The Forward Unbiasedness Hypothesis and the Forward Premium: A Nonlinear Analysis

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September 2003

* This paper, which is part of the author's PhD thesis, has been prepared for the annual Money, Macro and Finance Conference in Cambridge in September 2003. I am grateful to Professor Lucio Sarno and participants at the finance seminar series at the Warwick Business School for helpful comments and feedback. Any errors and omissions are my own.

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Abstract

This paper considers nonlinear aspects of testing the mean-reversion of the forward premium. In contrast to standard linear methods, we consider a novel approach that allows for the joint testing of nonlinearity and nonstationarity. Within this approach, we employ nonlinear threshold autoregressive (TAR) unit root tests to investigate whether the 1 and 3 month forward premium series for seven industrial countries are mean-reverting. Overall, we are able to reject the null hypotheses of linearity and nonstationarity indicating nonlinear mean-reversion of the forward premium. Further, large deviations of the forward premium from its long-run equilibrium level are found to have faster speed of mean-reversion than small deviations.

JEL Classification: F30, F31, C32, C33.

Keywords: Forward Premium, Panel Unit Root Tests, Half-lives and Nonlinear Threshold Autoregressive Unit Root Tests.

1. Introduction

The relationship between spot and forward exchange rates has been studied extensively in the international finance literature. However, the literature has produced conflicting results related to a number of issues. One of these controversial issues concerns the stationarity of the forward premium and cointegration between forward and spot exchange rates (Kutan and Zhou, 2002). The stationarity of the forward premium is important for a number of reasons. First, if the forward premium is a realisation of a unit root process it makes the commonly employed market efficiency test of Fama (1984) inappropriate since it suffers from the spurious regression type critique (Granger and Newbold, 1974; Baillie, 1996). Second, since Hakkio and Rush (1989), it has become common that tests of the forward rate unbiasedness hypothesis (FRUH), which states that the forward exchange rate should be an unbiased predictor of the future spot rate, require that the future spot rate s_{t+1} and the forward rate f_t be cointegrated with a cointegrating vector of [1,-1][1]. It is also true that the FRUH requires the spot rate s_t and forward rate f_t to be cointegrated with cointegrating vector of [1,-1] or equivalently that the forward premium ($f_t - s_t$) be stationary (Engel, 1996; Zivot, 2000).

In a recent survey, Engel (1996) summarized the evidence on tests of cointegration between spot and forward exchange rates and stationarity of the forward premium and forecast error / excess return as follows: "Some have found (s_t-f_t) is I (0); some have found it is I (1); some have found it is fractionally integrated. Some have found s_{t+1} and f_t are cointegrated with cointegrating vector [1,-1]; some have found they are cointegrating but not with cointegrating vector [1,-1]; some have found they are not cointegrated. These conflicting results hold on tests for the same set of currencies," (pp. 141). Engel (1996) argues that these conflicting results may arise from different sampling periods, but more likely they result from different properties of the various tests employed.

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[1] Cointegration between the future spot rate s_{t+1} and the forward rate f_t with a cointegrating vector of [1,-1] implies that the foreign exchange forecast error / excess return (s_{t+1}-f_t) is stationary and vice versa (see definition of cointegration in Engle and Granger, 1987). This also applies to the case of spot and forward exchange rates and the forward premium.

In this paper we focus on the properties of some of the tests used to investigate the stationarity of the forward premium. We argue that the major problem with the standard unit root and cointegration tests that have been used so far to identify the order of integration of the forward premium is that they implicitly assume linear adjustment. However, there are good reasons why, if spot and forward rates co-move and are linked by a cointegrating relationship, adjustment toward equilibrium may be nonlinear. For example, transaction costs and market frictions are often cited as the major sources of nonlinearities. According to Dumas (1992) transaction costs create a band of inaction within which deviations from the forward premium long run equilibrium level are left uncorrected as they are not large enough to be profitable. It is only deviations that are outside the band that are arbitraged by market agents. In this framework, the forward premium follows a nonlinear process that is mean-reverting. Nonlinearities in foreign exchange markets are supported by evidence of nonlinearities in nominal and real exchange rates (Obstfeld and Taylor, 1997; Kilian and Taylor, 2001; Taylor *et al.*, 2001) and in the adjustment of the nominal exchange rate towards the long run equilibrium level suggested by monetary fundamentals (Meese and Rose, 1991; Taylor and Peel, 2000). However, there is no clear way of differentiating between nonlinearity and nonstationarity in the above literature. Indeed, these studies assume stationarity prior to fitting a nonlinear model and find evidence of fast nonlinear mean-reversion when these nonlinear models are used. However, if stationarity is not valid and the variable under study has a unit root, then tests of linearity versus a nonlinear threshold alternative will lead to incorrect inferences as these tests will have non-standard asymptotic distributions. Also, results found in the above studies regarding nonlinearity and mean-reversion must be interpreted cautiously as long as a proper test of unit root against a stationary nonlinear (threshold) alternative has not been performed (Bec *et al.*, 2002). A recent paper by Taylor (2001) has shown that standard unit root tests have low power in the presence of nonlinear adjustment, however.

To circumvent these problems, we use a novel approach whose tests and distribution theory have been recently developed by Caner and Hansen (2001) and which allows for the joint testing of nonlinearity and nonstationarity of the forward premium[2]. This approach

[2] Sekioua (2003a) uses this approach to model the deviation of the nominal exchange rate from the level predicted by monetary fundamentals such as the money supply and income. Using nonlinear

uses a two-regime symmetric threshold autoregressive (TAR) model with an autoregressive unit root which allows for an inner no-arbitrage band for small disequilibria and captures mean-reversion in response to shocks outside the no-arbitrage band. Within this model, we study Wald tests for nonlinear adjustment and Wald and *t*-tests for nonstationarity. We also allow for general autoregressive orders and do not artificially restrict coefficients across regimes (Basci and Caner, 2002).

In this paper, we first examine the univariate time series properties of the forward premium for seven industrial countries and for the recent floating exchange rate period. Unsurprisingly, we find little evidence of mean-reversion in the forward premium and this is explained by the low power of the univariate unit root tests to reject the unit root null hypothesis especially when using short data samples[3]. To overcome the power problems of the univariate tests we increase the number of observations and test for the non-stationarity of the forward premium in a panel framework[4]. However, despite the use of panel tests such as the Maddala and Wu (1999) and Hadri (2000) tests which are appropriately sized and have excellent power against close local alternatives, we were still unable to reject the null hypothesis of a unit root in the forward premium. Given nonlinear exchange rate adjustment, the dynamic relationships implicit in testing the stationarity of the forward premium series using these linear univariate and panel tests are misspecified and it should come as no surprise that we cannot detect any mean-reversion in the forward premium. Allowing for the possibility of nonlinear adjustment and using nonlinear TAR unit root and bootstrap

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unit root tests, the author uncovers evidence of nonlinear mean reversion of the deviation series. This evidence reinforces those of studies which detected nonlinearities in exchange rate adjustment. However, the results are more robust since one of the drawbacks of previous studies which was the failure to test for unit root was dealt with.

[3] The failure to establish mean reversion in the forward premium may be due to lack of power of conventional unit root tests. Taylor (1995) argued that unit root tests are biased downward in the sense they are more likely to reject stationarity when the underlying process driving the data is in fact stationary. This bias is due to the small size of available data samples.

[4] The use of panel data allows us far greater flexibility and this has also been proven quite satisfactory in improving the power of unit root tests. The additional cross sectional dimension in the panel leads to better power properties of the panel tests as compared to the lower power of the standard individual unit root test against near unit root alternatives especially when using short data samples.

likelihood tests, we first obtain strong evidence of nonlinearities in the forward premium. Second, using Wald and *t*-tests we were able to reject the unit root null hypothesis in at least one regime (the regimes being inside and outside the transaction cost band). Specifically, we find that the forward premium displays unit root behavior when it is within the transaction costs band and when there is a large shock it is mean-reverting. Finally, the computed half-lives of deviations indicate that the speed of adjustment outside the band is faster than than inside the transaction cost band. Finally, the evidence of nonlinear mean-reversion of the forward premium, and the forecast error / excess return, uncovered in this paper provides support for the forward rate unbiasedness hypothesis (FRUH) and the uncovered interest rate parity (UIP) hypothesis which requires a stationary forecast error. This evidence also indicates that the interest rate differential, which is linked to the forward premium through the covered interest rate parity (CIP) condition, is also a stationary TAR process.

This paper is structured as follows: Section 2 provides a description of the workhorse of our analysis, namely the forward rate unbiasedness hypothesis (FRUH). The empirical methodology is described in Section 3. Data and empirical results are provided in Section 4. The last section concludes.

2. Forward Rate Unbiasedness Hypothesis (FRUH) and the Forward Premium

The forward rate unbiasedness hypothesis (FRUH), which states that the forward exchange rate is an unbiased predictor of the future spot exchange rate, has been studied extensively in the empirical literature of international finance (Engel, 1996). This hypothesis was developed as a corollary to the efficient market hypothesis.

Assuming rational expectations we get:

$$E_t\,(s_{t+1}/I_t)=f_t \qquad (1)$$

where f_t is the log forward rate for a 1 or 3-month contract at time t, s_{t+1} is the corresponding log spot rate at time $t+1$, and E_t (.) is the expectations operator given information at time t, I_t. Given the assumption of rational expectations, we get:

$$s_{t+1}=E_t(s_{t+1})+u_{t+1} \qquad (2)$$

where u_{t+1}, is a zero mean, white noise process. Now, substituting (1) into (2) yields:

$$s_{t+1}=f_t+u_{t+1} \qquad (3)$$

Equation (3) delivers the FRUH. This hypothesis is tested empirically using the following regression:

$$s_{t+1}=\eta+\beta f_t+u_{t+1} \qquad (4)$$

where FRUH holds if $\eta=0, \beta=1$ and $E_t[u_{t+1}]=0$. Naka and Whitney (1995) and Hai *et al.* (1997) refer to testing $\eta=0, \beta=1$ as testing the forward rate unbiasedness hypothesis. Testing the orthogonality condition $E_t[u_{t+1}]=0$, conditional on not rejecting FRUH, is then referred to as testing market efficiency under rational expectations and risk neutrality (Zivot, 2000). Several researchers over the years have examined regression (4) with various improvements in econometric techniques and several early results are supportive of the unbiasedness hypothesis (Cornell, 1977; Frenkel, 1977). However, subsequent research on testing for unit root drew doubt as to the appropriateness of (4), as exchange rates display unit root behaviour. Concerns over spurious regressions prompted researchers to look for stationary alternatives to (4). This resulted in the following specification:

$$s_{t+1}-s_t=\eta'+\beta'(f_t-s_t)+\varepsilon_{t+1} \qquad (5)$$

Equation (5) shows that the forward premium (f_t-s_t) is an unbiased predictor of the future currency depreciation ($s_{t+1}-s_t$) if $\eta'=0, \beta'=1$ and $E_t[\varepsilon_{t+1}]=0$. If the spot rate is difference stationary then for regression (5) to be balanced and the FRUH hypothesis to hold, (f_t-s_t) must be stationary, assuming $\eta'=0, \beta'=1$ and $E_t[\varepsilon_{t+1}]=0$. This can be shown by applying the hypothesis that the risk premium ($rp_t=f_t-E_t[s_{t+1}]$) is null and that expectations are rational ($E_t[s_{t+1}]=s_{t+1}-u_{t+1}$). We can then write:

$$f_t-s_t=E_t[s_{t+1}]-s_t=s_{t+1}-s_t-u_{t+1} \qquad (6)$$

where u_{t+1} is the forecast error, which must be stationary under rational expectations. Hence, if ($s_{t+1}-s_t$) and u_{t+1} are stationary, the forward premium (f_t-s_t) must also be stationary. The latter condition implies that the spot and forward rates have to be cointegrated with a cointegrating vector of [1,-1].

