Uncovered Interest Rate Parity in a Time-Varying Smooth Transition Regression Model

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ABSTRACT

There are general findings by many studies criticizing the linear model proposed by Fama (1984) is unable to validate the uncovered interest rate parity (UIP) hypothesis. Baillie and Bollerslev (2000) also show that the magnitude and sign of the estimated slope coefficient in regressing spot rate changes on the forward premiums appeared to be slowly time-varying. Particularly, if the estimated slope is negative, existing literature called it "forward premium anomaly." This paper applies the time-varying smooth transition regression (TV-STR) model proposed by Lundbergh, Teräsvirta and van Dijk (2003) to re-examine and distinguish the nonlinearity and structural changes simultaneously in the UIP relationship with monthly observations on the spot and 1-month forward exchange rates of the UK sterling, Canadian dollar and Japanese Yen vis-a-vis the US dollar, respectively, over 1980:12 to 2007:12. The main finding of this study concludes that there occurred nonlinear and time-varying relationships in the forward premium regression for sample currencies. Moreover, the UIP hypothesis and forward premium anomaly alternated along with the time-variation and dynamics of speculative behavior.

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1. Introduction

Uncovered interest rate parity (UIP) states if the financial instruments, says treasury bonds, in two nations are perfect substitutes of each other, the real returns on these financial assets will tend to equal. Applying to the foreign exchange markets, the UIP implies the nominal interest rate on the foreign bonds netting the interest rate on home bonds will be offset by the expected future foreign exchange depreciation at the time bonds matured. In other words, a nation with depreciation expectation on its currency in the future has to stabilize its funds from out flowing by maintaining a relatively higher nominal interest rate currently so as to eliminate the depreciation expectation. Moreover, with the assumptions of rational expectation and no arbitrage, in equilibrium the nominal interest differential on similar financial assets in two nations is equal to the expected return on the spot exchange rate or equivalent to the forward exchange premium/discount between the two nations. Alternatively speaking, under risk neutral hypothesis and foreign exchange market efficiency, the forward premium can be viewed as an unbiased estimator of the return on spot exchange rate, or say the forward exchange rate is an unbiased forecast of future spot rate. This is what we called “the forward rate unbiasedness hypothesis” (FRUH) in the literature, e.g., Bilson (1981), Fama (1984), Froot and Frankel (1989), Engel (1996) and Zivot (2000), etc.

In empirics, since the UIP presumes the expected deprecation or appreciation of the foreign currency should equal the forward premium in two nations, econometrician (e.g. Clarida and Taylor, 1997; Carlson and Osler, 1999, etc.) used to regress the rate of change of the spot rate on the forward exchange premium and to verify if the slope term is one, and the intercept is zero, as well as the existence of serial correlation in the regression residuals to determine the sustainability of UIP hypothesis. Many studies extended Fama (1984) linear model of spot rate returns on forward premium or interest differential in two nations, and generally they were unable to support the UIP. Among these, for example, Froot and

Thaler (1990), Taylor (1995), Marston (1995) and Engel (1996) examined the UIP for different time span in various nations respectively, and all concluded the slope term on forward premium is far less than one and insignificant from zero or even “abnormally negative,” the so-called “anomaly” in the literature. Investigations by Meredith and Chinn (1998), and Flood and Rose (2002) came out with similar UIP anomalies in the 90s for several countries. The anomaly can also be found in sample periods for nations with perfect floating exchange rate regime.

Bekaert and Hodrick (1993) provided evidences showing the estimated slope coefficient fluctuated increasingly during the 80s and tended to be more likely negative. Moreover, Carlson and Osler (1999) investigated fifteen nations under perfect floating exchange rate regime and concluded thirteen of them bearing the UIP anomaly. Froot and Thaler (1990) even surveyed 75 publications regarding UIP hypothesis, the average of the slope estimates is -0.88. Though data applied in these studies are not fully independent causing the average biased, it does give somewhat informal concept the empirics of UIP study have engaged. Therefore, the forward premium anomaly, i.e., the presence of negative slope estimate of the UIP estimation when the interest differential in two nations is replaced by the forward exchange rate premium, proposes a diverse change between forward premium and future exchange rate which indicates a nation with high interest rates may experience a currency appreciation rather than depreciation. To provide the rationality underlying forward premium anomaly, many studies, such as transaction cost hypothesis by Sercu and Wu (2000), the occurrence of speculative restriction by Lyons (2001) and intervention by monetary authority of Mark and Moh (2007), etc., gave interpretation on the appearance of possible nonlinear relationship between spot rate changes and forward exchange rate premiums. This paper, thus, firstly aimed to analyze the UIP hypothesis by taking consideration of transaction and arbitrage costs in that the Sharpe ratio, an index measuring the excess returns via transaction strategy, was taken as a proxy of forward premium (Lyons, 2001; Lucio et. al, 2006) so as to see how the arbitrage behavior resulted from a high value Sharpe ratio over the threshold will affect the forward premium regression estimate causing the slope parameter deviating from or approaching to the theoretical value implied by the UIP. The movement and transition of the parameter estimates associated with different intervals of Sharpe ratio then can help to clarify the nonlinearity of the UIP. Besides, Lucas (1976) criticized a rational individual's expectation on exchange rate behavior will adjust in accordance with the change of time-varying monetary policies. Thereby, the regression estimates of the spot rate change on forward exchange premium should be time-variant and changed along with the time monetary authority intervening and with the duration of monetary policy. Empirical findings by Baillie and Bollerslev (2000) showed evidences supporting a mild transition along with time on the magnitude and sign of forward premium slope estimate. Zhou (2002) also inspected the anomalous effect of time-varying monetary policies on forward premium by including the time trend in the usual regression model and concluded the parameter estimates were not time-invariant.

This study, therefore, secondly tried to simultaneously disclose whether there occur structural changes on the forward premium slope. The key concept of distinguishing the two kinds of nonlinearities is to examine the parameter nonconstancy (van Dijk et al., 2002), i.e., if there remains the second nonlinear regime switching process along with time, in the nonlinear model having Sharpe ratio as the transition variable in the first round of nonlinear estimation. To achieve this objective, we employ Lundbergh et al. (2003)'s time-varying smooth transition regression, TV-STR, to discriminate the two characteristics, i.e., Sharpe ratio nonlinearity and time-varying structural change, at the same time and to offer a reliable interpretation on the forward premium anomaly and why the UIP dynamics maintained or violated.

Our main findings indicate the TV-STR model originated by Lundbergh et al. (2003) not only enables the objectives being achievable but also substantially outperforms the linear and other competitive nonlinear models. Then we employ the TV-STR model to investigate the UIP hypothesis and forward premium anomaly for the UK pound sterling, Canadian dollar and Japanese Yen exchange rates against the US dollar. The empirical results show that the UIP sustained for some periods in each nation, and the UIP hypothesis and forward premium anomaly alternated along with the time-variation and dynamics of speculative behavior.

