}$$

where $\delta = \beta - 1$. The above regression has been used to study the predictability of UIP deviations using the forward premium as predictor variable; e.g. see Bilson (1981), Fama (1984) and Backus, Gregory and Telmer (1993). Under UIP δ should be zero; and a negative estimate of β in equation (3) implies that the estimate of δ in equation (10) is negative. This will clearly imply strong predicability of the deviations from UIP.

The panels of Table (3) present the results of the Monte Carlo experiments. In the first two rows of the table, the estimates of α and β obtained from the actual data are reported again (taken from Table (2)); while the other rows report the sample mean, denoted by $\bar{\alpha}^{sim}$, $\bar{\beta}^{sim}$; the median, denoted by $\tilde{\alpha}^{sim}$, $\tilde{\beta}^{sim}$ of the estimates over the 50,000 replications, together with their .05 and .95 percentiles from the simulated empirical distribution, which are denoted by $\alpha^{sim}_{5\%}$, $\beta^{sim}_{5\%}$ and $\alpha^{sim}_{95\%}$, $\beta^{sim}_{95\%}$. The seventh and twelfth rows of the table report

[8]These results can be seen from the estimated transition function graphs by looking at the region for the forward premium for which the transition takes value unity.

the value of the t statistic for testing the null hypothesis that $\bar{\alpha}^{sim} = \bar{\alpha}$ and $\bar{\beta}^{sim} = \bar{\beta}$ respectively.

The results in Table (3) reveal that true d.g.p.s of the LSTR models are consistent with data that generates the forward premium anomaly, since the estimated values of the β parameters are frequently negative. The average and median values of α and β are close to the estimates from the regression (3). In fact the t tests from seventh and twelfth rows of the table indicate that the average value of the estimates of α and β from the simulated data are not statistically significantly different from the estimates from the forward premium anomaly regression. Note also that the .05 and .95 percentiles of the simulated α^{sim} and β^{sim} contain the estimates of α and β from the forward premium anomaly regression (3).

The first two rows of the last panel of Table 3 report the estimate of δ and the value of the t statistic for testing the null hypothesis that $\delta = 0$ obtained from the actual data for each currency from the predictability regression (10). Consistent with earlier empirical evidence, the reported results indicate that the estimates of δ are more negative than the estimates of β for all cases and they are significantly different from zero for all cases except for the FF and the IL at conventional levels of significance. The last column of Panel D reports the values of the t statistic for testing the null hypothesis of $\bar{\delta}^{sim} = \hat{\delta}$. The results suggest that the average values of δ estimates from the simulated data are not statistically different from estimates obtained from the predictability regression (10). Hence the above simulation experiments suggest that a true d.g.p. of an LSTR model typically leads to rejection of the UIP hypothesis.

6 Conclusion

This paper has considered some of the asymmetries and more obvious non linearities present in the relationship between spot returns and the forward premium. The paper estimates LSTR models with the transition variable being the Risk Adjusted Forward Premium for nine currencies. An inner regime with foreign interest rates exceeding US rates is found to be consistent with data that is likely to generate the forward premium anomaly. A third and outer regime, where US interest rates exceed the corresponding foreign rates, is consistent with uncovered interest parity having a higher probability of holding. The estimated models are quite consistent with some recent theoretical work on hysteresis effects resulting from some types of transactions costs and/or the presence of limits to speculation in the currency market. The LSTR model also has statistical appeal and detailed Monte Carlo experiments further suggest that an LSTR data generating process can generate data that will induce the forward premium anomaly. The results appear a step in the direction of gaining further insight into the causes of the anomaly and also uncover some of non linearity and asymmetry present in the relationship. However, it should be noted that the estimated transition functions have considerable sampling variability. Hence, prediction of regimes where UIP has a high probability of holding; or the regime where the anomaly is most pronounced, is quite hard. Further research might usefully focus on developing models with transition functions grounded in more economic theory, and also more efficient estimation of transition functions and definitions of regimes.

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Table 1: Forward Premium Anomaly Regressions

	BF	CD	DG	FF	GM	IL	JY	SF	UKP
Reg 1									
α	0.001	0.002	-0.002	0.001	-0.002	0.001	-0.010	-0.004	0.006
	(0.002)	(0.001)	(0.002)	(0.002)	(0.002)	(0.003)	(0.003)	(0.003)	(0.002)
β	-0.814	-1.132	-1.598	0.032	-0.894	0.448	-2.728	-1.395	-2.526
	(0.703)	(0.378)	(0.730)	(0.650)	(0.669)	(0.751)	(0.684)	(0.605)	(0.819)
$t_{\beta=1}$	-2.580	-5.640	-3.559	-1.578	-2.831	-1.928	-5.450	-4.305	-4.305
$\overline{R}^2$	0.006	0.028	0.024	0.001	0.009	0.003	0.051	0.022	0.049
Reg 2									
α	-0.006	0.003	-0.013	-0.004	-0.010	-0.000	-0.014	-0.008	0.003
	(0.003)	(0.001)	(0.003)	(0.002)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
β^+	1.517	-1.457	3.736	1.294	2.912	0.646	5.582	2.396	-1.612
	(0.851)	(0.551)	(1.513)	(0.581)	(1.467)	(0.769)	(2.117)	(2.768)	(1.037)
$t_{\beta^+=1}$	0.001	-4.459	1.808	0.506	1.303	-0.460	2.117	0.504	-2.519
β^-	-7.059	-0.268	-5.620	-5.463	-3.594	-7.102	-3.672	-2.280	-5.480
	(1.798)	(1.085)	(1.144)	(1.628)	(1.097)	(6.973)	(0.761)	(0.765)	(1.993)
$t_{\beta^-=1}$	-4.482	1.169	-5.789	-3.970	-4.188	1.162	-6.139	-4.288	-2.762
W	13.585	0.699	17.457	12.792	8.776	1.127	6.092	2.178	2.346
$\overline{R}^2$	0.055	0.030	0.094	0.050	0.044	0.008	0.064	0.032	0.059