Hakkio and Rush (1989) and Engel (1996) stressed that testing the unbiasedness hypothesis from (4) or (5) involves cointegration between either s_{t+1} and f_t or f_t and s_t with the same cointegrating vector of [1,-1]. It is clear from

$$s_{t+1} - f_t = \Delta s_{t+1} - (f_t - s_t) \tag{7}$$

that if the change in the spot exchange rate is stationary then cointegration between the future spot rate and the forward rate with a cointegrating vector of [1,-1] involves cointegration with the cointegrating vector [1,-1] between the spot and forward exchange rate. Essentially, stationarity of the forecast error / excess return implies stationarity of the forward premium and vice versa. In this paper we focus on the relationship between the forward and spot rate only, i.e. the forward premium, since this relationship has implications for regression tests based on (5) and the unbiasedness hypothesis.

Surveying the empirical evidence on the stationarity of the forward premium and cointegration between spot and forward exchange rates we find that the results are somewhat mixed. Clarida and Taylor (1997), Hai *et al.* (1997) and Barkoulas *et al.* (2003) support the stationarity of the forward premium. Others obtain evidence supportive of unit root behaviour of the forward premium (Crowder, 1994; Evans and Lewis, 1995; Horvath and Watson, 1995; Luintel and Paudyal, 1998). In general, the results depend on the testing procedure, the data frequency and time period. Further, the high persistence and heterogeneity of the forward premium reduce the power of unit root tests and distort the size of stationarity tests (Engel, 1996). These problems have led some researchers to consider non-standard models of unit root and cointegration testing. For example, Baillie and Bollerslev (1994) and Byers and Peel (1996) concur that the forward premium is fractionally integrated, whereas Baillie (1994) considers fractional cointegration between spot and forward exchange rates. Kutan and Zhou (2002), on the other hand, use rolling cointegration techniques and find that the relationship between spot and forward rate broke down in the 1980s. Although, they became cointegrated again in the 1990s, they no longer co-moved proportionally, however.

The main problem with previous studies of the stationarity of the forward premium is that they implicitly assume that exchange rate adjustment is linear. Nonlinearities in the relationship between spot and forward exchange rates can be due to a variable speed of adjustment towards a long-run equilibrium. This may arise because small deviations are not considered important by the market and the authorities, whereas for larger deviations, the pressure from the market to return the exchange rate near its equilibrium value becomes larger (Taylor and Allen, 1992; Taylor and Peel, 2000). The same reasoning applies in the presence of transaction costs (Dumas, 1992). Transaction costs keep arbitrage on hold until the forward premium between two currencies departs sufficiently above a threshold of inaction. Nevertheless, in spite of the evidence supportive of nonlinear exchange rate adjustment, there are only a few studies of FRUH which take nonlinearities into account. Coakley and Fuertes (2001 a, b) address the forward premium puzzle[5] in a nonlinear framework and argue that the insignificant coefficients from recent data spans can be explained by an unbalanced regression problems caused by asymmetries and nonlinearities in spot returns.

3. Empirical Methodology

3.1. The Maddala and Wu (1999) Panel Unit Root Test

Maddala and Wu (1999) proposed a panel unit root test that is based on the p-values of N independent tests of a hypothesis. This test, therefore, can be used with any kind of unit root or stationarity test. The significance levels p_i (i = 1, 2, 3, …, N) are independent uniform (0, 1) variables and $-2\ln p_i$ has a χ^2 distribution with two degrees of freedom. Using the additive property of the χ^2 variables, the following test statistic (with a χ^2 distribution and $2N$ degrees of freedom) can be derived:

$$MW = -2\sum_{i=1}^{N} \ln p_i \tag{8}$$

This test statistic can be used with any kind of unit root or stationarity test. Indeed, the test's null hypothesis is the same as that of the univariate test used to obtain the significance levels p_i. This is its biggest advantage when compared with other panel unit root data tests, such as those proposed by Im *et al.* (1997) and Pedroni (1998). Further this test does not require a balanced panel and is non-parametric.

[5] In a regression of the change in the spot exchange rate on the forward premium, the slope coefficient is not only different from one but it is also negative. This is the forward premium puzzle.

3.2. The Hadri (2000) Panel Stationarity Test

The Hadri panel stationarity test is an extension of the univariate KPSS test to the panel framework with N cross-sections. Following single time series stationary testing procedures proposed by Kwiatkowski *et al.* (1992), consider the following two models:

$$fp_{it} = r_{it} + \varepsilon_{it}\ ,\ i=1,\dots,T \tag{9}$$

or

$$fp_{it} = r_{it} + \beta_i t + \varepsilon_{it} \tag{10}$$

where r_{it} is a random walk defined through $r_{it} = r_{it-1} + u_{it}$ with fixed initial values r_{i0}. To test the null of stationarity, consider the following hypothesis:

$$H_o\!:\ \lambda = 0 \text{ against } H_a\!:\ \lambda > 0 \tag{11}$$

where $\lambda = \sigma_u^2 \,/\, \sigma_\varepsilon^2$ and $\sigma_u^2 = 0$ under the null. The panel can be presented as:

$$fp_i = X_i B_i + e_{it} \tag{12}$$

where $fp_i^{'} = [fp_{i1}...fp_{iT}]$, $e_i^{'} = [e_{i1}...e_{iT}]$ and X_i is a $T\times 1$ unit (1) vector. The Hadri LM test statistic is:

$$LM = \frac{1}{N}\sum_{i=1}^{N}\left(\frac{\frac{1}{T^2}\sum_{t=1}^{T} S_{it}^2}{\sigma_{\varepsilon,i}^2}\right) \tag{13}$$

where S_{it}^2 and $\sigma_{\varepsilon,i}^2$ are the partial sum of residuals and the variance from each individual series. Hadri (2000) suggests the following semi-parametric correction for serial correlation to be applied to each variance term in the panel:

$$\sigma_i^{*2}(x) = \gamma_0 + 2\sum_{s=1}^{T-1} \varkappa(x)\gamma_s \tag{14}$$

where $\sigma_i^{*2} = \gamma_0$, the bandwidth $x = s/l+1$, l is the lag truncation and $\gamma_s = \frac{1}{T}\sum_{t=s+1}^{T} e_{it}e_{it-s}$. A number of choices are available for the Kernel $\varkappa(x)$. Hadri has suggested that the Quadratic-Spectral (QS) kernel might be optimal, but for comparison Bartlett (BT) and Tukey-Hanning (TH) kernels are also included. The following finite sample correction to the Hadri LM test statistic is used:

$$Z_{\mu} = \frac{\sqrt{N}(LM - \xi_{\mu})}{\varsigma_{\mu}} \quad (15)$$

where $\xi_{\mu} = 1/6,\ \varsigma_{\mu} = \sqrt{1/45}$ when an intercept is included and $\xi_{\mu} = 1/15,\ \varsigma_{\mu} = \sqrt{11/6300}$ when an intercept and a trend are included.

3.3. The Caner and Hansen (2001) Threshold Autoregressive (TAR) Unit Root Test

The model suggested by Caner and Hansen (2001) is a threshold autoregressive (TAR) process of the form:

$$\Delta fp_t = \theta_1' x_{t-1} 1_{\{Z_{t-1} < \lambda\}} + \theta_2' x_{t-1} 1_{\{Z_{t-1} \geq \lambda\}} + e_t \quad (16)$$

t=1,…, T, where $x_{t-1} = (fp_{t-1} r_t' \Delta fp_{t-1} ... \Delta fp_{t-k})'$, $1_{\{.\}}$ is the indicator function, e_t is an *iid* error, $Z_t = fp_t - fp_{t-m-1}$ for some $m \geq 1$ and r_t is a vector of deterministic components including an intercept and possibly a linear time trend. The threshold λ is unknown and takes on values in the interval $\lambda \in \Lambda = [\lambda_1, \lambda_2]$ where λ_1 and λ_2 are picked so that $P(Z_t \leq \lambda_1) = \pi_1 > 0$ and $P(Z_t \leq \lambda_2) = \pi_2 < 1$ [6]. It is convenient to show the components of θ_1 and θ_2 as:

$$\theta_1 = \begin{pmatrix} \mu_1 \\ \beta_1 \\ \varrho_1 \end{pmatrix} \text{ and } \theta_2 = \begin{pmatrix} \mu_2 \\ \beta_2 \\ \varrho_2 \end{pmatrix} \quad (17)$$

The TAR model is estimated by least squares (LS):

$$\Delta fp_t = \hat{\theta}_1(\lambda)' x_{t-1} 1_{\{Z_{t-1} < \lambda\}} + \hat{\theta}_2(\lambda)' x_{t-1} 1_{\{Z_{t-1} \geq \lambda\}} + \hat{e}_t(\lambda) \quad (18)$$

Let

$$\hat{\sigma}^2(\lambda) = T^{-1} \sum_{1}^{T} \hat{e}_t(\lambda)^2 \quad (19)$$

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[6] Since only the magnitude of the change in the forward premium that matters and not the sign, we consider the absolute value of the change as the switching variable. We, therefore, retain a symmetric threshold $\lambda_1 = -\lambda_2 = \lambda$. This makes our TAR a two-regime symmetric model.

be the OLS estimate of σ^2 for fixed λ. The least squares estimate of the threshold λ is found by minimizing $\sigma^2(\lambda)$:

$$\hat{\lambda} = \underset{\lambda \in \Lambda}{\arg\min}\ \hat{\sigma}^2(\lambda) \tag{20}$$