The article is structured as follows. The first section introduces the theoretical foundation and existing findings for the development of UIP research. After which research methodology is presented, with full details regarding the UIP empirical modeling, diagnostic testing and the estimation procedures used. Empirical results are then presented, with a thorough diagnostic comparison among linear and various nonlinear models, as well as with a complete description on the performance of the TV-STR model in allocating the regime transition caused from the Sharpe ratio fluctuation or time structural change for the UIP dynamics and foreign exchange market efficiency over the sample nations and time span. Finally the conclusions are drawn.

2 . Interest Parity

The covered interest rate parity (CIP) states,

$$f_{t,k} - s_t = i_{t,k} - i^*_{t,k} \quad , \tag{1}$$

where $f_{t,k}$ and s_t denote the natural logarithm of forward exchange rate matured k periods later from time t and of the spot rate (domestic price of foreign currency) at time

t, respectively. $i_{t,k}$ and $i^*_{t,k}$ are defined as the natural logarithm of one plus risk free bond rates, $r_{t,k}$ and $r^*_{t,k}$ matured k periods from time t at home and foreign countries, respectively, i.e., $i_{t,k} = \ln(1 + r_{t,k})$, $i^*_{t,k} = \ln(1 + r^*_{t,k})$. In other words, under efficiency market assumption, the arbitrage, i.e., interest differential, if any from holding the same risk of home and foreign bonds will be covered immediately against future exchange risk by trading forward exchange in the forward exchange market. Researches by most studies, for instance, Taylor (1987), and Sarno and Taylor (2003), supported the CIP hypothesis. Moreover under the assumptions of rational expectation and CIP, the UIP,

$$\underset{t}{E}(s_{t+k}) - s_t = i_{t,k} - i^*_{t,k} = f_{t,k} - s_t \quad , \tag{2}$$

where $E(s_{t+k})$ is the expected spot rate at time $t+k$ based on all information available at date t, indicates FRUH, i.e., the forward exchange rate is an unbiased predictor of future spot rate. Thus, the UIP presumes the expected k-period spot rate change, $E(s_{t+k}) - s_t$ equals the k-period forward premium, $f_{t,k} - s_t$, that is, if a nation's currency is expected to depreciate (appreciate), then it should maintain a higher (lower) interest rate than its foreign counterpart to keep its funds from out flowing.

Most empirical UIP studies modified the linear regression model by Fama (1984) to investigate the relationship between spot rate returns (setting $k = 1$ in equation (2)) and the forward premiums, they examined [1]

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

where $f_{t,1} - s_t$ is the natural logarithm difference between one-month forward exchange rate and spot rate, or say forward premium. The null hypothesis of UIP means $\alpha = 0$,

[1] Please also refer to Baillie and Kilic (2006).

$\beta = 1$ and no serial correlation for the residuals, ε_t. However, if the investors or speculators in the foreign exchange market are not risk neutral, the forward rate may not equal the expected spot rate which will result in a differential, i.e., $f_{t,k} - E(s_{t+k}) = \eta_{t,k}$ or called risk premium, between them. Fama (1984), Froot and Frankel (1989) argued there must exist more or less the deviation of the coefficient estimates from the theoretical values in the empirical UIP regression because of the time-varying risk premiums. Many studies, thus, not only rejected the UIP null hypothesis, but even obtained negative slope estimates of β which is called "forward premium anomaly."

Nonetheless, Goldberg (2000) proposed that the time-varying risk premium is not the only source affecting the UIP regression estimates. McCallum (1994)'s findings suggested the time trend related monetary policy changes may also bring out variations in the UIP regression estimates (see also Zhou, 2002; Sakoulis and Zivot, 2001). This article accordingly utilized Lundbergh, Teräsvirta and van Dijk (2003)'s TV-STR (time-varying smooth transition regression) modified and originated from Teräsvirta (1994)'s smooth transition regression (STR) to clarify simultaneously the influences of nonlinearity and time structure that generate the forward premium anomaly. Then we compare the model performance among linear, STR and TV-STR models to verify if the regression estimates are consistent with the FRUH and UIP conditions.

3. Econometrical Methodology

The first step in estimating the exchange rate data is to test the possibility that the data involve a linear process containing unit roots, since the Luukkonen et al. (1988) linearity tests validated only having conventional asymptotic distributions under assumption of stationarity (under the null of linearity). Only when having rejected the unit root hypothesis for the exchange rate data can the Lundbergh et al. (2003) estimation be applied to investigate the statistical support for the competing models.

3.1 Smooth Transition Regression (STR) Model

Having born the stationary exchange rate data the benchmark nonlinear UIP relationship between spot exchange rate change, Δs_t, and forward premium, $f_{t,1} - s_t$, can be represented in terms of Granger and Teräsvirta (1993), and Teräsvirta (1994) STR model as follows,

$$\Delta s_t = \phi' \mathbf{x_t} + \theta' \mathbf{x_t} \cdot G(\gamma, \mathbf{c}, z_t) + \varepsilon_t \quad , \tag{4}$$

where $\mathbf{x_t} \equiv (1, f_{t,1} - s_t)'$ is the vector of explanatory variables, $\phi \equiv (\phi_0, \phi_1)'$, and $\theta \equiv (\theta_0, \theta_1)'$ are parameter vectors, and $\varepsilon_t \sim iid\ N(0, \sigma^2)$. The general form of transition function used in the literature is the logistic function with K thresholds,

$$G(\gamma, \mathbf{c}, z_t) = \left(1 + \exp\left(-\gamma \prod_{k=1}^{K} \left(z_t - c_k \right) \right) \right)^{-1} \ , \ c_1 \le c_2 \le \ldots \le c_K \, , \ \gamma > 0 \tag{5}$$

where γ is the adjustment speed of a transition function, $\mathbf{c} = [c_1, \cdots, c_k]'$ is location parameter, called threshold values, and z_t is transition variable and very often an element of dependent, or independent variables or some important variables, e.g., the Sharpe ratio in this study. The most common uses of $G(\gamma, \mathbf{c}, z_t)$ are $K = 1$ and $K = 2$. When $K = 1$ the transition function has asymmetric character and is monotonically increasing from 0 to 1 as z_t increases, i.e., the parameter $\phi + \theta \cdot G$ changes monotonically from ϕ to $\phi + \theta$. The adjustment speed γ controls the quickness of the movement on $\phi + \theta \cdot G$. If $\gamma \to \infty$, the model nests the two-regime threshold regression model proposed by Tong (1978). However, when $\gamma = 0$, the STR model degenerates into a linear model. If $K = 2$, i.e., two thresholds (c_1 and c_2), the transition function changes symmetrically around the midpoint $(c_1 + c_2)/2$, where the logistic function attains it minimum value. It is appropriate in situations in which the local dynamic behavior of the process is similar at both large and small values of z_t and different in the middle. When $\gamma \to \infty$, the result is another switching regression model with three regimes such that the two outer regimes are identical and the mid-regime is different from the other two. On the other hand, when $\gamma = 0$, then $G(\gamma, c_1, c_2, z_t) = 1/2$, and the STR model turns to the linear model as the case when $K = 1$.