Key: Asymptotic standard errors are in parentheses below the corresponding parameter estimates. W is the robust Wald test statistic for testing $H_0 : \beta^+ = \beta^-$ in (9) and has an asymptotic χ^2 distribution.

Table 2: Results from Unrestricted LSTR UIP Regression:
$\Delta s_{t+1} = [\alpha_1 + \beta_1(f_t^1 - s_t)] + [\alpha_2 + \beta_2(f_t^1 - s_t)]F(z_t, \gamma, c) + u_{t+1}$,
where $F(.) = [1 + \exp(-\gamma(z_t - c)/\sigma_{z_t})]^{-1}$.

	BF	CD	DG	FF	GM	IL	JY	SF	UKP
α_1	-0.004	0.004	-0.016	0.002	-0.012	0.006	-0.008	-0.012	0.006
	(0.008)	(0.006)	(0.004)	(0.007)	(0.005)	(0.005)	(0.019)	(0.004)	(0.002)
β_1	-6.355	0.258	-6.498	-3.328	-4.137	-2.965	-2.951	-2.923	-4.373
	(2.742)	(2.793)	(1.412)	(2.430)	(1.448)	(1.699)	(2.403)	(0.857)	(0.978)
α_2	-0.003	-0.002	0.011	-0.009	0.009	0.034	-0.007	0.064	0.080
	(0.010)	(0.012)	(0.010)	(0.010)	(0.014)	(0.029)	(0.038)	(0.045)	(0.090)
β_2	8.010	-1.369	8.221	4.987	5.322	1.448	10.481	13.158	9.386
	(2.689)	(1.641)	(2.613)	(2.371)	(2.660)	(3.252)	(5.278)	(11.203)	(13.836)
γ	4.753	2.418	19.356	6.779	8.931	1.872	1.790	8.017	8.766
	(2.709)	(3.697)	(3.045)	(2.061)	(3.605)	(0.632)	(0.871)	(0.835)	(1.70)
c	.	0.002	.	0.001	.	0.007	.	.	0.005
	.	(0.003)	.	(0.003)	.	(0.002)	.	.	(0.001)
$Wald$	2.196	6.686	0.693	3.721	0.334	3.883	0.581	5.780	1.513
$LM(4)$	5.478	1.091	4.423	4.937	5.748	5.263	4.138	5.970	4.880
$LM(8)$	3.478	1.033	2.295	2.665	2.919	2.883	2.643	3.068	2.597
$pRNL$	0.144	0.294	0.249	0.188	0.568	0.599	0.779	0.068	0.934
AIC	-7.171	-9.100	-7.232	-7.209	-7.164	-7.258	-7.049	-7.017	-7.350
SIC	-7.084	-9.021	-7.144	-7.122	-7.077	-7.170	-6.971	-6.938	-7.271
Sample	241	278	241	241	241	241	278	278	278

Key: Asymptotic standard errors are in parentheses below the corresponding parameter estimates. The transition variable is the lagged risk adjusted forward premium. $Wald$ denotes the robust Wald statistic for the null hypothesis that $\alpha_2 + \alpha_1 = 0$ and $\beta_2 + \beta_1 = 1$; while $pRNL$ is the p value for the test of no remaining nonlinearity in the residuals; and $LM(4)$ and $LM(8)$ are LM tests for testing for the presence of serial correlation in the residuals up to lags 4 and 8 respectively. These tests are constructed as in Eitrheim and Terasvirta (1996). AIC and SIC are the Akaike and Schwartz Information Criteria respectively.

Table 3: Results from Restricted LSTR UIP Regression:
$\Delta s_{t+1} = [\alpha_1 + \beta_1(f_t^1 - s_t)] + [\alpha_2 + \beta_2(f_t^1 - s_t)]F(z_t, \gamma, c) + u_{t+1}$, where $\alpha_2 = -\alpha_1$, $\beta_1 = 1 - \beta_2$ and $F(.) = [1 + \exp(-\gamma(z_t - c)/\sigma_{z_t})]^{-1}$.