In model (5), an important issue is whether there is a threshold effect. The threshold effect disappears under the hypothesis that:

$$H_0: \mu_1 = \mu_2, \varrho_1 = \varrho_2 \tag{21}$$

This hypothesis is tested using a standard Wald test statistic W_T. This statistic is written as:

$$W_T = T(\hat{\sigma}_0^2 / \hat{\sigma}^2 - 1) \tag{22}$$

where $\hat{\sigma}^2$ is the residual variance from (18), and $\hat{\sigma}_0^2$ is the residual variance from OLS estimation of the null linear model. Let (23) denote the Wald statistic of hypothesis (21) for fixed threshold.

$$W_T(\lambda) = T(\hat{\sigma}_0^2 / \hat{\sigma}^2(\lambda) - 1) \tag{23}$$

Then since $W_T(\lambda)$ is a decreasing function of $\hat{\sigma}^2(\lambda)$, we find that the Wald statistic is:

$$W_T = W_T(\hat{\lambda}) = \sup_{\lambda \in \Lambda} W_T(\lambda) \tag{24}$$

In model (18), the parameters ϱ_1 and ϱ_2 control the stationarity of the process fp_t. A leading case is when fp_t is a unit root process such that:

$$H_0: \varrho_1 = \varrho_2 = 0 \tag{25}$$

The standard test for (20) against the unrestricted alternative $\varrho_1 \neq 0$ or $\varrho_2 \neq 0$ is the Wald statistic:

$$R_{2T} = t_1^2 + t_2^2 \tag{26}$$

where t_1 and t_2 are t ratios for $\hat{\varrho}_1$ and $\hat{\varrho}_2$ from the OLS regression of the TAR model. While it is unclear how to form an optimal one-sided Wald test, Caner and Hansen (2001) recommend focusing on negative values of $\hat{\varrho}_1$ and $\hat{\varrho}_2$ to end up with a simple one-sided Wald test statistic:

$$R_{1T} = t_1^2 1_{\{\hat{\varrho}_1 < 0\}} + t_2^2 1_{\{\hat{\varrho}_2 < 0\}} \tag{27}$$

which is testing the unit root null hypothesis against the one-sided alternative $\varrho_1 < 0$ or $\varrho_2 < 0$. Generally, Caner and Hansen suggest examining the individual *t* statistics

t_1 and t_2 such that an insignificant t statistic provides evidence in favour of the presence of a unit root in the TAR process. While the distributions of R_{1T} and R_{2T} have asymptotic approximations, improved finite sample inference may be conducted using a bootstrap distribution. In this paper, we obtain the exact *p*-values of the R_{1T} and R_{2T} statistics using 10000 bootstrap simulations consistent with Caner and Hansen (2001).

4. Data and Empirical Results

4.1. Data

Monthly spot and forward exchange rates data for seven industrial countries, namely, Australia, Canada, France, Germany, Italy, Japan, Switzerland and the USA, against the UK pound sterling are used in the empirical analysis for this paper. The data is for the period spanning from January 1979 to December 1998 and is obtained from DataStream. We use these currencies because they are heavily utilized in earlier studies. By using monthly data and 1- and 3-month forward contracts, overlapping data effects are avoided. The empirical analysis in this paper is performed using version 3.2. of the GAUSS programming language.

4.2. Univariate and Panel Unit Root Tests

Mean-reversion of the forward premium series would provide evidence of cointegration between spot and forward exchange rates. Table 1 presents the values of the Augmented Dickey Fuller (ADF) test performed on the 1- and 3-month forward premium series. The ADF test on the log levels of the 1-month forward premium series shows that the unit-root null hypothesis can only be rejected for Switzerland and Japan at the 5% significance level and for Canada at the 10% level. Results of the ADF test on the 3-month forward premium series are quantitatively similar to those on the 1-month forward premium series. However, the inability to reject the unit root null hypothesis using the ADF test should not be regarded as prima facie evidence against the stationarity of the forward premium. Indeed, results of the ADF test have to be interpreted with care since a major criticism of the ADF unit root testing procedure is that it cannot distinguish between unit root and near unit root processes due to its low power in the presence of such a persistent variable as the forward premium (Engel, 1996).

Given the conflicting findings of the univariate ADF unit root test, we now apply panel unit root and stationarity tests to systems of forward premium series. By exploiting cross-equation dependencies and increasing the span of the data by jointly testing for a unit root and stationarity across a number of series, panel tests are likely to lead to substantial gains in test power. In table 1 we implement correlation analysis to identify cross-sectional dependencies among the system variables. As reported in table 1, there are substantial correlations among the panel forward premium series, with the notable exception of the Swiss / USA and Swiss / Italy forward premium series. Overall, these results suggest that implementation of panel unit root and stationarity tests would lead to substantial power gains.

In table 2 reports the individual ADF *t*-statistics and their corresponding *p*-values for each one of the forward premium series. These values are used to compute the Maddala and Wu (1999) test statistic. The results of the Maddala and Wu (1999) statistic show that there is support for the rejection of the unit root null hypothesis as the Maddala and Wu statistic is greater than the critical value at the 1% significance level. However, it is possible that this test is influenced by outliers in the panels, and that, in the current application testing is of limited value due to heterogeneity of the data of different countries, some being stationary and others integrated of order 1. To illustrate this point, reconsider the Maddala and Wu (1999) test statistic without the forward premium series whose ADF statistics are significant at the 5% level. If we removed these series and considered a panel of 4 forward premium series the calculated Maddala and Wu (1999) test statistic falls from 31.2838 to 10.4849 for 1-month forward premium series, and now not significant even at the 10% significance level. We reach a similar conclusion for the 3-month forward premium.

Given the mixed results obtained with the Maddala and Wu (1999) test statistic, we now turn our attention to testing the stationarity of the forward premium series using a test that does not rely upon the *t*-statistics of the ADF test. For this panel test, we use the non-parametric correction to the stationarity test due to Hadri (2000) to take account of heterogeneous serial dependence across our panel that is made up of 7 forward premium series. The test considers the null of stationarity and for a sample with more than 50 observations Hadri (2000) shows that this test has high power against close local alternatives

(Hunter and Simpson, 2001). The results of the Hadri (2000) are reported in table 3. Ordering the tests by their specification, the test statistics reject the null hypothesis of stationarity for the 1-month forward exchange rate premium. As for the 3-month forward premium the evidence is mixed. The test statistics corrected for serial correlation accept the null of stationarity whereas the null hypothesis is rejected when there is no correction for serial correlation. Therefore, despite the fact that this test appears suited to test whether the forward premium series are stationary the results are inconclusive.

4.3. Threshold Autoregressive (TAR) Bootstrap Tests

The major problem with the tests that we have used so far is that they implicitly assume linear adjustment. Nonlinearities in the relationship between spot and forward exchange rates can be due to a variable speed of adjustment towards a long-run equilibrium. This may arise because small deviations are not considered important by the market and the authorities, whereas for larger deviations, the pressure from the market to return the exchange rate near its equilibrium value becomes larger (Taylor and Allen, 1992; Taylor and Peel, 2000). Nonlinearities can also arise as a consequence of transaction costs and market frictions (Dumas, 1992).

To explore the forward premium for potential frictions captured by nonlinearities; we employ the nonlinear threshold autoregressive (TAR) framework of Caner and Hansen (2001). The model used is a two-regime symmetric TAR model with an autoregressive root that is local-to-unity. The first step in our analysis involves testing for linearity and the appropriate test statistic for this is a Wald test. In tables 4 to 7 we report the Wald test, 1% bootstrap critical values and bootstrap *p*-values for threshold variables of the form $Z_t = fp_t - fp_{t-m-1}$ for delay parameters m from 1 to 12. However, the optimal delay parameter is chosen so that it minimises the residual variance of the TAR model of each forward premium series. The delay parameters for the 1 - (3-) month forward premium series are 4 (7) for the USA, 1 (8) for Switzerland, 7 (4) for Japan, 2 (1) for Canada, 5 (5) for France, 11 (8) for Germany and 6 (7) for Italy. From tables 4 and 6, it is clear that each statistic corresponding to the optimal delay parameter is highly significant and easily rejects the null hypothesis of linearity in favour of the nonlinear threshold model.

In tables 8 and 9 we report the threshold unit root t_1 and t_2 statistics for each delay parameter m from 1 to 12, and their p-values obtained using 10000 bootstrap simulations. However, the t-statistics of interest are those that correspond to the delay parameters that minimise the residual variances. These delay parameters are the same as those used to test for linearity. For these delay parameters the bootstrap p-values for t_1 for the 1- (3-) month forward premium series are 0.0979 (0.0431) for the UK, 0.3670 (0.0849) for Switzerland, 0.0391 (0.1880) for Japan, 0.5500 (0.9530) for Canada, 0.4500 (0.0331) for France, 0.5610 (0.1120) for Germany and 0.0057 (0.0894) for Italy. The significant t_1 statistics for 1-month forward premium series for the UK, Japan and Italy indicate that we are able to reject the unit root null hypothesis in favour of $\rho_1 < 0$ in the first regime. For the 3-month forward premium we are able to reject the null hypothesis for the UK, Switzerland, France and Italy. The bootstrap p-values for the t_2 statistic for the 1- (3-) month forward premium series are 0.3400 (0.3200) for the UK, 0.3550 (0.1540) for Switzerland, 0.2180 (0.0047) for Japan, 0.0908 (0.0112) for Canada, 0.6060 (0.5550) for France, 0.1380 (0.2440) for Germany and 0.8260 (0.6550) for Italy. In this case, p-values indicate that we are able to reject the unit root null hypothesis in the second regime for 1- and 3-month forward premium series for Canada and 1-month forward premium for Japan only.

However, we are able to settle the unit root question using the R_{1T} and R_{2T} statistics which are reported in tables 9 and 11. The one-sided Wald test R_{1T}, which tests unit root against a two-regime stationary nonlinear model, is rejected in 2 out of 7 cases for the 1-month forward premium. For the 3-month forward premium, the unit root null hypothesis is rejected in 5 out of 7 cases. However, on closer inspection we find that the results for the one-sided unit root test are due to the fact that the first regime is mean-reverting whereas the second regime has a unit root and this regime is dominant. Taking this reasoning into consideration and following Basci and Caner (2002) we reject the unit root null in favour of a stationary process for 4 out of 7 1-month forward premium series. For the 3-month forward premium series, we reject the unit root null for 6 out of the 7 cases considered. This is an important result since it indicates that once allowance is made for nonlinearities we are able to reject the unit root null hypothesis for the forward premium. This would explain why we were unable to reject the unit root null hypothesis using linear univariate and panel unit root testing techniques.

For the chosen delay parameter we report in tables 13 and 14 the least squares estimates of the TAR models for each forward premium series. For the US 3-month forward premium, in particular, the point estimate of the threshold is – 0.0033. This value indicates that the TAR splits the regression function depending on whether the change in the forward premium lies above the threshold of –0.0033 in absolute terms. If the change in the forward premium in the past 7 months is above the threshold in absolute terms we switch to the second regime. The first regime (outside the band) behaves as a stationary process and has 18.9% of observations, whereas the second regime (inside the band) is a random walk with a drift and has 81.1% of observations[7]. Overall, the US 3-month forward premium is shown to be globally stationary since we are able to reject the unit root null for the R_{1T} statistic with a 5.78% bootstrap *p*-value.