An alternative to the logistic STR model with $K = 2$ (LSTR2) is the so-called exponential STR (ESTR) model in which the transition function has the form as follows:

$$G_E(\gamma, c, z_t) = 1 - \exp\left(-\gamma(z_t - c)\right), \quad \gamma > 0.$$

$G_E(\gamma, c, z_t)$ is symmetric around $z_t = c$ and has similar shape as the transition function of LSTR2 model. When $\gamma = 0$, then $G_E(\gamma, c, z_t) = 1$, ESTR turns into a linear model, too. It has a drawback, however. When $\gamma \to \infty$, ESTR model becomes practically linear, because $G_E(\gamma, c, z_t) = 0$ at $z_t = c$ (mid-regime is a point) and unity elsewhere (the outer regimes include the whole domain of z_t except $z_t = c$). Due to this almost linearity, we prefer using LSTR2 model instead to track economic dynamics of variables that behave symmetrically.

3.2 Time-Varying Regression (TV-R) Model

Nevertheless, LSTR is not sufficient enough to specify all kinds of nonlinear characteristics for the data sometimes. Lucas' critique reminds that without considering the possibility of time-varying parameter estimates due to policy changes will lead to seriously logical mistakes when we estimate economical behavior equations with time series data. Accordingly, the economical regression parameters may change smoothly or promptly along with the time-varying government policies altering. It turns out the regression estimates may exhibit parameter nonconstancy of structural change and can then be framed simply by replacing the transition variable z_t in equation (5) with $t^* = t/T$, where T is the sample size, i.e.,

$$H(\gamma_h, \mathbf{c}_h, t^*) = \left(1 + \exp\left(-\gamma_h \prod_{k=1}^{K_h}\left(t^* - c_{h,k}\right)\right)\right)^{-1}, \quad c_{h,1} \le c_{h,2} \le \ldots \le c_{h,K_h}, \quad \gamma_h > 0. \qquad (6)$$

Substituting $G(\cdot)$ in equation (4) with the transition function $H(\cdot)$ above gives the time-varying regression (TV-R) of Lin and Teräsvirta (1994). The magnitude of γ_h governs the rate of structural change. As γ_h approaches zero, the parameters vary inactively, on the contrary, if $\gamma_h \to \infty$, the parameters change instantly.

3.3 Time-Varying Smooth Transition Regression (TV-STR) Model

To simultaneously specify the dynamics of the time-varying coefficients and their nonlinearity within regime transmission, TV-STR (time-varying STR) model modifies the assumption of parameter constancy with the following extension from STR model,

$$\Delta s_t = \phi(t)'\mathbf{x_t} + \theta(t)'\mathbf{x_t} \cdot G(\gamma_g, \mathbf{c}_g, z_t) + \varepsilon_t \quad , \qquad (7)$$

$$\phi(t) \equiv \phi + \lambda_\phi H(\gamma_h, \mathbf{c}_h, t^*) \quad , \qquad (7a)$$

$$\theta(t) \equiv \theta + \lambda_\theta H(\gamma_h, \mathbf{c}_h, t^*) \quad , \qquad (7b)$$

where $\mathbf{x_t} = (1, \tilde{x}_t)' = (1, f_{t,1} - s_t)'$ is the vector of explanatory variables, and $\varepsilon_t \sim iid\ N(0, \sigma^2)$. $\phi = (\phi_1, \phi_2)'$, $\theta = (\phi_3, \phi_4)'$, $\lambda_\phi = (\phi_5, \phi_6)'$, and $\lambda_\theta = (\phi_7, \phi_8)'$ are all (2×1) parameter vectors. $G(\gamma_g, \mathbf{c}_g, z_t)$ and $H(\gamma_h, \mathbf{c}_h, t^*)$ are transition functions handling the parameter nonlinearity between regimes and parameter nonconstancy along with time changes, respectively. Both $G(\cdot)$ and $H(\cdot)$ are ranged in between 0 and 1 with z_t, t^* being the transition variables, γ_g, γ_h being the speed of transition, and $\mathbf{c}_g$ and $\mathbf{c}_h$ being the thresholds, respectively. Either the time-varying coefficients or smooth nonlinear parameters or both can be characterized by the two parameter vectors, $\phi(t)$ and

$\theta(t)$, where $\phi(t)$ changes within the closed interval $[\phi, \phi+\lambda_\phi]$, and $[\theta, \theta+\lambda_\theta]$ for $\theta(t)$.

3.4 Estimation and Evaluation

In building up the investigation on possible UIP nonlinearity and structural change, we refer to Lundberg et al. (2003)'s specific-to-general procedure that begins with testing linearity against the STR model.

3.4.1 Linearity test

From the STR specification in section 3.1, we realize that STR model reduces to a linear model once the speed parameter of the transition, γ, approaches zero. Therefore, it is straight forward to identify the linearity of a regression model by simply test if γ is vanished in equation (5).

Recalling the standard STR model of equations (4) and (5), Luukkonen et al. (1988) and Teräsvirta (1994) suggested using a second order and third order Taylor expansion to approximate the LSTR1 and LSTR2 models around $\gamma = 0$ separately. Then the auxiliary regression for testing the linearity takes the form

$$\Delta s_t = \beta_0{}' \mathbf{x_t} + \sum_{j=1}^{3} \beta_j{}' \cdot \tilde{x}_t \cdot z_t^j + \varepsilon_t^*, \quad t = 1, \cdots, T, \tag{8}$$

where β_0' and β_j' are coefficient matrices, $\varepsilon_t^* = \varepsilon_t + R_3(\gamma, \mathbf{c}, z_t) \cdot \theta' \cdot \tilde{x}_t$ is white noise and $R_3(\gamma, \mathbf{c}, z_t)$ is a Taylor remainder. The null hypothesis of linearity is $\mathrm{H}_0: \beta_1 = \beta_2 = \beta_3 = 0$, because each β_j, $j = 1, 2, 3$, is of the form $\gamma \cdot \tilde{\beta}_j$, where $\tilde{\beta}_j \neq 0$ is a function of θ and $\mathbf{c}$. The auxiliary regression in equation (8) could also be used to select the appropriate specification of the transition function. The reparameterizations of the Taylor expansion of LSTR1 imply that rejection of the null of $\beta_1 = 0$ or $\beta_3 = 0$, whereas the Taylor expansion of the LSTR2 or ESTR under the null involves only the second order term, that is $\beta_2 \neq 0$. Following Teräsvirta (1994), the null and alternative hypotheses in testing the linearity can thus be summarized below:

$\mathrm{H}_0: \beta_1 = \beta_2 = \beta_3 = 0$ vs. H_1: nonlinear model

$\mathrm{H}_0^1: \beta_3 = 0$ vs. H_1^1: LSTR1 model

$\mathrm{H}_0^2: \beta_2 = 0 \mid \beta_3 = 0$ vs. H_1^2: LSTR2/ESTR model

$\mathrm{H}_0^3: \beta_1 = 0 \mid \beta_2 = \beta_3 = 0$ vs. H_1^3: LSTR1 model

First of all, if $\mathrm{H}_0: \beta_1 = \beta_2 = \beta_3 = 0$ is rejected, a nonlinear regression is more appropriate to fit for the empirical data. Next, we need to perform tests on H_0^1, H_0^2, and H_0^3, and choose the one for which the p-value of the test is minimized then determine for which the best nonlinear model specification is going to be implemented.