	BF	CD	DG	FF	GM	IL	JY	SF	UKP
$\alpha_1 = -\alpha_2$	-0.007	0.002	-0.017	-0.001	-0.013	0.006	-0.016	-0.011	0.006
	(0.010)	(0.001)	(0.004)	(0.001)	(0.005)	(0.004)	(0.004)	(0.005)	(0.002)
$\beta_1 = 1 - \beta_2$	-8.060	-0.785	-6.653	-3.649	-4.250	-1.264	-4.176	-2.700	-4.190
	(4.261)	(0.590)	(1.469)	(2.250)	(1.449)	(0.977)	(1.081)	(0.982)	(0.869)
γ	2.092	7.648	13.420	2.475	8.117	5.420	6.028	7.094	12.212
	(1.374)	(3.483)	(1.893)	(1.378)	(2.533)	(3.046)	(2.171)	(2.584)	(1.899)
c	.	0.003	.	0.002	.	0.008	.	.	0.005
	.	(0.001)	.	(0.003)	.	(0.002)	.	.	(0.000)
$LM(4)$	5.529	1.051	4.414	4.326	5.131	6.004	4.327	6.234	4.906
$LM(8)$	3.504	1.012	2.294	2.250	2.718	3.139	2.656	3.175	2.516
$pRNL$	0.160	0.185	0.231	0.264	0.564	0.696	0.372	0.064	0.903
AIC	-7.166	-9.071	-7.228	-7.204	-7.162	-7.228	-7.045	-6.984	-7.340
SIC	-7.078	-8.992	-7.140	-7.116	-7.075	-7.141	-6.967	-6.905	-7.261
Sample	241	278	241	241	241	241	278	278	278