4.4. Computed Half-lives of Deviations[8]

In table 14 we report the half-lives of deviations from the relationship between spot and forward exchange rates. The half-lives are defined as the number of months it takes for deviations to subside permanently below 50% in response to a unit shock in the level of the series and are computed because they essentially provide a measure of the degree of mean-reversion. In the first two columns we report the half-lives estimated using the linear autoregressive model and for the two maturity periods i.e. 1 and 3 months. It is clear from these two columns that the speed of mean-reversion is very slow. The estimates of the half-lives vary from a low of 4.26 months to a high of 15.77 months. In the remaining columns we report the half-lives estimated for the nonlinear TAR models. The half-lives are computed for the first regime (outside the band) and the second regime (inside the band). The values of the half-lives indicate that deviations outside the band of inaction are

[7] Given that most observations lie within the unit root regime, this would explain why standard univariate and panel unit root tests fail to reject the null hypothesis.

[8] Although it would have been useful to compare our results with other studies, we could not find any previous research which computed the half-life of deviations from the relationship between forward and spot exchange rates using either linear or nonlinear models. Therefore, out results represent an important contribution.

corrected or die out very quickly and the speed of adjustment is as low as 0.80 months. The estimates for the half-lives outside the band vary from 0.80 months to 14 months and this is much faster than the values we obtained using the linear model. As for the inner regime, adjustment is extremely slow. However, this is not surprising since the inner regime behaves essentially as a unit root process and its root is insignificantly different from zero in most cases. In general, it is only when the forward premium is in the neighbourhood of its long run equilibrium level do the nonlinear TAR models yield very slow speeds of adjustment. For large deviations, arbitrage kicks in and adjustment is almost instantaneous. Indeed, these half-lives are very fast when compared to the kind of estimates found in linear exchange rate literature (for example the PPP literature).

5. Conclusion

This paper addresses the issue of conflicting results associated with studies trying to detect mean-reversion in the forward premium. It is argued that the negative results are due to nonlinearities in the adjustment towards the long run. These stem from market frictions such as transaction costs. However, unlike previous studies, which tested for nonlinearities only, we propose an approach that allows the joint testing of nonlinearity and nonstationarity.

Using univariate threshold autoregressive (TAR) unit root and bootstrap tests, which allow for the possibility of nonlinear adjustment, we find evidence to support the hypothesis that the forward premium is a stationary TAR process. Specifically, the forward premium follows a nonlinear process and is found to display unit root behaviour when it is inside the transaction cost band whereas outside this band the forward premium is mean-reverting. The speed of adjustment outside the band is much faster than inside the band. These results indicate that once we allow for the possibility of nonlinear adjustment we are able to detect rapid mean-reversion of the forward premium.

Overall, there are two significant conclusions that we can draw from the empirical analysis in this paper. First, since the forward premium has been found to be nonlinearly mean-reverting a corollary of this is that the forward market forecast error / excess return is also a nonlinear mean-reverting process. This provides evidence in favour of the forward rate unbiasedness hypothesis (FRUH) and also confirms the validity of the uncovered interest rate parity (UIP) hypothesis which requires a stationary excess return. Second, the domestic and foreign nominal interest rate differentials are stationary TAR processes, given the covered interest rate parity (CIP) condition.

6. References

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7. Empirical Results

Table 1 Descriptive statistics for forward exchange rate premium series.

	USA	Switzerland	Japan	Canada	France	Germany	Italy	USA	Switzerland	Japan	Canada	France	Germany	Italy
k				1							3			
N	240	240	240	240	240	240	240	240	240	240	240	240	240	240
Mean	0.0019	0.0049	0.0048	0.0009	0.0004	0.0035	-0.0023	0.0053	0.0137	0.0137	0.0024	0.0009	0.0099	-0.0071
Maximum	0.0081	0.0131	0.0112	0.0056	0.0054	0.0094	0.0048	0.0181	0.0342	0.0305	0.0143	0.0145	0.0248	0.0105
Minimum	-0.0058	0.0002	0.0016	-0.0078	-0.0198	-0.0020	-0.0183	-0.0141	0.0009	0.0029	-0.0196	-0.0371	-0.0055	-0.0505
S.D.	0.0023	0.0025	0.0017	0.0021	0.0038	0.0024	0.0036	0.0066	0.0071	0.0045	0.0056	0.0100	0.0068	0.0101
Skewness	-0.3261	0.4539	0.3909	-1.0579	-2.0687	-0.2797	-1.2935	-0.2511	0.4744	0.3287	-1.0850	-1.5275	-0.3298	-1.0321
Kurtosis	3.4987	3.2604	3.0482	5.6020	9.1920	2.3782	5.8170	3.1002	3.2621	2.9292	5.3686	5.5449	2.3955	4.5543
Jarque-Bera	6.7403**	8.9180**	6.1345**	112.4640*	554.5894*	6.9957**	146.2864*	2.6216	9.6899*	4.3731	103.1894*	158.0938*	8.0058**	66.7686*
	(0.0344)	(0.0116)	(0.0465)	(0.0000)	(0.0000)	(0.0303)	(0.0000)	(0.2696)	(0.0079)	(0.1123)	(0.0000)	(0.0000)	(0.0183)	(0.0000)
ADF: Test	-1.8544	-2.8926**	-3.4584**	-2.7737***	-2.3022	-2.2923	-1.5722	-1.7273	-2.6807***	-3.6037*	-2.5487	-2.0266	-2.2584	-1.6476
	(0.3535)	(0.0478)	(0.0100)	(0.0637)	(0.1722)	(0.1754)	(0.4952)	(0.4161)	(0.0789)	(0.0064)	(0.1054)	(0.2753)	(0.1866)	(0.4565)
	-0.0588	-0.0758	-0.1500	-0.1080	-0.1200	-0.0543	-0.0653	-0.0479	-0.0622	-0.1430	-0.0973	-0.0752	-0.0430	-0.0479
ϱ_1	(0.0317)	(0.0262)	(0.0432)	(0.0389)	(0.0481)	(0.0237)	(0.0415)	(0.0277)	(0.0232)	(0.0397)	(0.0382)	(0.0371)	(0.0190)	(0.0291)
Half-life	11.4381	8.7933	4.2650	6.0649	5.4223	12.4153	10.2643	14.1213	10.7936	4.4917	6.7713	8.8663	15.7706	14.1213
Correlations														
USA	1	**0.0446**	0.2204	0.7354	0.4709	0.2634	0.4861	1	**0.0214**	0.2575	0.7475	0.4957	0.2394	0.5467
Switzerland	**0.0446**	1	0.5416	0.1818	0.2110	0.8657	**0.0738**	**0.0214**	1	0.5121	0.2034	0.2482	0.8804	**0.0940**
Japan	0.2204	0.5416	1	0.1895	0.3555	0.5889	0.3585	0.2575	0.5121	1	0.2461	0.4528	0.5813	0.4104
Canada	0.7354	0.1818	0.1895	1	0.6450	0.3193	0.5676	0.7475	0.2034	0.2461	1	0.6386	0.3121	0.5668
France	0.4709	0.2110	0.3555	0.6450	1	0.4099	0.8406	0.4957	0.2482	0.4528	0.6386	1	0.3956	0.8475
Germany	0.2634	0.8657	0.5889	0.3193	0.4099	1	0.3091	0.2394	0.8804	0.5813	0.3121	0.3956	1	0.2784
Italy	0.4861	**0.0738**	0.3585	0.5676	0.8406	0.3091	1	0.5467	**0.0940**	0.4104	0.5668	0.8475	0.2784	1

Note: ϱ_1 is the root of the autoregressive (AR) model. Critical values for the ADF test with an intercept are -3.4591 (1%), -2.8740 (5%) and -2.5735 (10%). *, ** and *** indicate significance at 1%, 5% and 10%.

Table 2 Maddala and Wu (1999) unit root test based on the *p*-values of the ADF unit root test.

Level	ADF *t*-statistic	*p*-value	ADF *t*-statistic	*p*-value	ADF *t*-statistic	*p*-value	ADF *t*-statistic	*p*-value
k		1				3		
USA	-1.8544	0.3535	-1.8544	0.3535	-1.7273	0.4161	-1.7273	0.4161
Switzerland	-2.8926**	0.0478	-	-	-2.6807***	0.0789	-	-
Japan	-3.4584**	0.0100	-	-	-3.6037*	0.0064	-	-
Canada	-2.7737***	0.0637	-	-	-2.5487	0.1054	-2.5487	0.1054
France	-2.3022	0.1722	-2.3022	0.1722	-2.0266	0.2753	-2.0266	0.2753
Germany	-2.2923	0.1754	-2.2923	0.1754	-2.2584	0.1866	-2.2584	0.1866
Italy	-1.5722	0.4952	-1.5722	0.4952	-1.6476	0.4565	-1.6476	0.4565
Maddala and Wu test statistic		31.2838*		10.4849		28.9414**		13.7593
		(0.0051)		(0.2326)		(0.0106)		(0.1843)
Degrees of Freedom		14		8		14		10

Note: Critical values for the ADF test with an intercept are -3.4591 (1%), -2.8740 (5%) and -2.5735 (10%). The asymptotic distribution of the Maddala and Wu (MW) test is $\chi^2(2N)$. *, ** and *** indicate significance at 1%, 5% and 10%.

Critical Values

d.f	0.1%	1%	2.5%	5%	10%
8	26.125	20.090	17.535	15.507	13.362
10	29.588	23.209	20.483	18.307	15.987
14	36.123	29.141	26.119	23.685	21.064

Table 3 Hadri (2000) panel stationarity test for heterogeneous panels.

	Test statistic, *k*=1	Test statistic, *k*=3
Intercept: LM (1)	2.6260*	0.8350
	(0.0043)	(0.2019)
Intercept: LM (2)	76.6981*	81.2337*
	(0.0000)	(0.0000)
Trend and Intercept: LM (3)	1.6883**	-0.3070
	(0.0457)	(0.6206)
Trend and Intercept: LM (4)	69.6017*	75.6749*
	(0.0000)	(0.0000)

Note: LM (1) denotes the value of the one-sided test LM test, testing for stationarity around an intercept with serially corrected errors. LM (2) denotes the value of the one-sided test LM test, testing for stationarity around an intercept. LM (3) denotes the value of the one-sided test LM test, testing for stationarity around an intercept and a time trend with serially corrected errors. LM (4) denotes the value of the one-sided test LM test, testing for stationarity around an intercept and a time trend. *, ** and *** indicate rejection of the stationarity null hypothesis at 1%, 5% and 10%.