After we finish up the regression estimation for the nonlinear STR model chosen following previous description, we may further verify parameter constancy and remaining nonlinearity for STR model estimation. Estimate a TV-STR model if the null of no remaining nonlinearity or parameter constancy is failed.

3.4.2 Parameter constancy test and no remaining nonlinearity test

The purpose of parameter constancy test (PC test) is to assess if the parameter estimates of a nonlinear regression are changing over time, i.e., whether there exists structural change on parameters. If the null hypothesis of parameter constancy can not be rejected, we conclude the parameters are time-invariant. However, when we use the long term sample data to estimate time series models, it may result in parameter nonconstancy and an increase of estimation errors. The implementation of PC test thus helps avoiding problematic estimation, inference and forecasting. Analogous to the formation of linearity test described previously, the null hypothesis of PC test for an STR specification for instance, i.e., $\mathrm{H}_0: \gamma_h = 0$ in equations (7a) and (7b), can be verified by plugging a first ord er Taylor approximation of $H(\cdot)$ around $\gamma_h = 0$ assuming $K_h = 3$ at most,

$$H(\gamma_h, \mathbf{c}_h, t^*) \approx 1 \Big/ 2 + (\gamma_h/2)\sum_{j=0}^{K_h} \delta_j (t^*)^j \ , \quad K_h = 1 \text{ or } 2 \text{ or } 3 \ , \tag{9}$$

into equation (7) to form the following auxiliary regression,

$$\Delta s_t = \beta_0' \mathbf{x_t} + \beta_1' \mathbf{x_t} \cdot G(\gamma_g, \mathbf{c}_g, z_t) + \sum_{j=1}^{3} \beta_{j+1}' \cdot \tilde{x}_t \cdot (t^*)^j + \sum_{j=1}^{3} \beta_{j+4}' \cdot \tilde{x}_t \cdot (t^*)^j \cdot G(\gamma_g, \mathbf{c}_g, z_t) + \varepsilon_t^* \ . \tag{10}$$

The PC test is equivalent to test the following new hypotheses,

$\mathrm{H}_0^1: \beta_4 = \beta_7 = 0 \ (K_h<3)$ vs. $\mathrm{H}_1^1: \beta_4 \neq 0 \text{ or } \beta_7 \neq 0 \ (K_h=3)$

$\mathrm{H}_0^2: \beta_3 = \beta_6 = 0 \,|\, \beta_4 = \beta_7 = 0 \ (K_h<2)$ vs. $\mathrm{H}_1^2: \beta_3 \neq 0 \text{ or } \beta_6 \neq 0 \,|\, \beta_4 = \beta_7 = 0 \ (K_h=2)$

$\mathrm{H}_0^3: \beta_2 = \beta_5 = 0 \,|\, \beta_3 = \beta_6 = \beta_4 = \beta_7 = 0 \ (K_h=0)$ vs. $\mathrm{H}_1^3: \beta_2 \neq 0 \text{ or } \beta_5 \neq 0 \,|\, \beta_3 = \beta_6 = \beta_4 = \beta_7 = 0 \ (K_h=1)$

using Wald test. If we reject either one of the nulls, H_0^1, H_0^2 and H_0^3, we conclude $\phi(t)$ and $\theta(t)$ in equation (7) are time-variant and exhibit structural change, otherwise the parameters are time-invariant. Paralleling with what we just developed above, the PC tests can be derived and tested for a TV-R and TV-STR specification as well.

Unlike the PC test, no remaining nonlinearity test (NRN test) examines whether there occurs other nonlinearities after controlling the first nonlinearity. The construction for NRN test statistic shares similar logic as that in the PC test. An auxiliary regression for NRN test can be conducted for an STR model, equation (4), as

$$\Delta s_t = \beta_0' \mathbf{x_t} + \theta' \mathbf{x_t} \cdot G(\gamma, \mathbf{c}, z_t) + \sum_{j=1}^{3} \beta_j' \cdot \tilde{x}_t \cdot w_t^j + \varepsilon_t^* \ , \tag{11}$$

where w_t denotes the transition variable for some possible transition function. If the null hypothesis, $\mathrm{H}_0: \beta_1 = \beta_2 = \beta_3 = 0$, can not hold, there may exist the other nonlinearity suggesting a refinement on the initial STR model is necessary.

3.5 Wald Test for UIP Hypothesis

In this subsection, we construct the Wald statistics for UIP hypothesis in a TV-STR environment. Similar derivation of the Wald tests for the UIP under an STR or TV-R estimation can be easily followed with the same nature. Since there are two transition functions, $G(\cdot)$ and $H(\cdot)$, in the TV-STR model of equations (7), (7a) and (7b) that can be rearranged terms as

$$\Delta s_t = (\phi_1 + \phi_3 G + \phi_5 H + \phi_7 G \cdot H) + (\phi_2 + \phi_4 G + \phi_6 H + \phi_8 G \cdot H)(f_{t,1} - s_t) + \varepsilon_t \ ,$$

$$\equiv \alpha_t + \beta_t \cdot (f_{t,1} - s_t) + \varepsilon_t \qquad (12)$$

The intercept, α_t and slope term, β_t, not only change along with time, $t^* \equiv t/T$ but vary in associate with the shift of transition variable z_t. To verify the UIP, the null hypothesis, $\mathrm{H}_0:\ \alpha_t = 0$ and $\beta_t = 1$ can be expressed in matrix form as

$$\mathrm{H}_0:\ \mathbf{R_t \Phi = r} \ , \qquad (13)$$

where $\mathbf{R_t} = \begin{bmatrix} 1 & 0 & G(\cdot) & 0 & H(\cdot) & 0 & G(\cdot)H(\cdot) & 0 \\ 0 & 1 & 0 & G(\cdot) & 0 & H(\cdot) & 0 & G(\cdot)H(\cdot) \end{bmatrix}$, $\mathbf{\Phi} = [\phi_1 \quad \phi_2 \quad \cdots \quad \phi_8]'$, and $\mathbf{r} = [0 \quad 1]'$. The Wald statistics at each point of time is then

$$Wald_t = (\mathbf{R_t \hat{\Phi} - r})'(\mathbf{R_t \hat{V} R}_t')^{-1}(\mathbf{R_t \hat{\Phi} - r}) \xrightarrow{\ a\ } \chi^2(2) \ , \qquad (14)$$

where $\hat{\mathbf{\Phi}}$ and $\hat{\mathbf{V}}$ are coefficient estimates and covariance matrix of the TV-STR model, respectively. A figure with horizontal axis the time schedule and vertical axis the $Wald_t$ statistics, the chronic switches of the sustainability of UIP can be easily traced and contrasted to the historical events that would cause a shift in $Wald_t$ over time.[2]

3.6 Confidence Intervals of the Spot Rate Changes

Another way to testify the appearance of the relationship between spot rate changes and forward premiums imposed by the UIP theory is to compute the confidence intervals (CI) for the spot rate changes and to check whether the actual values of forward exchange

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[2] A more rigorous test of the causality between a specific event and UIP change need to be implemented to reach a reliable conclusion. However, it is beyond our goal of this study.

premiums fall in between the lower and upper bounds of the confidence intervals with some predetermined level of significance. If the CI cover forward premiums, it directly gives evidences of UIP and CIP, otherwise there may involve some degrees of monetary intervention or other exogenous shocks to the foreign exchange market that fail the UIP relationship.