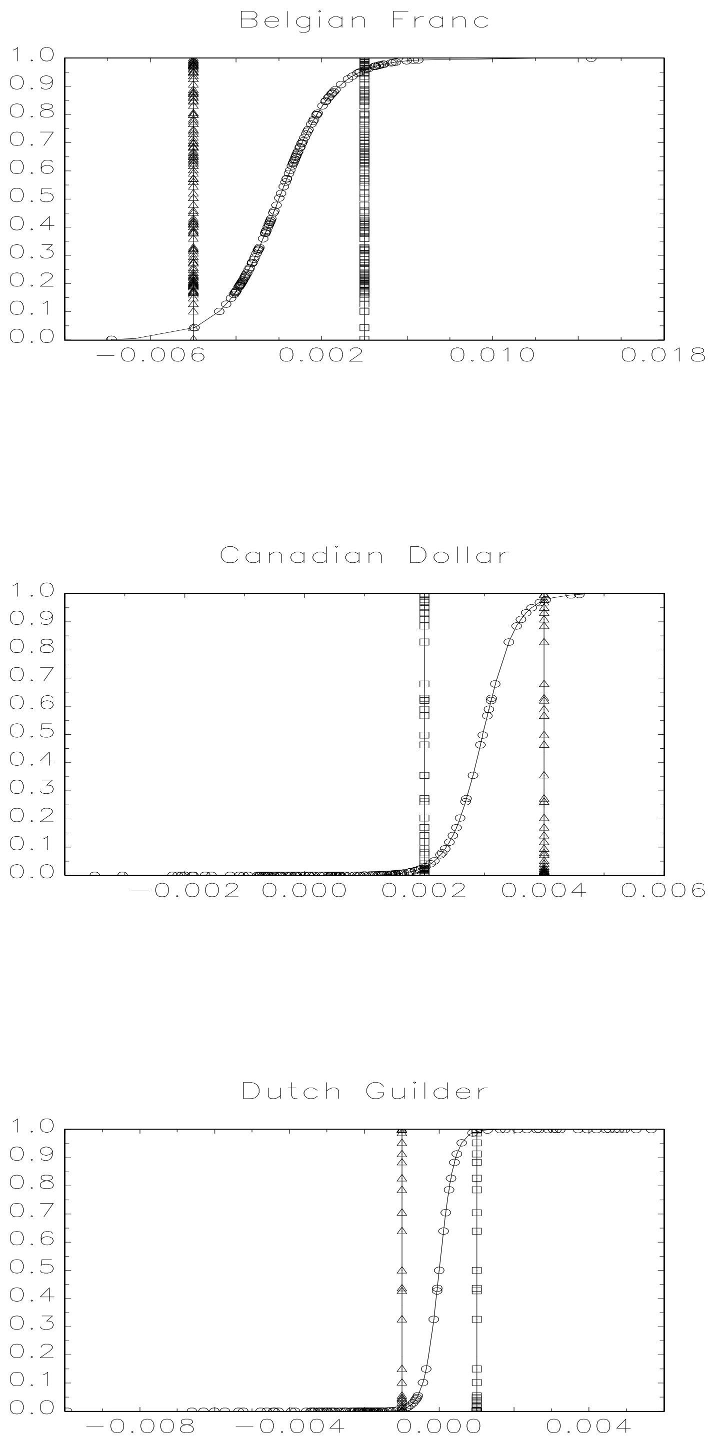
Key: As for Table 2

Table 4: Simulation Results: Results of Calibration of an LSTR Model

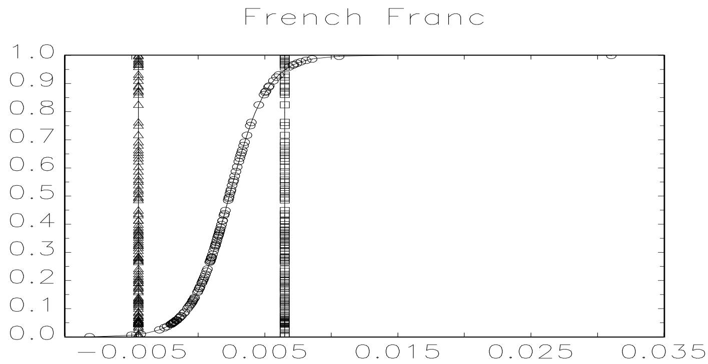
	BF	CD	DG	FF	GM	IL	JY	SF	UKP
α	0.001	0.002	-0.002	0.001	-0.002	0.001	-0.010	-0.004	0.006
β	-0.814	-1.132	-1.598	0.032	-0.894	0.448	-2.728	-1.395	-2.526
$\bar{\alpha}^{sim}$	-0.003	0.001	-0.008	-0.001	-0.006	0.004	-0.007	-0.006	0.004
$\tilde{\alpha}^{sim}$	-0.004	0.001	-0.008	-0.001	-0.006	0.004	-0.007	-0.005	0.004
$\alpha_{5\%}^{sim}$	-0.113	-0.111	-0.125	-0.115	-0.128	-0.177	-0.164	-0.138	-0.092
$\alpha_{95\%}^{sim}$	0.108	0.086	0.109	0.116	0.116	0.185	0.149	0.128	0.099
$t_{\bar{\alpha}}^{sim}$	-0.060	-0.015	-0.086	-0.022	-0.058	0.028	0.024	-0.018	-0.031
$\bar{\beta}^{sim}$	-3.710	0.178	-2.877	-1.091	-1.731	-0.295	-1.515	-0.870	-1.607
$\tilde{\beta}^{sim}$	-3.656	0.057	-2.982	-1.081	-1.847	-0.381	-1.483	-0.893	-1.507
$\beta_{5\%}^{sim}$	-43.321	-63.344	-42.396	-30.948	-38.393	-33.993	-41.510	-31.522	-35.631.
$\beta_{95\%}^{sim}$	35.831	63.703	36.815	28.891	34.976	33.810	38.355	29.872	32.054
$t_{\bar{\beta}}^{sim}$	-0.120	0.034	-0.053	-0.062	-0.038	-0.016	0.050	0.028	0.035
$\hat{\delta}$	-1.824	-2.123	-2.598	-0.968	-1.894	-0.552	-3.728	-2.395	-3.526
t_{δ}	-2.601	-5.654	-3.551	-1.491	-2.825	-0.737	-5.687	-3.954	-4.309
$\bar{\delta}^{sim}$	-4.710	-0.821	-3.877	-2.091	-2.731	-1.295	-2.515	-1.870	-2.607
$\tilde{\delta}^{sim}$	-4.656	-0.086	-3.982	-2.081	-2.847	-1.381	-2.483	-1.893	-2.507
$t_{\bar{\delta}}^{sim}$	-0.456	-0.007	0.155	-0.071	-0.029	-0.030	0.050	0.035	0.035

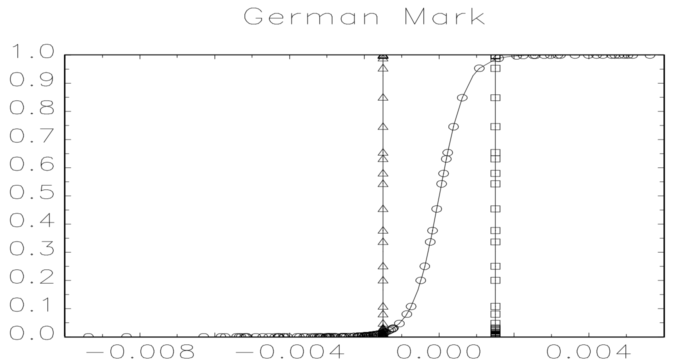
Key: The symbols α and β denote estimated parameters from the conventional forward premium regression in (1). The quantities $\bar{\alpha}^{sim}$, $\bar{\beta}^{sim}$ and $\tilde{\alpha}^{sim}$, $\tilde{\beta}^{sim}$ denote the mean and median of the empirical distribution (based on 50,000 replications) of the coefficients α and β respectively, obtained from estimating the regression (3) using simulated data from the LSTR UIP regression given in (4). ($\alpha_{5\%}^{sim}$, $\beta_{5\%}^{sim}$) and ($\alpha_{95\%}^{sim}$, $\beta_{95\%}^{sim}$) are the 5th and 95th percentiles of the empirical distribution of the parameters α^{sim}, β^{sim} respectively. $t_{\bar{\alpha}}^{sim}$ and $t_{\bar{\beta}}^{sim}$ are the $t-$values for the null hypothesis that $\bar{\alpha}^{sim} = \hat{\alpha}$ and $\bar{\beta}^{sim} = \hat{\beta}$, respectively. $\bar{\delta}^{sim}$ and $\tilde{\delta}^{sim}$ are the mean and median of the simulated distribution of the slope parameter in a regression of excess return $(s_{t+1} - f_t)$ on a constant and forward premium $(f_t - s_t)$. t_δ. $t_{\bar{\delta}}^{sim}$ is the value of the $t-$statistic to test the null hypothesis that $\bar{\delta}^{sim} = \hat{\delta}$.