Table 4 Threshold tests with fixed delay parameter, *m*. 1-month forward exchange rate premium series.

k=1	USA		Switzerland		Japan		Canada		France		Germany		Italy	
Delay Order	Wald Statistic	Statistic *p*-value	Wald Statistic	Statistic *p*-value	Wald Statistic	Statistic *p*-value	Wald Statistic	Statistic *p*-value	Wald Statistic	Statistic *p*-value	Wald Statistic	Statistic *p*-value	Wald Statistic	Statistic *p*-value
1	93.6000	0.0000*	63.7000	0.0001*	28.5000	0.2040	60.6000	0.0003*	112.0000	0.0000*	68.9000	0.0003*	65.0000	0.0040*
2	74.8000	0.0000*	42.2000	0.0061*	29.3000	0.1710	78.9000	0.0000*	86.7000	0.0005*	46.8000	0.0113**	66.0000	0.0021*
3	69.8000	0.0003*	40.5000	0.0135**	39.5000	0.0235**	68.7000	0.0001*	101.0000	0.0002*	56.4000	0.0022*	71.5000	0.0018*
4	94.1000	0.0000*	49.3000	0.0012*	29.3000	0.1710	54.0000	0.0017*	163.0000	0.0000*	38.9000	0.0389**	76.6000	0.0009*
5	89.4000	0.0000*	35.5000	0.0421**	37.1000	0.0335**	52.9000	0.0017*	276.0000	0.0000*	64.8000	0.0007*	68.2000	0.0020*
6	65.4000	0.0003*	35.5000	0.0435**	24.5000	0.3830	53.0000	0.0012*	189.0000	0.0000*	41.2000	0.0247**	136.0000	0.0000*
7	90.5000	0.0000*	26.0000	0.2850	49.4000	0.0027*	47.4000	0.0041*	66.3000	0.0026*	60.9000	0.0010*	93.4000	0.0003*
8	67.4000	0.0005*	33.1000	0.0658***	42.7000	0.0112**	53.4000	0.0012*	128.0000	0.0000*	45.1000	0.0130**	92.2000	0.0000*
9	93.2000	0.0000*	33.7000	0.0583***	42.9000	0.0109**	57.2000	0.0002*	108.0000	0.0000*	50.1000	0.0061*	44.5000	0.0346**
10	67.4000	0.0005*	30.5000	0.1160	48.9000	0.0032*	36.3000	0.0462**	129.0000	0.0001*	75.2000	0.0001*	83.3000	0.0010*
11	67.0000	0.0006*	31.9000	0.0842***	46.7000	0.0036*	27.3000	0.2600	91.6000	0.0000*	76.7000	0.0002*	67.6000	0.0029*
12	58.6000	0.0016*	36.0000	0.0339**	48.0000	0.0029*	33.6000	0.0852***	125.0000	0.0000*	64.3000	0.0007*	115.0000	0.0000*

Table 5 1% Bootstrap critical values

	USA	Switzerland	Japan	Canada	France	Germany	Italy
m	1% C.V.	1% C.V.	1% C.V.	1% C.V.	1% C.V.	1% C.V.	1% C.V.
1	48.8000	41.8000	43.5000	44.8000	58.1000	47.7000	55.9000
2	47.7000	40.5000	42.8000	44.4000	56.1000	47.5000	53.4000
3	48.0000	41.7000	43.3000	44.9000	56.6000	46.6000	53.9000
4	46.4000	41.2000	43.0000	44.0000	55.3000	46.4000	52.8000
5	46.5000	41.6000	42.2000	44.4000	57.7000	47.4000	54.7000
6	46.7000	41.9000	43.0000	44.3000	55.4000	46.7000	54.7000
7	47.2000	41.8000	42.8000	43.4000	55.9000	46.7000	55.4000
8	47.4000	41.7000	43.2000	43.9000	54.2000	46.4000	56.0000
9	46.7000	41.7000	43.4000	44.3000	55.6000	46.8000	54.6000
10	48.0000	41.1000	42.8000	44.6000	55.6000	46.7000	55.7000
11	47.1000	41.0000	43.0000	43.5000	54.3000	47.3000	55.9000
12	46.2000	41.0000	42.0000	44.0000	54.4000	46.5000	55.1000

Note: Bootstrap critical values calculated from 10000 replications.

Table 6 Threshold tests with fixed delay parameter, *m*. 3-month forward exchange rate premium series.

k=3	USA		Switzerland		Japan		Canada		France		Germany		Italy	
Delay Order	Wald Statistic	Statistic *p*-value	Wald Statistic	Statistic *p*-value	Wald Statistic	Statistic *p*-value	Wald Statistic	Statistic *p*-value	Wald Statistic	Statistic *p*-value	Wald Statistic	Statistic *p*-value	Wald Statistic	Statistic *p*-value
1	80.3000	0.0002*	35.5000	0.0545***	39.6000	0.0389**	47.5000	0.0058*	121.0000	0.0000*	79.1000	0.0001*	57.0000	0.0066*
2	62.9000	0.0008*	39.0000	0.0266**	50.2000	0.0075*	45.1000	0.0086*	50.2000	0.0075*	37.3000	0.0462**	59.8000	0.0039*
3	56.0000	0.0023*	65.1000	0.0000*	49.1000	0.0093*	35.0000	0.0603***	132.0000	0.0000*	37.8000	0.0396**	105.0000	0.0000*
4	96.0000	0.0000*	49.5000	0.0026*	62.2000	0.0016*	36.6000	0.0436**	129.0000	0.0000*	32.3000	0.1100	92.4000	0.0000*
5	76.3000	0.0002*	41.1000	0.0132**	45.1000	0.0149**	33.3000	0.0825***	152.0000	0.0000*	43.0000	0.0167**	80.6000	0.0000*
6	85.8000	0.0000*	42.7000	0.0095*	55.2000	0.0027*	32.0000	0.1090	147.0000	0.0000*	49.3000	0.0053*	106.0000	0.0000*
7	119.0000	0.0000*	74.9000	0.0000*	41.5000	0.0261**	34.5000	0.0655***	120.0000	0.0000*	76.4000	0.0000*	111.0000	0.0000*
8	82.6000	0.0000*	97.4000	0.0000*	33.9000	0.0980***	38.1000	0.0334**	118.0000	0.0000*	101.0000	0.0000*	104.0000	0.0000*
9	96.9000	0.0000*	59.0000	0.0004*	41.5000	0.0276**	36.7000	0.0399**	112.0000	0.0000*	82.5000	0.0000*	98.5000	0.0000*
10	66.0000	0.0006*	56.8000	0.0008*	36.1000	0.0639***	30.0000	0.1530	77.7000	0.0002*	78.9000	0.0001*	104.0000	0.0000*
11	67.4000	0.0003*	37.6000	0.0308**	44.8000	0.0148**	34.0000	0.0723***	96.5000	0.0000*	77.8000	0.0000*	72.8000	0.0002*
12	39.0000	0.0343**	43.9000	0.0084*	25.9000	0.3150	34.8000	0.0596***	103.0000	0.0000*	83.3000	0.0000*	98.6000	0.0000*

Table 7 1% Bootstrap critical values

	USA	Switzerland	Japan	Canada	France	Germany	Italy
m	1% C.V.	1% C.V.	1% C.V.	1% C.V.	1% C.V.	1% C.V.	1% C.V.
1	47.7000	43.1000	47.6000	44.3000	49.3000	46.3000	53.7000
2	47.7000	43.2000	48.5000	44.3000	48.7000	46.3000	51.5000
3	47.5000	42.8000	48.5000	43.4000	48.7000	45.8000	50.9000
4	46.8000	43.1000	48.2000	44.0000	49.6000	45.7000	51.1000
5	46.3000	42.9000	47.7000	44.3000	49.7000	46.0000	50.6000
6	46.5000	42.5000	48.8000	44.5000	49.0000	45.6000	51.1000
7	46.9000	42.9000	47.3000	43.7000	49.0000	45.9000	50.3000
8	45.8000	43.5000	47.9000	43.9000	49.1000	45.4000	49.6000
9	45.9000	43.5000	47.4000	43.6000	49.4000	45.0000	50.3000
10	46.6000	43.0000	47.8000	43.3000	48.6000	45.2000	50.4000
11	45.4000	42.6000	47.4000	43.9000	48.6000	45.1000	50.0000
12	46.4000	43.3000	46.3000	43.9000	49.3000	45.1000	51.0000

Note: Bootstrap critical values calculated from 10000 replications.

Table 8 Threshold autoregressive (TAR) unit root test. 1-month forward exchange rate premium series.