The Delta method developed by Oehlert (1992) can be utilized to compute the CI for spot rate changes at each point of time. The TV-STR model of equation (7) can be rewritten as

$$\Delta s_t = F(\boldsymbol{\theta}, \mathbf{x_t}, z_t, t^*) + \varepsilon_t \ , \ \ t = 1, \cdots, T \ , \tag{15}$$

where $\boldsymbol{\theta}$ denotes the coefficient vector of all parameters in the TV-STR model. The first order Taylor expansion of $F(\boldsymbol{\theta}, \mathbf{x_t}, z_t, t^*)$ on $\hat{\boldsymbol{\theta}}$ can be approximated as

$$F(\boldsymbol{\theta}, \mathbf{x_t}, z_t, t^*) \approx F(\hat{\boldsymbol{\theta}}, \mathbf{x_t}, z_t, t^*) + \hat{D}^T \cdot (\boldsymbol{\theta} - \hat{\boldsymbol{\theta}}) \ , \tag{16}$$

where $\hat{D}^T = \nabla_{\boldsymbol{\theta}} F(\boldsymbol{\theta}, \mathbf{x_t}, z_t, t^*)\big|_{\hat{\boldsymbol{\theta}}} = \left[\partial F/\partial \boldsymbol{\Phi} \ \ \partial F/\partial \gamma_g \ \ \partial F/\partial \gamma_h \ \ \partial F/\partial \mathbf{c}_g \ \ \partial F/\partial \mathbf{c}_h\right]_{\hat{\boldsymbol{\theta}}}$. Under the assumption that $F(\cdot)$ is differentiable and $\nabla_{\boldsymbol{\theta}} F(\cdot)$ is nonzero, the discrepancy between actual value and forecast of Δs_t is simply

$$\Delta s_t - \Delta \hat{s}_t \approx F(\hat{\boldsymbol{\theta}}, \mathbf{x_t}, z_t, t^*) + \hat{D}^T \cdot (\boldsymbol{\theta} - \hat{\boldsymbol{\theta}}) + \varepsilon_t - F(\hat{\boldsymbol{\theta}}, \mathbf{x_t}, z_t, t^*) = \hat{D}^T \cdot (\boldsymbol{\theta} - \hat{\boldsymbol{\theta}}) + \varepsilon_t \ . \tag{17}$$

According to the asymptotic theory, $\hat{\boldsymbol{\theta}}$ is asymptotically normally distributed, i.e., $\sqrt{T}(\hat{\boldsymbol{\theta}} - \boldsymbol{\theta}) \xrightarrow{a} N(0, \Sigma_{\hat{\boldsymbol{\theta}}})$. Then $F(\boldsymbol{\theta}, \mathbf{x_t}, z_t, t^*)$ is asymptotically normal, too,

$$\sqrt{T}\left(F(\hat{\boldsymbol{\theta}}, \mathbf{x_t}, z_t, t^*) - F(\boldsymbol{\theta}, \mathbf{x_t}, z_t, t^*)\right) \xrightarrow{a} N(0, \Sigma_{F(\hat{\boldsymbol{\theta}})}) \ , \tag{18}$$

where $\Sigma_{F(\hat{\boldsymbol{\theta}})} = \hat{D}^T \Sigma_{\hat{\boldsymbol{\theta}}} \hat{D}$. The $100 \times (1 - \alpha)$ percent CI can then be derived as

$$F(\hat{\boldsymbol{\theta}}, \mathbf{x_t}, z_t, t^*) \pm t_{\alpha/2, T-k} \left(\Sigma_{F(\hat{\boldsymbol{\theta}})} \right)^{1/2}, \tag{19}$$

where k is the number of parameters in the TV-STR estimation, and α is the level of significance.

4. Empirical Results and Analyses

The main purpose of this study is to exam the long term relationship of spot rate changes and forward exchange premiums raised from the UIP hypothesis. We hence restrict our investigation for countries with more completely financial regulation on the foreign exchange market. The exchange rates of British pound sterling, Canadian dollar and Japanese Yen against one unit of US dollar are chosen as the objective rates. The data set consists of the natural logarithms of monthly spot exchange rates (s_t) and of one-month forward exchange rates $(f_{t,1})$ over the period 1980:12-2007:12 total of 325 monthly observations from the Global Financial Database. The spot rate change, $\Delta s_t \equiv s_{t+1} - s_t$, is constructed as the first difference of the logarithm of monthly spot rates. Forward exchange premium, $f_{t,1} - s_t$, and Sharpe ratio ($\equiv ER_t / \sigma_{ER}$, where $ER_t \equiv s_{t+1} - f_{t,1}$) can also be conducted following the definition, where σ_{ER} represents the standard deviation of the excess returns, *ER*, via transaction arbitrage. As mentioned previously in section 3, variables used have to be assured of stationarity to proceed Luukkonen et al. (1988) linearity test. Table 1 reports the unit root test results from two popular methods applied by most practical studies. Both the ADF (Dickey and Fuller, 1979) and PP (Phillips and Perron, 1988) tests reject the null hypothesis of existing nonstationarity for all time series in each nation.

[Table 1 here]

Having satisfied with stationarity for all variables, Table 2 reveals the p-values of the linearity F-tests associated with null hypotheses of $\mathrm{H}_0, \mathrm{H}_0^1, \mathrm{H}_0^2$ and H_0^3, respectively, described in section 3.4.1. The possible candidates for the transition variable considered for being used in the STR estimation are the forward premium and Sharpe ratio in the current period, and one lag before without loss of generality. Following Teräsvirta (1994)'s criteria, it suggests an LSTR2 (logistic smooth transition regression with two thresholds) with transition variable the Sharpe ratio one period before is the most qualified model for estimating symmetric dynamics of both the British pounds and Canadian dollars. An LSTR1 (LSTR model with on threshold), on the other hand, with transition variable still the Sharpe ratio of lag one, is proposed to model asymmetric behavior of UIP relationship. In addition to the consideration of using Sharpe ratio as the transition variable, we also inspect the possible structural changes with $t^* = t/T$ being the transition variable, and the outcome indicate a possible time-varying UIP relationship for the UK and Canada, but not for Japan. The TV-R model, therefore, are carried to compare with other STR models for the UK and Canada only.