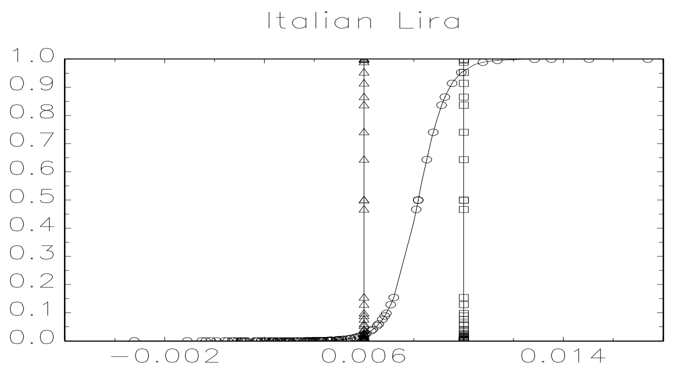
Figure 1: Estimated transition functions over transition variable

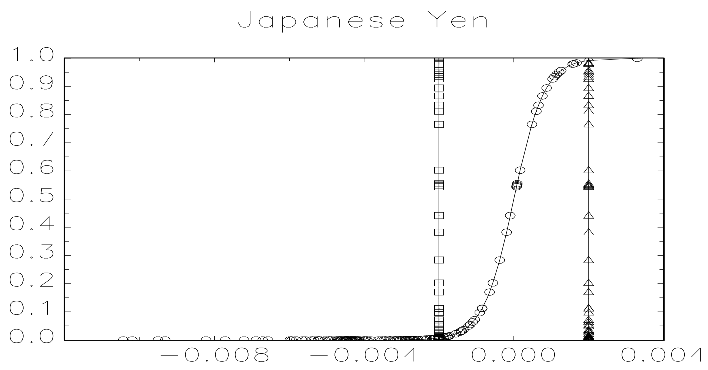


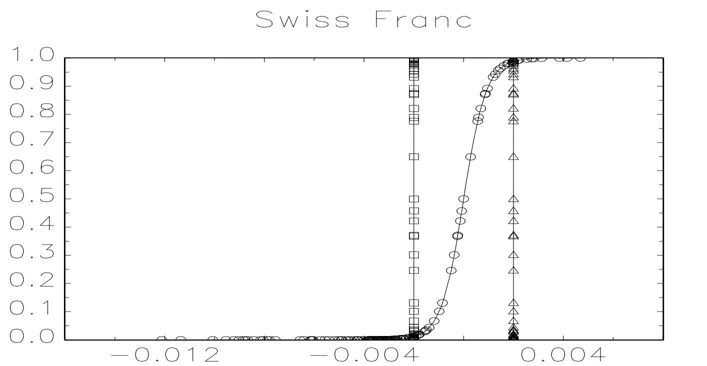
Key: Estimated transition functions over lagged forward premium are displayed. The vertical lines are approximate bands for which $F(.) \approx 0$ and $F(.) \approx 1$.


French Franc
1.0
0.9
0.8
0.7
0.6
0.5
0.4
0.3
0.2
0.1
0.0
−0.005
0.005
0.015
0.025
0.035


German Mark
1.0
0.9
0.8
0.7
0.6
0.5
0.4
0.3
0.2
0.1
0.0
−0.008
−0.004
0.000
0.004


Italian Lira
1.0
0.9
0.8
0.7
0.6
0.5
0.4
0.3
0.2
0.1
0.0
−0.002
0.006
0.014


Japanese Yen
1.0
0.9
0.8
0.7
0.6
0.5
0.4
0.3
0.2
0.1
0.0
−0.008
−0.004
0.000
0.004