k=1	USA		Switzerland		Japan		Canada		France		Germany		Italy	
Delay Order	t_1 - statistic (Boot p-value)	t_2 - statistic (Boot p-value)	t_1 - statistic (Boot p-value)	t_2 - statistic (Boot p-value)	t_1 - statistic (Boot p-value)	t_2 - statistic (Boot p-value)	t_1 - statistic (Boot p-value)	t_2 - statistic (Boot p-value)	t_1 - statistic (Boot p-value)	t_2 - statistic (Boot p-value)	t_1 - statistic (Boot p-value)	t_2 - statistic (Boot p-value)	t_1 - statistic (Boot p-value)	t_2 - statistic (Boot p-value)
1	-0.1740	0.7950	1.4200	1.4900	2.4600	3.2800**	1.3200	2.3800	3.5600**	-0.1670	1.5300	1.5000	1.1600	0.5610
	(0.8350)	(0.6000)	(0.3670)	(0.3550)	(0.1020)	(0.0244)	(0.4100)	(0.1150)	(0.0206)	(0.8460)	(0.3440)	(0.3510)	(0.4480)	(0.6670)
2	1.8000	1.5500	1.9200	2.2600	3.0600**	1.9600	0.8900	2.5400***	1.0100	0.6540	0.9480	2.6300***	2.1400	1.1300
	(0.2530)	(0.3280)	(0.2230)	(0.1390)	(0.0365)	(0.2180)	(0.5500)	(0.0908)	(0.5070)	(0.6320)	(0.5410)	(0.0799)	(0.1700)	(0.4780)
3	1.6600	2.2800	2.0800	1.3700	2.9000***	1.9500	0.0545	2.7900***	2.9900***	0.7080	1.5300	2.5900***	0.6710	0.0094
	(0.3010)	(0.1410)	(0.1710)	(0.3950)	(0.0527)	(0.2170)	(0.7910)	(0.0602)	(0.0546)	(0.6190)	(0.3470)	(0.0856)	(0.6110)	(0.8060)
4	2.5300***	1.5200	2.2100	1.8600	3.4100**	1.6100	0.3880	2.7900***	7.7300*	0.8350	-0.6070	2.5300***	-0.6280	1.7200
	(0.0979)	(0.3400)	(0.1470)	(0.2440)	(0.0191)	(0.3220)	(0.7110)	(0.0616)	(0.0000)	(0.5870)	(0.9050)	(0.0957)	(0.9040)	(0.2880)
5	1.8400	1.1800	2.6000***	2.2100	2.5300***	3.3400**	2.2800	2.3500	1.1900	0.7610	0.4420	2.5900***	-0.4510	1.1700
	(0.2470)	(0.4560)	(0.0754)	(0.1540)	(0.0996)	(0.0215)	(0.1380)	(0.1290)	(0.4500)	(0.6060)	(0.7030)	(0.0837)	(0.8790)	(0.4710)
6	0.1960	0.7230	2.2400	1.7400	2.1700	3.0100**	4.3800*	1.9900	2.9200**	2.4400	-0.5340	2.3100	4.5800*	-0.0557
	(0.7570)	(0.6060)	(0.1400)	(0.2760)	(0.1760)	(0.0458)	(0.0021)	(0.2100)	(0.0588)	(0.1250)	(0.8960)	(0.1330)	(0.0057)	(0.8260)
7	2.0200	0.6150	0.5580	2.8000***	3.1100**	2.0100	2.8600**	2.2300	3.5900**	-0.5040	0.0559	2.3900	2.5300***	0.2620
	(0.2040)	(0.6390)	(0.6610)	(0.0565)	(0.0391)	(0.2180)	(0.0587)	(0.1560)	(0.0235)	(0.8970)	(0.7990)	(0.1220)	(0.0991)	(0.7520)
8	2.6200***	1.3500	2.4900***	2.5000**	3.2200**	2.0500	4.6500*	1.9200	7.7800*	-1.1500	1.6900	2.1200	0.3600	1.1700
	(0.0923)	(0.4070)	(0.0963)	(0.0952)	(0.0338)	(0.1990)	(0.0011)	(0.2340)	(0.0000)	(0.9550)	(0.3100)	(0.1770)	(0.7020)	(0.4770)
9	3.8100**	1.1700	1.7900	2.9400**	2.0700	3.2200**	4.0100*	2.5600***	8.1300*	-0.2060	1.5500	1.9500	2.8200***	1.4000
	(0.0115)	(0.4710)	(0.2560)	(0.0442)	(0.2070)	(0.0316)	(0.0065)	(0.0955)	(0.0000)	(0.8540)	(0.3540)	(0.2170)	(0.0703)	(0.3960)
10	2.6100***	1.0900	1.6000	3.9100*	2.2500	2.3600	3.5500**	2.5100	4.0200**	0.3580	3.5300**	1.6300	1.3600	0.0862
	(0.0928)	(0.4860)	(0.3290)	(0.0054)	(0.1600)	(0.1390)	(0.0182)	(0.1070)	(0.0122)	(0.7270)	(0.0206)	(0.3060)	(0.3780)	(0.7960)
11	1.8900	1.2900	2.9300**	1.0200	1.2400	2.9000***	3.0400**	1.8100	4.9200*	-0.9040	0.9170	2.3300	-2.1300	0.9670
	(0.2390)	(0.4230)	(0.0455)	(0.5210)	(0.4570)	(0.0635)	(0.0456)	(0.2610)	(0.0021)	(0.9370)	(0.5610)	(0.1380)	(0.9930)	(0.5470)
12	1.9300	0.7280	2.8000***	2.0900	3.1200**	1.4900	2.9700***	1.6000	5.8400*	-0.0419	2.5800	2.1100	-1.1300	2.2700
	(0.2300)	(0.6120)	(0.0632)	(0.1830)	(0.0412)	(0.3910)	(0.0515)	(0.3340)	(0.0007)	(0.8230)	(0.1070)	(0.1850)	(0.9560)	(0.1610)

Note: Bootstrap critical values calculated from 10000 replications.

Table 9 Threshold autoregressive (TAR) unit root test. 1-month forward exchange rate premium series.

k=1	USA		Switzerland		Japan		Canada		France		Germany		Italy	
Delay Order	R_{1T} - statistic (Boot p-value)	R_{2T} - statistic (Boot p-value)	R_{1T} - statistic (Boot p-value)	R_{2T} - statistic (Boot p-value)	R_{1T} - statistic (Boot p-value)	R_{2T} - statistic (Boot p-value)	R_{1T} - statistic (Boot p-value)	R_{2T} - statistic (Boot p-value)	R_{1T} - statistic (Boot p-value)	R_{2T} - statistic (Boot p-value)	R_{1T} - statistic (Boot p-value)	R_{2T} - statistic (Boot p-value)	R_{1T} - statistic (Boot p-value)	R_{2T} - statistic (Boot p-value)
1	0.8940	0.9300	0.4330	0.4680	0.0104**	0.0119**	0.1850	0.2040	0.0541***	0.0593***	0.4050	0.4410	0.7500	0.7960
2	0.3020	0.3320	0.1190	0.1350	0.0341**	0.0386**	0.2020	0.2240	0.7870	0.8300	0.1710	0.1940	0.2920	0.3210
3	0.1720	0.1930	0.2480	0.2750	0.0476**	0.0545***	0.1660	0.1870	0.1240	0.1350	0.1210	0.1360	0.9110	0.9530
4	0.1320	0.1500	0.1380	0.1560	0.0270**	0.0303**	0.1590	0.1810	0.0000*	0.0000*	0.2550	0.2500	0.5860	0.5760
5	0.3860	0.4230	0.0454**	0.0545***	0.0102**	0.0117**	0.0713***	0.0824***	0.7190	0.7620	0.2170	0.2430	0.7950	0.8150
6	0.8980	0.9400	0.1500	0.1700	0.0354**	0.0402**	0.0028*	0.0030*	0.0362**	0.0390**	0.3390	0.3520	0.0105**	0.0121**
7	0.4190	0.4570	0.1470	0.1640	0.0346**	0.0392**	0.0406**	0.0473**	0.0581***	0.0613***	0.3150	0.3500	0.2550	0.2850
8	0.1410	0.1630	0.0385**	0.0460**	0.0299**	0.0337**	0.0014*	0.0014*	0.0000*	0.0000*	0.2100	0.2380	0.7840	0.8320
9	0.0227**	0.0256**	0.0475**	0.0551***	0.0291**	0.0331**	0.0028*	0.0032*	0.0000*	0.0000*	0.2790	0.3110	0.1120	0.1240
10	0.1720	0.1930	0.0067*	0.0081*	0.0939***	0.1060	0.0102**	0.0110**	0.0293**	0.0318**	0.0285**	0.0322**	0.7330	0.7790
11	0.3490	0.3860	0.0990***	0.1110	0.1200	0.1340	0.0516***	0.0580***	0.0069*	0.0062*	0.2830	0.3150	0.8540	0.3600
12	0.4470	0.4860	0.0472**	0.0540***	0.0689***	0.0761***	0.0729***	0.0813***	0.0016*	0.0017*	0.0861***	0.0957***	0.3630	0.2960

Note: Bootstrap critical values calculated from 10000 replications.

Table 10 Threshold autoregressive (TAR) unit root test. 3-month forward exchange rate premium series.

k=3	USA		Switzerland		Japan		Canada		France		Germany		Italy	
Delay Order	t_1 - statistic (Boot p-value)	t_2 - statistic (Boot p-value)	t_1 - statistic (Boot p-value)	t_2 - statistic (Boot p-value)	t_1 - statistic (Boot p-value)	t_2 - statistic (Boot p-value)	t_1 - statistic (Boot p-value)	t_2 - statistic (Boot p-value)	t_1 - statistic (Boot p-value)	t_2 - statistic (Boot p-value)	t_1 - statistic (Boot p-value)	t_2 - statistic (Boot p-value)	t_1 - statistic (Boot p-value)	t_2 - statistic (Boot p-value)
1	2.8700***	0.0461	1.8600	1.1300	3.6200**	2.1800	-1.0500	3.6000*	4.2000*	-0.6170	0.7680	3.1900*	-0.5130	1.5500
	(0.0539)	(0.8040)	(0.2320)	(0.4810)	(0.0163)	(0.1640)	(0.9530)	(0.0112)	(0.0044)	(0.9080)	(0.5960)	(0.0286)	(0.9010)	(0.3360)
2	1.4300	1.6100	3.2400**	2.0800	3.2900**	0.7070	-0.1620	2.0400	3.6800**	0.6480	0.7860	2.4700	-1.8600	1.2900
	(0.3650)	(0.3130)	(0.0235)	(0.1840)	(0.0282)	(0.6110)	(0.8410)	(0.1890)	(0.0124)	(0.6390)	(0.5870)	(0.1020)	(0.9890)	(0.4200)
3	2.9200***	1.0500	-1.4600	3.3400**	3.0100**	2.6900***	2.4600	3.2500**	2.4900	0.4470	1.1500	2.4300	0.9300	1.3000
	(0.0514)	(0.5020)	(0.9730)	(0.0201)	(0.0465)	(0.0806)	(0.1070)	(0.0234)	(0.1010)	(0.7010)	(0.4580)	(0.1100)	(0.5330)	(0.4270)
4	1.4700	0.6190	0.8100	3.2100**	2.1200	4.1200*	1.9000	2.9300**	2.3000	0.4280	0.4390	1.8900	-0.4660	0.6400
	(0.3540)	(0.6490)	(0.5830)	(0.0278)	(0.1880)	(0.0047)	(0.2260)	(0.0459)	(0.1370)	(0.7000)	(0.6880)	(0.2300)	(0.8820)	(0.6410)
5	0.3740	1.1700	-0.1030	2.4800	2.4000	2.6300***	2.2200	2.9000***	3.2700**	0.9290	1.0700	1.9900	0.7370	0.9020
	(0.7090)	(0.4690)	(0.8250)	(0.1030)	(0.1310)	(0.0921)	(0.1510)	(0.0503)	(0.0331)	(0.5550)	(0.4890)	(0.2030)	(0.5830)	(0.5680)
6	0.8400	1.5300	-0.0184	2.3800	3.5800**	-0.5300	2.0400	3.7000**	5.3100*	0.1400	1.0400	2.1500	-0.3810	0.5350
	(0.5690)	(0.3450)	(0.8040)	(0.1200)	(0.0239)	(0.8830)	(0.1920)	(0.0102)	(0.0005)	(0.7790)	(0.5110)	(0.1640)	(0.8700)	(0.6790)
7	3.0600**	1.6300	2.5900****	2.9500**	2.8600***	1.7200	2.0000	2.5600***	4.0200*	-1.1400	1.4300	2.0100	2.6300***	0.6130
	(0.0431)	(0.3200)	(0.0835)	(0.0474)	(0.0755)	(0.3030)	(0.2100)	(0.0957)	(0.0069)	(0.9610)	(0.3810)	(0.1950)	(0.0894)	(0.6550)
8	2.9500**	0.7600	2.5800***	2.2500	1.4100	3.1900**	4.1800*	1.5900	4.8500*	-0.2280	2.5000	1.8400	2.4400	0.7110
	(0.0499)	(0.6010)	(0.0849)	(0.1540)	(0.3970)	(0.0416)	(0.0034)	(0.3300)	(0.0016)	(0.8580)	(0.1120)	(0.2440)	(0.1180)	(0.6300)
9	2.5500***	0.5870	1.5500	3.1000**	1.6800	3.0700***	2.9400***	1.9500	0.7920	4.1600*	1.3800	2.1200	0.0246	0.0932
	(0.0956)	(0.6600)	(0.3430)	(0.0403)	(0.3200)	(0.0519)	(0.0510)	(0.2260)	(0.5820)	(0.0074)	(0.4000)	(0.1760)	(0.7970)	(0.7950)
10	2.7900***	0.7640	1.4200	2.9200***	2.0000	2.8600***	1.9200	2.4300	5.5900*	-0.6570	1.8200	1.8800	-0.5780	-0.1870
	(0.0658)	(0.6060)	(0.3850)	(0.0547)	(0.2380)	(0.0763)	(0.2360)	(0.1150)	(0.0007)	(0.9130)	(0.2660)	(0.2400)	(0.9010)	(0.8450)
11	2.8500***	0.8210	-0.1270	2.7900***	2.7300***	2.0300	2.2000	2.2900	6.5200*	-0.6490	1.6000	2.0800	0.3850	-0.1060
	(0.0597)	(0.5860)	(0.8310)	(0.0761)	(0.0999)	(0.2410)	(0.1670)	(0.1410)	(0.0001)	(0.9140)	(0.3280)	(0.1950)	(0.7050)	(0.8340)
12	1.2900	0.7420	0.4970	2.5900***	3.5400**	1.1100	2.4400	1.4600	4.4200*	-0.2690	1.6800	2.1200	0.6310	0.6170
	(0.4190)	(0.6130)	(0.6850)	(0.0988)	(0.0318)	(0.5250)	(0.1170)	(0.3810)	(0.0058)	(0.8650)	(0.3010)	(0.1860)	(0.6320)	(0.6640)