[Table 2 here]

Before regressing STR, TV-R or TV-STR models with either forward premium or Sharpe ratio or both as the transition variable, Granger and Teräsvirta (1993) and Teräsvirta (1994) pointed out that the parameter estimation may encounter slow convergency and overestimation problems due to the inappropriate measure unit of the transition variable. They suggested standardizing the transition variable by its own standard deviation before implementing any empirical estimation to avoid the problem. The transition functions, $G(\cdot)$ and $H(\cdot)$, can be re-expressed as

$$G(\gamma_g, \mathbf{c}_g, z_t) = \left(1 + \exp\left(-\gamma_g \prod_{k=1}^{K_g}\left(\frac{z_t - c_{g,k}}{\hat{\sigma}_z}\right)\right)\right)^{-1} \quad \text{and}$$

$$H(\gamma_h, \mathbf{c}_h, t^*) = \left(1 + \exp\left(-\gamma_h \prod_{k=1}^{K_h} \left(\frac{t^* - c_{h,k}}{\hat{\sigma}_{t^*}}\right)\right)\right)^{-1},$$

respectively.

Tables 3, 4 and 5 present the estimation outcome of various nonlinear models (STR, TV-R and TV-STR), likewise the results of linear model for easily making comparisons among them, for the UK, Canada and Japan, respectively. Several informal criteria shown in these tables usually are used to roughly compare the performance of the regression models, for example, we prefer the regression model possessing higher values of adjusted R^2, and log likelihood, but lower values of regression residual standard deviation $\hat{\sigma}_\varepsilon$, the ratio of the actual to estimated standard errors of the spot rate changes $\hat{\sigma}_{\Delta s}/\hat{\sigma}_{\Delta\hat{s}}$, AIC (Akaike information criterion), and CAIC (corrected AIC, Hurvich and Tsai, 1989). Also we favor the nonlinear model than the linear model, if the ratio of nonlinear to linear model residual standard deviation $\hat{\sigma}_{\varepsilon,NL}/\hat{\sigma}_{\varepsilon,L}$ is low, and vice versa. According to the above informal criteria, they all suggested TV-STR model outperforming STR, TV-R models, and the linear model being the worst.

[Tables 3, 4, 5 here]

Though the evidences stated are all in favor of TV-STR model without exception, they did not give a precise measure or significance level of how better the TV-STR is relative to the other models. Fortunately, the TV-STR model will nest to STR, TV-R and even linear models under suitable restrictions of $\gamma_h = 0$, $\gamma_g = 0$, and $\gamma_h = \gamma_g = 0$, respectively. It is straight forward to apply the likelihood ratio (LR) test with some specific levels of significance to get statistical supports for the informal conclusions achieved above. For instance, if we want to check the performance of TV-STR against STR model, the null hypothesis is as H_0: $\gamma_g = 0$ (i.e., STR, the restricted model) and the alternative will be

$H_1: \gamma_g > 0$ (i.e., TV-STR, the unrestricted model). Then the likelihood ratio statistic is simply $LR = -2(L_R - L_U)$, where L_R and L_U are the log likelihood under H_0 and H_1, respectively, and is distributed as χ^2 with degrees of freedom m, the difference of the parameter numbers estimated in TV-STR and STR models. For the other two tests, i.e., TV-STR against TV-R, and TV-STR against linear model, the null and alternative hypotheses can be constructed analogously as

$$\begin{cases} H_0: \gamma_g = 0 \text{ (TV-R)} \\ H_1: \gamma_g > 0 \text{ (TV-STR)} \end{cases} \quad \text{and} \quad \begin{cases} H_0: \gamma_g = \gamma_h = 0 \quad \text{(Linear)} \\ H_1: \gamma_g > 0,\ \gamma_h > 0 \text{ (TV-STR)} \end{cases},$$

respectively. The last row of Tables 3, 4 and 5 give LR statistics of the tests for linear, STR and TV-R models against TV-STR in each nation. All of the LR tests consistently demonstrate a better model specification in favor of TV-STR model in each country.

To evaluate the performance of the regression models estimated, Tables 6, 7 and 8 perform various diagnostic tests, including LM-test for no error autocorrelation, Ljung-Box Q test for serial correlation, and ARCH-LM test for residual heteroskedasticity in Table 6, test of parameter constancy in Table 7, and test of no additive nonlinearity in Table 8. In sum, the results in Table 6 indicate no residual autocorrelation remaining left for both linear and nonlinear models in all currencies examined. However, it seems that except for the TV-STR in all nations and Japan's STR model, all other regression models faced the residual heteroskedasticity. The parameter constancy tests for all nations in Table 7 indicate except for TV-STR model, there remain more or less of structural changes in the parameters. Table 8 also confirms a better character of no remaining nonlinearity for TV-STR model in all exchange rates. Summarizing form the discussion above, the TV-STR model specification satisfies various kinds of diagnostic examinations and prevails over other models for the UK pound, Canadian dollar and Japanese Yen.

[Tables 6, 7 and 8 here]

After well inspecting the evidences verifying that TV-STR model is the most prominent choice among all models investigated in handling the UIP relationship for the UK, Canada and Japan, we then in the final step is to reveal how the UIP relationships between spot rate changes and forward premiums change over time from the TV-STR estimation results. The forward premium anomaly also appears in our linear models for the UK and Canada, especially the UK. The UIP hypothesis was rejected definitely in the linear model for all nations, however, by implementing the Wald test introduced in section 3.5 for UIP hypothesis at each point of time, we can track down the path of Wald statistics over time to reveal the behavior of UIP transition, then to possibly map the historical economic/political events or shocks to the timing of big shifts on the Wald statistics. Figure 1 plots the dynamics of Wald statistics against time for the UK (upper panel), Canada (mid-panel) and Japan (lower panel). The dashed line represents the critical value of the Wald test (χ^2 with degrees of freedom 2) at 5% of significance level. From these three panels of $Wald_t$ dynamics, there do occur structural change, e.g., the UK break point was about 1988:09, the break points for Canada were around 1982:06, 1991:09, and 2001:03, and for Japan the most evident structural break took place about 1995:09. Starting from 1988:09, the UIP relationships for the UK exchange rate appeared to make a big and long lasting change for some reasons from a nonsustained state to a sustainable state. Analogous to the UK pound, the UIP dynamics for Canadian dollar and Japanese Yen can be inferred directly by looking at the $Wald_t$ movements in mid-panel and lower panel of Figure 1, respectively. The UIP hypothesis succeeds for Canadian dollar in the periods before 1982:06 and during 1991:09-2001:03. However, the UIP sustainability for Japanese Yen switches rapidly back and forth over the whole sample period.

[Figure 1 here]

Dooley and Isard (1980) and Spiegel (1990) stated the monetary intervention, capital controls and financial liberalization affecting the international capital mobility could be some of important factors that cause structural changes in the financial market efficiency. One can check with the historical records and find some possible linkage for the events and the break dates just found. However, the causal relations between the event and date still need a more rigorous discussion and test to assure the real causality between them which is beyond our goal of this study.