Swiss Franc
1.0
0.9
0.8
0.7
0.6
0.5
0.4
0.3
0.2
0.1
0.0
−0.012
−0.004
0.004

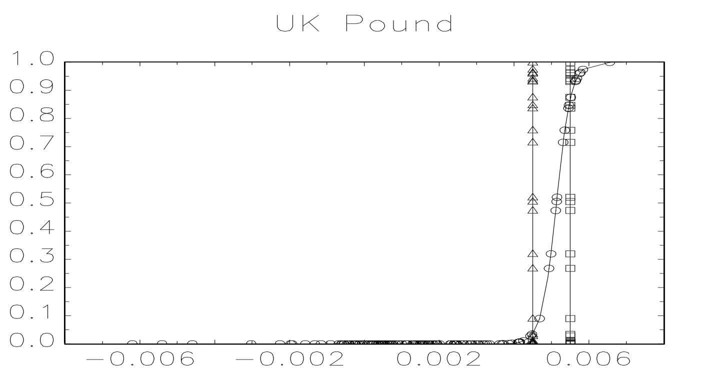
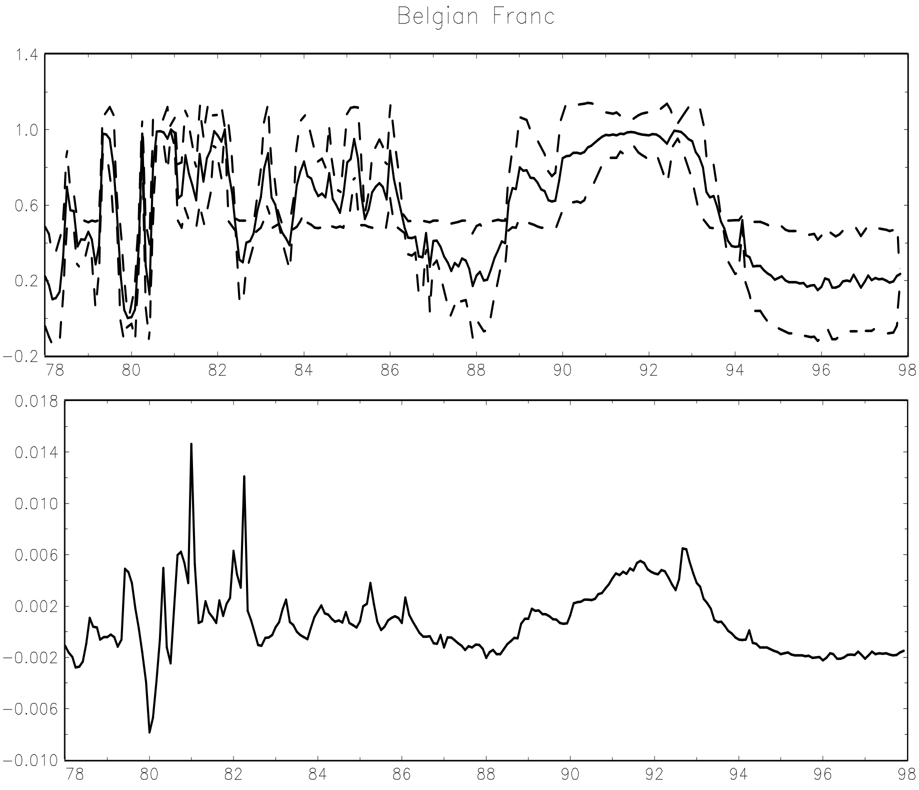
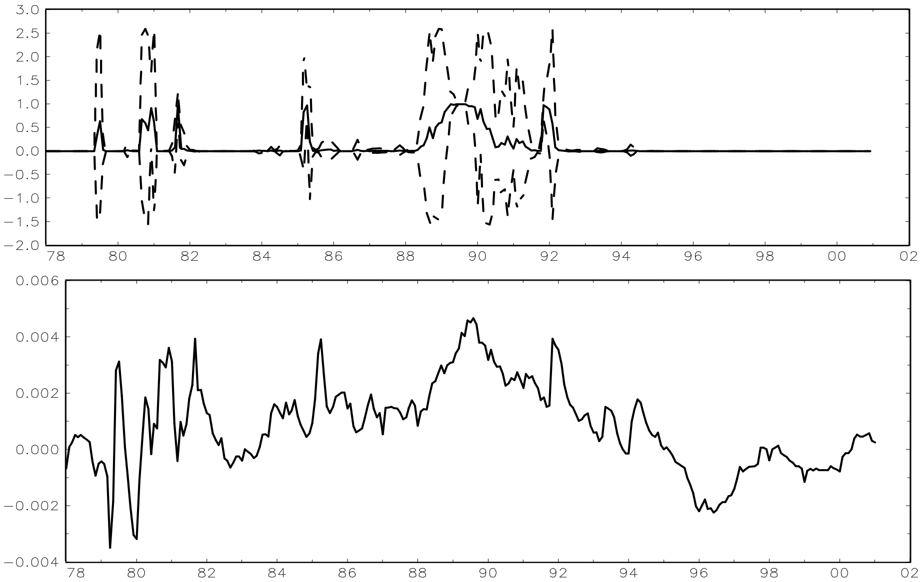

UK Pound
1.0
0.9
0.8
0.7
0.6
0.5
0.4
0.3
0.2
0.1
0.0
−0.006
−0.002
0.002
0.006

Figure 2: Estimated transition functions and forward premia over time

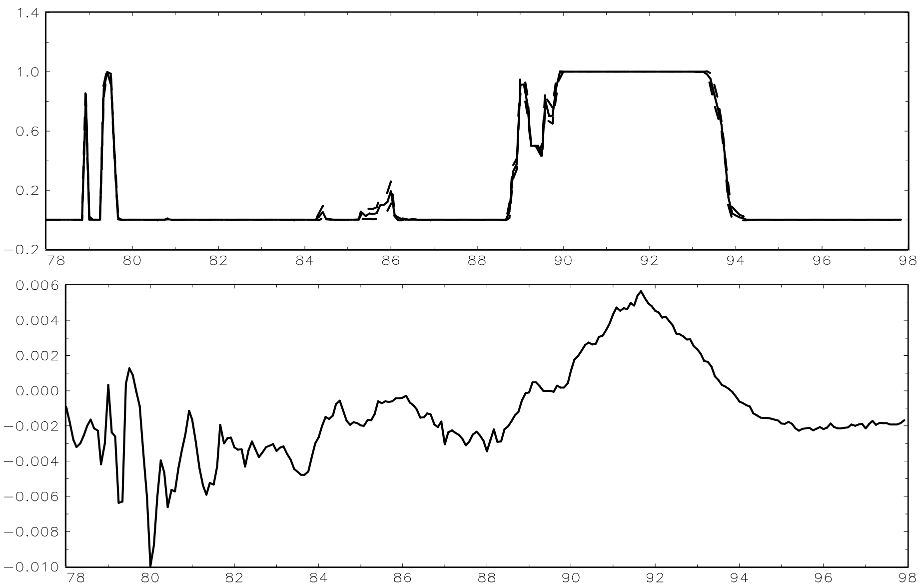


Key: Upper panels display estimated transition functions with 95% confidence intervals over the sample period and lower panels display forward premium over the sample period.