Note: Bootstrap critical values calculated from 10000 replications.

Table 11 Threshold autoregressive (TAR) unit root test. 3-month forward exchange rate premium series.

k=1	USA		Switzerland		Japan		Canada		France		Germany		Italy	
Delay Order	R_{1T} - statistic (Boot *p*-value)	R_{2T} - statistic (Boot *p*-value)	R_{1T} - statistic (Boot *p*-value)	R_{2T} - statistic (Boot *p*-value)	R_{1T} - statistic (Boot *p*-value)	R_{2T} - statistic (Boot *p*-value)	R_{1T} - statistic (Boot *p*-value)	R_{2T} - statistic (Boot *p*-value)	R_{1T} - statistic (Boot *p*-value)	R_{2T} - statistic (Boot *p*-value)	R_{1T} - statistic (Boot *p*-value)	R_{2T} - statistic (Boot *p*-value)	R_{1T} - statistic (Boot *p*-value)	R_{2T} - statistic (Boot *p*-value)
1	0.1440	0.1610	0.3860	0.4210	0.0121**	0.0138**	0.0338**	0.0272**	0.0126**	0.0127**	0.0711***	0.0795***	0.6590	0.6680
2	0.3930	0.4270	0.0194**	0.0224**	0.0702***	0.0790***	0.4430	0.4750	0.0335**	0.0370**	0.2270	0.2550	0.7570	0.3910
3	0.1010	0.1130	0.0582***	0.0340**	0.0205**	0.0227**	0.0111**	0.0129**	0.2570	0.2830	0.1920	0.2170	0.6520	0.6950
4	0.6390	0.6830	0.0661***	0.0749***	0.0071*	0.0078*	0.0467**	0.0552***	0.3250	0.3570	0.4890	0.5340	0.9150	0.9310
5	0.7820	0.8270	0.2640	0.2920	0.0607***	0.0669***	0.0335**	0.0400**	0.0674***	0.0738***	0.3480	0.3850	0.8000	0.8450
6	0.5790	0.6210	0.3040	0.3380	0.0628***	0.0665***	0.0086*	0.0100**	0.0013*	0.0015*	0.3050	0.3390	0.9310	0.9550
7	0.0578***	0.0655***	0.0180**	0.0202**	0.0960***	0.1080	0.0821***	0.0945***	0.0207**	0.0161**	0.2790	0.3110	0.2170	0.2420
8	0.1160	0.1280	0.0552***	0.0632***	0.0818***	0.0911***	0.0060*	0.0065*	0.0043*	0.0042*	0.1080	0.1240	0.2720	0.2990
9	0.2230	0.2530	0.0524***	0.0590**	0.0804***	0.0893***	0.0488**	0.0563***	0.0183**	0.0198**	0.2650	0.2980	0.9770	0.9990
10	0.1480	0.1680	0.0825***	0.0948***	0.0831***	0.0937***	0.1110	0.1260	0.0012*	0.0011*	0.2380	0.2690	0.9830	0.9640
11	0.1330	0.1510	0.1870	0.2080	0.1030	0.1140	0.0968***	0.1080	0.0001*	0.0001*	0.2400	0.2690	0.9500	0.9830
12	0.6810	0.7270	0.2320	0.2560	0.0621***	0.0686***	0.1680	0.1870	0.0123	0.0127**	0.2160	0.2430	0.8820	0.9230

Note: Bootstrap critical values calculated from 10000 replications.

Table 12 Least Squares estimates of the threshold model. 1-month forward exchange rate premium series.

k=1	USA		Switzerland		Japan		Canada		France		Germany		Italy	
Regime	$Z_{t-1} < \lambda$	$Z_{t-1} > \lambda$	$Z_{t-1} < \lambda$	$Z_{t-1} > \lambda$	$Z_{t-1} < \lambda$	$Z_{t-1} > \lambda$	$Z_{t-1} < \lambda$	$Z_{t-1} > \lambda$	$Z_{t-1} < \lambda$	$Z_{t-1} > \lambda$	$Z_{t-1} < \lambda$	$Z_{t-1} > \lambda$	$Z_{t-1} < \lambda$	$Z_{t-1} > \lambda$
Threshold	-0.0013		0.0002		0.0002		-0.0010		-0.0016		-0.0021		-0.0021	
m	4		1		7		2		5		11		6	
(%) Obs	13.7	83.6	65.2	34.8	54.6	45.4	13.2	86.8	13.5	86.5	14.5	85.5	15.4	84.6
Intercept	0.0008	0.0001	0.0003	0.0005	0.0007	0.0008	0.0010	0.0001	-0.0003	0.0001	0.0002	0.0002	-0.0006	0.0000
	(0.0003)	(0.0001)	(0.0002)	(0.0002)	(0.0003)	(0.0003)	(0.0003)	(0.0001)	(0.0006)	(0.0001)	(0.0004)	(0.0001)	(0.0006)	(0.0001)
fp_{t-1}	-0.2070	-0.0468	-0.0468	-0.0653	-0.1610	-0.1330	-0.1030	-0.0954	-0.1820	-0.0361	-0.0735	-0.0515	-0.5770	0.0023
	(0.0816)	(0.0308)	(0.0329)	(0.0438)	(0.0519)	(0.0664)	(0.1150)	(0.0376)	(0.1530)	(0.0474)	(0.0801)	(0.0221)	(0.1260)	(0.0408)
Δfp_{t-1}	-0.1070	0.1440	0.0268	-0.6090	-0.2100	-0.2130	0.2690	0.0987	-0.2530	-0.2710	-0.0943	-0.1960	0.1530	-0.3520
	(0.1710)	(0.0708)	(0.1160)	(0.1670)	(0.1070)	(0.1190)	(0.1870)	(0.0840)	(0.1720)	(0.0857)	(0.1540)	(0.0805)	(0.1690)	(0.0868)
Δfp_{t-2}	0.1650	-0.1510	-0.0643	0.0948	0.2500	-0.1200	0.5390	0.0372	0.4620	-0.1720	-0.3200	0.0716	0.6720	-0.1240
	(0.2390)	(0.0693)	(0.0780)	(0.1290)	(0.1170)	(0.1300)	(0.1910)	(0.0765)	(0.1890)	(0.0874)	(0.1510)	(0.0756)	(0.2280)	(0.0942)
Δfp_{t-3}	0.2310	-0.1570	-0.1260	-0.2620	0.0181	-0.2200	-0.0068	-0.0140	0.7960	-0.0206	0.4850	0.0224	0.5550	-0.0402
	(0.1380)	(0.0703)	(0.0755)	(0.1300)	(0.1160)	(0.1290)	(0.1500)	(0.0695)	(0.2360)	(0.0814)	(0.1420)	(0.0809)	(0.2250)	(0.0902)
Δfp_{t-4}	0.3610	-0.1210	0.0472	-0.3500	0.0139	-0.0162	0.5550	-0.0925	0.4740	-0.1510	0.1460	-0.0616	-0.2090	-0.0696
	(0.1500)	(0.0738)	(0.0795)	(0.1150)	(0.1140)	(0.1290)	(0.2220)	(0.0613)	(0.2290)	(0.0773)	(0.1320)	(0.0772)	(0.1760)	(0.0837)
Δfp_{t-5}	0.0231	0.0293	0.0161	0.0205	-0.2160	-0.1140	-0.2420	-0.0802	-1.1000	-0.3560	0.0266	-0.0794	0.5720	-0.1360
	(0.1410)	(0.0651)	(0.0775)	(0.1190)	(0.1130)	(0.1290)	(0.1940)	(0.0649)	(0.2690)	(0.0751)	(0.1290)	(0.0755)	(0.2620)	(0.0783)
Δfp_{t-6}	-0.5700	0.1680	0.0803	-0.2580	0.1710	-0.0941	0.7200	-0.1640	-0.3220	-0.1080	-0.3120	-0.0607	0.5630	-0.0943
	(0.1120)	(0.0718)	(0.0876)	(0.0931)	(0.1030)	(0.1210)	(0.1580)	(0.0645)	(0.1600)	(0.0794)	(0.1460)	(0.0753)	(0.2790)	(0.0883)
Δfp_{t-7}	-0.0639	0.1050	-0.0194	0.2650	0.0546	0.1070	-0.2480	-0.0177	-0.1060	0.2200	-0.5220	0.1590	-0.3060	0.1180
	(0.0979)	(0.0800)	(0.0748)	(0.1180)	(0.1010)	(0.1320)	(0.1150)	(0.0695)	(0.1250)	(0.0790)	(0.1630)	(0.0749)	(0.1290)	(0.0812)
Δfp_{t-8}	0.1080	-0.0910	0.0464	0.2870	0.0157	0.2620	0.2380	0.0463	1.1800	-0.1780	-0.5630	0.0534	0.0795	0.0075
	(0.0930)	(0.0842)	(0.0783)	(0.1110)	(0.0914)	(0.1080)	(0.1170)	(0.0656)	(0.1480)	(0.0677)	(0.1710)	(0.0714)	(0.1540)	(0.0746)
Δfp_{t-9}	-0.0056	-0.0575	-0.0805	0.1550	-0.1800	0.4390	0.0857	-0.2110	0.0879	-0.1120	0.2840	-0.0279	-0.4670	0.1460
	(0.1480)	(0.0701)	(0.0826)	(0.1040)	(0.0853)	(0.1200)	(0.1510)	(0.0634)	(0.1740)	(0.0782)	(0.2560)	(0.0665)	(0.1350)	(0.0668)
Δfp_{t-10}	0.0007	-0.2490	0.0958	0.1190	0.1480	-0.0051	-0.0446	-0.1950	-1.0100	0.0715	0.2000	0.0767	-0.7630	0.0494
	(0.1680)	(0.0707)	(0.0836)	(0.0987)	(0.0956)	(0.0982)	(0.1350)	(0.0651)	(0.2340)	(0.0574)	(0.2060)	(0.0696)	(0.1140)	(0.0713)
Δfp_{t-11}	0.0137	-0.0648	0.0544	-0.2380	-0.0136	-0.0463	0.0394	0.0823	-1.2300	0.0000	1.1300	-0.0586	-0.4490	-0.0046
	(0.1450)	(0.0668)	(0.0731)	(0.1230)	(0.0954)	(0.0852)	(0.1890)	(0.0641)	(0.3420)	(0.0000)	(0.2120)	(0.0670)	(0.1060)	(0.0793)
Δfp_{t-12}	-0.1920	-0.1860	0.1460	-0.2620	0.3360	0.1080	0.0767	-0.0202	0.2980	0.0000	0.9590	0.0039	0.2800	0.0493
	(0.1510)	(0.0667)	(0.0743)	(0.1030)	(0.0858)	(0.0898)	(0.2340)	(0.0623)	(0.1610)	(0.0000)	(0.2050)	(0.0611)	(0.1230)	(0.0677)