Another useful tool in identifying the UIP sustainability is to form the confidence intervals for spot rate changes over time under a prespecified level of significance, then to observe whether the confidence interval contains the actual value of forward premium in each period. Let's take the UK pound as an example to double confirm the UIP behavior discussed by the Wald test. Applying the Delta method (Oehlert, 1992) described in section 3.6, we can compute the 95% confidence interval at each point of time using the formula in equation (19). Figure 2 presents the dynamics of actual forward exchange premiums and 95% confidence intervals of estimated spot rate changes for TV-STR model for the UK pound (upper panel), Canadian dollar (mid-panel) and Japanese Yen (lower panel). From the UK phase diagram, prior to 1988:09, the structural break detected by $Wald_t$ in Figure 1, there were only 31 out of 93 months the actual forward exchange premiums fall in between the CI of estimated spot rate changes. In words, for most of the time, two thirds of the period 1981:01-1988:09, the UIP relationship failed to support the efficiency foreign exchange market hypothesis. On the other hand, after 1988:10 till 2007:11, there were 216 out of 230 months (i.e., 94% of the observations) the CI of estimated spot rate changes enclosed actual forward premiums which is quite consistent with our ex-ante level of confidence specification, 95%, that is, the relation between spot rate changes and forward exchange premiums implied by the UIP theory, i.e., efficient foreign exchange market for the UK pound, worked quite successfully during 1988:10-2007:11. Similar to the UK pound addressed, the UIP hypotheses for Canadian dollar and Japanese Yen exchange rates can be easily carried out to reconfirm the structural behavior found by the Wald statistics.

[Figure 2 here]

Figure 3 depicts the movements of transition functions, $G(\cdot)$ and $H(\cdot)$, along with time line, they also show obvious structural breaks in the dynamic path of $H(\cdot)$ for each nation, however the speedy fluctuations of $G(\cdot)$ over the sample period for the UK and Canada seem contributing slight effect on the switches of UIP sustainability. Nevertheless, the rapid changes of $G(\cdot)$ caused by the frequently shifting Sharpe ratio in Japan do generate enormous influences on the swift changes of UIP sustainability. It suggests that the Japanese UIP tends to be tenable (untenable) with a higher (lower) value of one-lag Sharpe ratio before the structural break at round 1995:09 or with a lower (higher) value of one-lag Sharpe ratio after 1995:09.

[Figure 3 here]

Also we calculate the forward premium slope at each point of time for TV-STR specification. The forward premium anomaly occurred constantly in the UK before 1988:09 but not after. In Canada, for most of the time except before 1982:06 and during 1991:09-2001:03, the anomaly seemed quite natural. Oppositely, Japan only encountered the forward premium anomaly for 2 out of 323 months, one in 1981:01 and the other in 1981:07, but its UIP relationships were influenced mainly by the magnitudes of Sharpe ratio over time.

5. Conclusions

The most distinctive contribution of current study from existing literatures is the simultaneous treatments on the parameter nonlinearity among regimes and parameter time-varying character, i.e., structural change along with time, for the UIP relation between spot rate changes and forward exchange premiums. The TV-STR model originated by Lundbergh et al. (2003) not only enables the objectives being achievable but also substantially outperforms other competitive nonlinear models, e.g., the STR and TV-R, in terms of fulfilling various diagnostic tests and model evaluation, such as ARCH-LM test for residual heteroskedasticity, test of parameter constancy, test of no remaining nonlinearity, etc.

This paper makes use of the exchange rate data of the UK pound sterling, Canadian dollar and Japanese Yen over the period 1980:12-2007:12 to analyze the UIP transition and forward premium anomaly. Our major findings from the Wald test on the UIP hypothesis at each point of time indicate there occurred evident structural changes for the UK around 1988:09, for Canada about 1982:06, 1991:09 and 2001:03, and seemed 1995:09 for Japan. In words, the relationships between spot rate changes and forward exchange premiums imposed by the UIP theory for the sample nations examined were not time-invariant due to possible local/global changes or shocks on economic/political conditions. Moreover, the forward premium anomaly took place repeatedly for the UK pound exchange rate prior to 1988:09, but not afterwards. For Canada, the forward premium anomaly happened rather often other than the periods before 1982:06 and during 1991:09-2001:03. On the contrary, the Japanese Yen seemed immunized from the anomaly for most of the sample periods except for January and July in 1981. In addition, the Sharpe ratio appeared to significantly affect the UIP sustainability for the Japanese Yen exchange rate. Especially, Japanese UIP tends to be tenable (untenable) with a higher (lower) value of one-lag Sharpe ratio before the structural break at round 1995:09 or with a lower (higher) value of one-lag Sharpe ratio after 1995:09. In regard to the UK and Canada, however, the Sharpe ratio played trivial role to the UIP sustainability.

To keep the contents of this article from too far-reaching to be untrackable, we do not match the economical chronicle with the structural breaks detected in our empirical findings for the sample nations. However, we cast future research in linking historical records with these breaks and the need for more rigorous examination on this issue.

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Table 1. ADF and PP unit root tests on spot rate changes, forward premiums and Sharpe ratio

		UK (pound)	Canada (dollar)	Japan (Yen)
ADF test	spot rate changes (Δs_t)	-16.9292***	-16.8556***	-16.7426***
	forward premiums ($f_{t,1} - s_t$)	-5.4960***	-9.2512***	-14.6427***
	Sharpe ratio	-5.3472***	-5.3472***	-10.9004***
PP test	spot rate changes (Δs_t)	-18.4371***	-16.8331***	-16.2346***
	forward premiums ($f_{t,1} - s_t$)	-18.4371***	-15.3614***	-16.8050***
	Sharpe ratio	-20.1640***	-15.8069***	-15.2308***

Notes: This table reports statistics for the ADF and PP tests of the null of unit root, respectively.

1. The optimal lag length in the unit root test regression with intercept and trend is chosen on the basis of the Schwartz Bayesian information criterion (SBC).
2. "*", "**", and "***" represent rejection of the null of nonstationarity, I(1), at the 10%, 5%, and 1% of significance levels, respectively.
3. The 10%, 5%, and 1% critical values from MacKinnon (1996) are -3.135 (*), -3.424 (**), and -3.987 (***), respectively, for both ADF and PP tests with intercept and trend.

Table 2. p-values for Teräsvirta (1994) linearity test

UK (pound)	H_0	H_0^1	H_0^2	H_0^3	Model
Forward premium (t)	0.2021	0.3578	0.0526	0.9131	Linear
Sharpe ratio (t)	—	—	0.1504	0.9926	Linear
Forward premium ($t-1$)	—	—	0.0459	0.3018	Linear
Sharpe ratio ($t-1$)	0.0277**	0.1340	0.0496**	0.1245	LSTR2
$t^*(\equiv t/T)$	0.0001***	0.0512*	0.3784	0.0004***	LSTR1
Canada (dollar)	H_0	H_0^1	H_0^2	H_0^3	Model
Forward premium (t)	0.8594	0.4045	0.8063	0.9765	Linear
Sharpe ratio (t)	—	—	0.0351	—	Linear
Forward premium ($t-1$)	—	—	0.7722	0.4867	Linear
Sharpe ratio ($t-1$)	0.0002***	0.7085	0.0001***	0.4443	LSTR2
$t^*(\equiv t/T)$	0.0368**	0.1864	0.1192	0.0544*	LSTR1
Japan (Yen)	H_0	H_0^1	H_0^2	H_0^3	Model
Forward premium (t)	0.0717*	0.0340**	0.8550	0.1161	LSTR1
Sharpe ratio (t)	—	0.2075	0.6896	—	Linear
Forward premium ($t-1$)	—	—	0.6171	0.0484	Linear
Sharpe ratio ($t-1$)	0.0696*	0.0319**	0.3379	0.2746	LSTR1
$t^*(\equiv t/T)$	0.1320	0.0597	0.5355	0.2314	Linear

Notes: This table reports p-values for Teräsvirta (1994) linearity test and the best model selected associated with various transition variables considered.