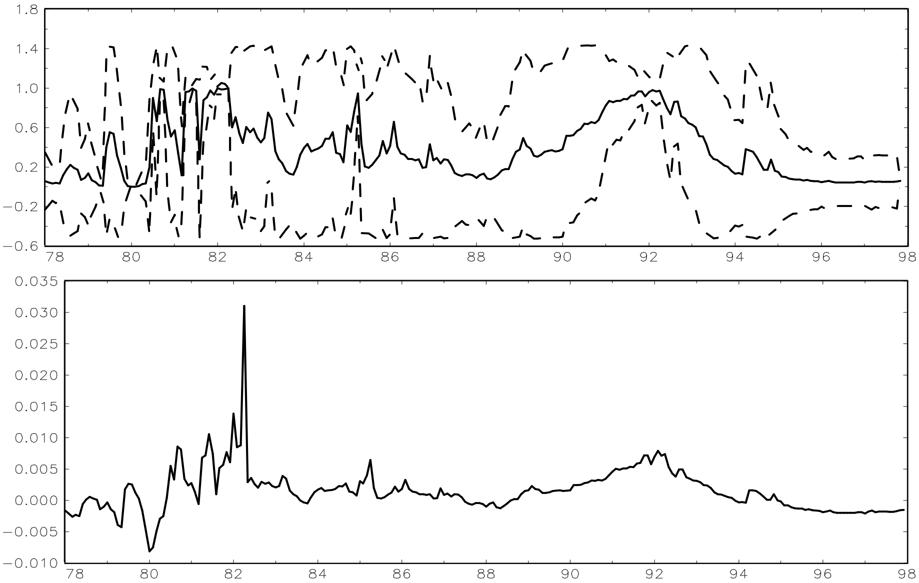
Canadian Dollar



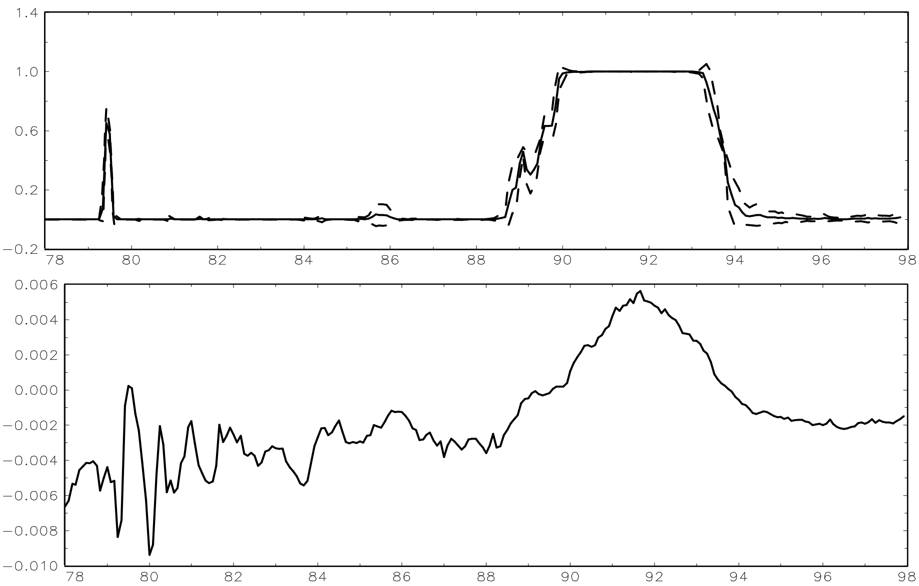
Dutch Guilder

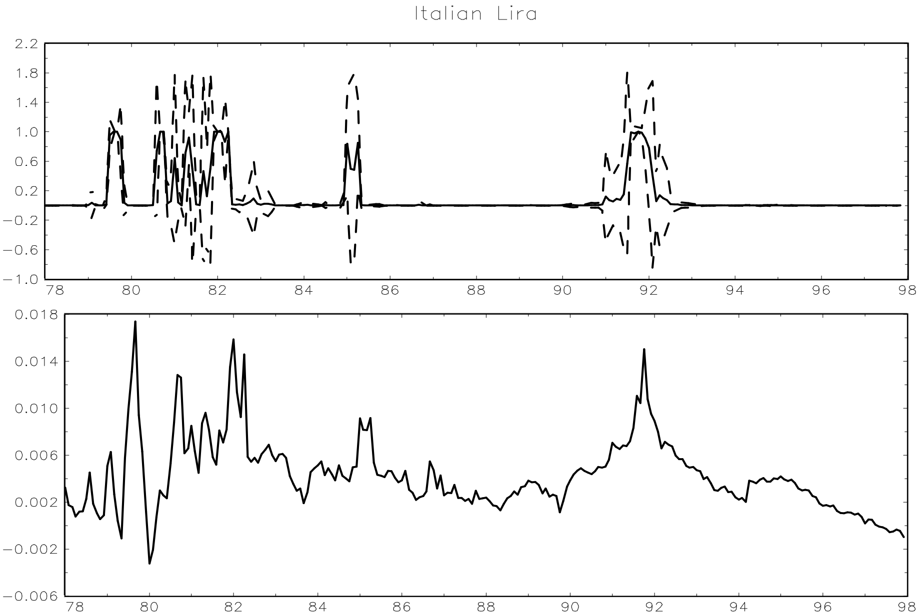


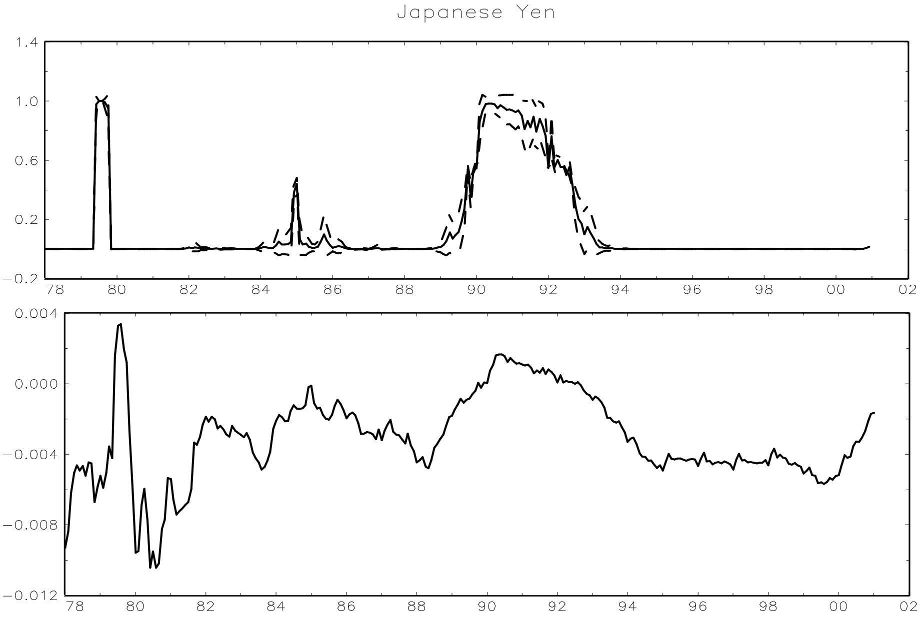
French Franc



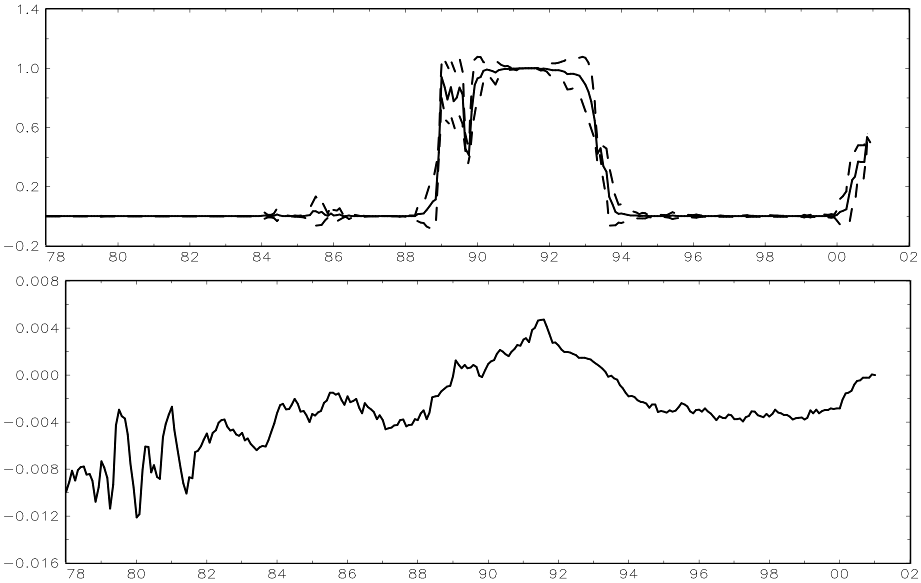
German mark



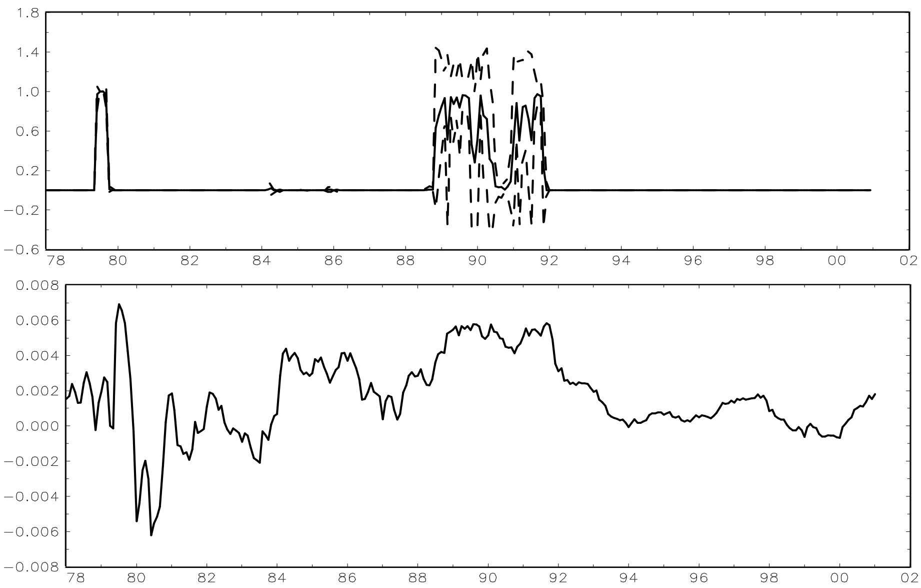

Italian Lira
2.2
1.8
1.4
1.0
0.6
0.2
−0.2
−0.6
−1.0
78 80 82 84 86 88 90 92 94 96 98
0.018
0.014
0.010
0.006
0.002
−0.002
−0.006
78 80 82 84 86 88 90 92 94 96 98


Japanese Yen
1.4
1.0
0.6
0.2
−0.2
78 80 82 84 86 88 90 92 94 96 98 00 02
0.004
0.000
−0.004
−0.008
−0.012
78 80 82 84 86 88 90 92 94 96 98 00 02

Swiss Franc



UK Pound



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