Note: Standard errors are shown in parentheses.

Table 13 Least Squares estimates of the threshold model. 3-month forward exchange rate premium series.

k=3	USA		Switzerland		Japan		Canada		France		Germany		Italy	
Regime	$Z_{t-1} < \lambda$	$Z_{t-1} > \lambda$	$Z_{t-1} < \lambda$	$Z_{t-1} > \lambda$	$Z_{t-1} < \lambda$	$Z_{t-1} > \lambda$	$Z_{t-1} < \lambda$	$Z_{t-1} > \lambda$	$Z_{t-1} < \lambda$	$Z_{t-1} > \lambda$	$Z_{t-1} < \lambda$	$Z_{t-1} > \lambda$	$Z_{t-1} < \lambda$	$Z_{t-1} > \lambda$
Threshold	-0.0033		-0.0044		0.0015		-0.0015		-0.0046		-0.0046		-0.0063	
m	7		8		4		1		5		8		7	
(%) Obs	18.9	81.1	15.0	85.0	67.0	33.0	18.5	81.5	15.0	85.0	16.3	83.7	14.1	85.9
Intercept	0.0016	0.0003	-8.60E-06	4.88E-04	0.0013	0.0043	0.0022	0.0005	0.0007	0.0002	-0.0010	0.0004	-0.0025	-0.0003
	(0.0007)	(0.0002)	(1.21E-03)	(3.02E-04)	(0.0006)	(0.0011)	(0.0007)	(0.0002)	(0.0015)	(0.0003)	(0.0006)	(0.0002)	(0.0014)	(0.0003)
fp_{t-1}	-0.1750	-0.0433	-1.44E-01	-4.87E-02	-0.0888	-0.3230	0.0936	-0.1560	-0.3320	-0.0389	-0.1140	-0.0323	-0.2910	-0.0173
	(0.0572)	(0.0266)	(5.58E-02)	(2.17E-02)	(0.0419)	(0.0783)	(0.0894)	(0.0433)	(0.1010)	(0.0419)	(0.0454)	(0.0175)	(0.1100)	(0.0282)
Δfp_{t-1}	0.0633	0.2980	-1.16E-01	-9.39E-03	-0.0159	0.0118	0.6610	-0.0039	0.0721	-0.3230	-0.0093	-0.0301	-0.0270	-0.0123
	(0.1740)	(0.0731)	(1.27E-01)	(7.59E-02)	(0.0917)	(0.1540)	(0.1950)	(0.0951)	(0.1300)	(0.0822)	(0.1380)	(0.0742)	(0.1940)	(0.0866)
Δfp_{t-2}	-0.1590	0.0252	7.75E-02	-9.97E-02	0.1370	0.2870	-0.1900	0.1640	0.3040	-0.0351	0.0332	0.0668	0.3990	0.0167
	(0.1470)	(0.0671)	(1.33E-01)	(7.45E-02)	(0.0889)	(0.1370)	(0.1310)	(0.0802)	(0.1660)	(0.0847)	(0.1280)	(0.0738)	(0.2260)	(0.0869)
Δfp_{t-3}	0.2450	-0.2380	3.64E-02	-4.04E-02	0.0137	0.0978	-0.0915	-0.0585	0.6920	-0.0279	0.2020	-0.1550	0.0454	0.0316
	(0.1480)	(0.0666)	(1.42E-01)	(7.59E-02)	(0.0881)	(0.1450)	(0.1320)	(0.0817)	(0.2280)	(0.0769)	(0.1200)	(0.0725)	(0.2280)	(0.0806)
Δfp_{t-4}	-0.2630	0.0252	3.04E-02	-7.61E-02	0.0450	0.2080	0.1430	-0.0150	0.5680	-0.0738	-0.1980	-0.1310	-0.1320	0.0019
	(0.1350)	(0.0671)	(1.40E-01)	(7.81E-02)	(0.0844)	(0.1590)	(0.1970)	(0.0708)	(0.1830)	(0.0809)	(0.1170)	(0.0704)	(0.2530)	(0.0781)
Δfp_{t-5}	0.6320	-0.1440	-5.08E-01	-7.03E-03	0.0017	-0.0658	-0.0555	-0.0466	-0.0547	-0.3280	-0.0805	-0.0974	-0.0904	-0.0339
	(0.1550)	(0.0677)	(1.35E-01)	(7.43E-02)	(0.0812)	(0.0997)	(0.1260)	(0.0810)	(0.1800)	(0.0767)	(0.1240)	(0.0690)	(0.2150)	(0.0721)
Δfp_{t-6}	0.0988	-0.0389	-3.69E-01	9.48E-02	0.0413	0.0279	0.2970	-0.0696	-0.3150	-0.1240	-0.2970	0.0128	-0.0617	0.0058
	(0.1750)	(0.0667)	(1.42E-01)	(7.49E-02)	(0.0825)	(0.1020)	(0.1230)	(0.0779)	(0.1380)	(0.0826)	(0.1240)	(0.0715)	(0.2730)	(0.0734)
Δfp_{t-7}	0.4360	-0.2330	-2.23E-01	1.54E-01	-0.1550	0.7030	-0.2980	0.0018	-0.4080	0.2130	-0.3540	0.1340	-0.2710	0.1060
	(0.1410)	(0.0712)	(1.59E-01)	(7.51E-02)	(0.0726)	(0.1260)	(0.1230)	(0.0780)	(0.1010)	(0.0887)	(0.1270)	(0.0724)	(0.3940)	(0.0671)
Δfp_{t-8}	-0.2200	0.0635	1.15E-01	1.20E-01	0.0250	0.3980	0.2000	0.0344	0.4340	-0.1310	0.1090	-0.0590	0.3190	0.0554
	(0.1240)	(0.0729)	(1.45E-01)	(6.96E-02)	(0.0803)	(0.1060)	(0.1280)	(0.0749)	(0.1390)	(0.0664)	(0.1390)	(0.0647)	(0.2270)	(0.0692)
Δfp_{t-9}	-0.0742	0.0782	3.85E-01	1.79E-02	-0.0897	0.2340	-0.1550	-0.0959	0.3930	-0.0040	-0.1300	-0.1050	-0.8720	0.2040
	(0.1380)	(0.0699)	(1.13E-01)	(6.56E-02)	(0.0780)	(0.1140)	(0.1080)	(0.0799)	(0.1140)	(0.0768)	(0.1250)	(0.0620)	(0.1970)	(0.0658)
Δfp_{t-10}	0.3990	-0.2560	5.18E-01	-8.58E-03	-0.0537	-0.2920	0.4400	-0.1470	-0.7660	0.0925	-0.2390	-0.0635	-0.7290	-0.0571
	(0.1200)	(0.0681)	(1.27E-01)	(6.92E-02)	(0.0741)	(0.1130)	(0.1380)	(0.0783)	(0.1550)	(0.0680)	(0.1230)	(0.0611)	(0.1880)	(0.0654)
Δfp_{t-11}	-0.1860	0.0913	9.54E-01	6.55E-02	0.1570	0.0742	0.1350	0.0174	-0.4690	-0.0493	0.7420	0.0847	0.1400	0.0353
	(0.1230)	(0.0672)	(1.90E-01)	(5.89E-02)	(0.0815)	(0.0888)	(0.1910)	(0.0695)	(0.2390)	(0.0606)	(0.1460)	(0.0590)	(0.1410)	(0.0673)
Δfp_{t-12}	0.0239	-0.1140	-8.04E-01	7.96E-02	0.0649	0.4680	0.1850	-0.1030	-0.0095	-0.0419	-0.4490	0.0460	0.6990	0.0528
	(0.1410)	(0.0623)	(1.87E-01)	(5.90E-02)	(0.0747)	(0.0948)	(0.1690)	(0.0706)	(0.1140)	(0.0686)	(0.1740)	(0.0590)	(0.1230)	(0.0676)

Note: Standard errors are shown in parentheses.

Table 14 Estimated half-lives of deviations in months.

	Linear		Nonlinear TAR		Nonlinear TAR	
k	1	3	1		3	
			Outside	Inside	Outside	Inside
USA	11.4381	14.1213	2.9886	14.4615	3.6032	15.6589
Switzerland	8.7933	10.7936	14.4615	10.2643	4.4580	13.8835
Japan	4.2650	4.4917	3.9486	4.8568	7.4538	1.7769
Canada	6.0649	6.7713	6.3767	6.9133	-	4.0869
France	5.4223	8.8663	3.4503	18.8521	1.7180	17.4698
Germany	12.4153	15.7706	9.0796	13.1095	5.7267	21.1112
Italy	10.2643	14.1213	0.8056	-	2.0156	39.7187

Note: The half-life is defined as the number of months it takes for deviations to subside permanently below 0.5 in response to a unit shock in the level of the series. All half-lives are reported in months.