1. "*", "**", and "***" represent rejection of the null of linear model at the 10%, 5%, and 1% of significance levels, respectively.
2. If the p-value of the linearity test, H_0, is less than the significance level, i.e., the null of linear model is rejected, then we proceed to choose the one for which the p-value of the test is minimized among H_0^1, H_0^2, and H_0^3 and determine for which the best nonlinear model specification is going to be implemented, otherwise if the null can't be rejected, i.e., the p-value of the linearity test, H_0, is greater than the significance level, linear model is the recommended model in the last column of this table.

Table 3. UK (pound): Estimation results for linear, STR, TV-R and TV-STR models

parameter	Linear	STR	TV-R	TV-STR
ϕ_1	0.0017 [0.9977]	0.0047 [-0.1377]	0.0061* [1.9261]	0.0112*** [3.0370]
ϕ_2	-0.8303** [-2.6556]	-0.8036* [-1.8710]	-3.3071*** [-4.9340]	-1.7094** [-2.2969]
ϕ_3	—	0.0230*** [3.7664]	—	-0.0080 [-0.9609]
ϕ_4	—	2.1960* [1.6725]	—	-6.4429*** [-3.7060]
γ_g	—	9.9933 [—]	—	297.3748 [—]
$c_{g,1}$	—	-0.6493 [—]	—	-1.9654*** [-49.6821]
$c_{g,2}$	—	3.5642 [—]		0.8802*** [100.5782]
ϕ_5	—	—	-0.0085** [-2.1492]	-0.0133*** [-3.0393]
ϕ_6	—	—	4.2885*** [4.5031]	2.5022** [2.5207]
ϕ_7	—	—	—	0.0079 [0.7664]
ϕ_8	—	—	—	6.9747*** [3.4338]
γ_h	—	—	26.1509 [—]	9.1052 [—]
c_h	—	—	0.2885 [—]	0.2866*** [8.7652]
Adjusted R^2	0.0106	0.0670	0.0882	0.1316
$\hat{\sigma}_{\varepsilon}$	0.0294	0.0288	0.0284	0.0274
$\hat{\sigma}_{\varepsilon,NL}/\hat{\sigma}_{\varepsilon,L}$	—	0.9796	0.9660	0.9320
$\hat{\sigma}_{\Delta S}/\hat{\sigma}_{\Delta\hat{S}}$	8.5033	3.9745	3.1467	2.4716
AIC	-7.0458	-7.0773	-7.1246	-7.1537
CAIC	-7.0487	-7.1134	-7.1543	-7.2272
Log likelihood	681.57	698.30	691.67	710.00
LR statistics	$56.86^{***}\sim\chi^2(11)$	$23.39^{***}\sim\chi^2(6)$	$36.66^{***}\sim\chi^2(7)$	—

Notes: This table reports estimation results for the UK's linear, STR, TV-R and TV-STR models, and some comparative criteria among models.

1. "*", "**", and "***" represent rejection of the null at the 10%, 5%, and 1% of significance levels, respectively.
2. Numbers in brackets, [ ], represent t values.
3. We prefer the regression model possessing higher values of adjusted R^2, and log

likelihood, but lower values of regression residual standard deviation $\hat{\sigma}_{\varepsilon}$, the ratio of the actual to estimated standard errors of the spot rate changes $\hat{\sigma}_{\Delta s}/\hat{\sigma}_{\Delta \hat{s}}$, AIC (Akaike information criterion), and CAIC (corrected AIC, Hurvich and Tsai, 1989). Also we favor the nonlinear model than the linear model, if the ratio of nonlinear to linear model residual standard deviation $\hat{\sigma}_{\varepsilon,NL}/\hat{\sigma}_{\varepsilon,L}$ is low, and vice versa.

4. Akaike information criterion $(\text{AIC}) = \ln(SSE/T) + 2k/T$, Corrected Akaike information criterion $(\text{CAIC}) = \ln(SSE/T) + 2(k+1)/(T-(k+2))$, where T is the number of observations, SSE is the sum of squared errors, and k is the number of parameters.
5. The null and alternative hypotheses for the three LR tests in the last row of this table are $\begin{cases} \text{H}_0: \gamma_g = \gamma_h = 0 & \text{(Linear)} \\ \text{H}_1: \gamma_g > 0,\ \gamma_h > 0 & \text{(TV-STR)} \end{cases} \sim \chi^2(11)$, $\begin{cases} \text{H}_0: \gamma_h = 0 \text{ (STR)} \\ \text{H}_1: \gamma_h > 0 \text{ (TV-STR)} \end{cases} \sim \chi^2(6)$, and $\begin{cases} \text{H}_0: \gamma_g = 0 \text{ (TV-R)} \\ \text{H}_1: \gamma_g > 0 \text{ (TV-STR)} \end{cases} \sim \chi^2(7)$, from left to right.
6. The critical values at 1% of significance levels for the LR tests above are $\chi^2(11) = 24.73$, $\chi^2(6) = 16.81$, and $\chi^2(7) = 18.48$, respectively.

Table 4. Canada (dollar): Estimation results for linear, STR, TV-R and TV-STR models

parameter	Linear	STR	TV-R	TV-STR
ϕ_1	0.0003 [0.3639]	-0.0017 [-1.7290]	-0.2161 [-0.7655]	-0.0031** [-2.2540]
ϕ_2	-0.2211 [-0.8167]	-0.3839 [-1.2440]	0.0087*** [3.7161]	0.6130 [1.3645]
ϕ_3	—	0.0206*** [6.4549]	—	-0.0017 [-0.2554]
ϕ_4	—	0.9028 [1.6092]	—	-3.8068 [-1.3671]
γ_g	—	20.9637 [0.2764]	—	14.2627 [—]
$c_{g,1}$	—	-1.6723 [—]	—	-1.5746*** [-4.6179]
$c_{g,2}$	—	1.8669 [—]	—	1.8855*** [34.6872]
ϕ_5	—	—	-1.2351 [-1.4058]	0.0015* [0.7872]
ϕ_6	—	—	674.8500 [—]	-1.9517*** [-2.8348]
ϕ_7	—	—	—	0.0262*** [3.3938]
ϕ_8	—	—	—	5.7593** [1.9952]
γ_h	—	—	0.81659 [—]	303.3895 [—]
$c_{h,1}$	—	—	-0.0013 [-1.2585]	0.0571* [1.8418]
$c_{h,2}$	—	—	—	0.4014*** [146.9030]
$c_{h,3}$	—	—	-0.0013 [-1.2585]	0.7512*** [12.3952]
Adjusted R^2	-0.0010	0.1265	0.0588	0.1519
$\hat{\sigma}_{\varepsilon}$	0.0163	0.0155	0.0160	0.0151
$\hat{\sigma}_{\varepsilon,NL}/\hat{\sigma}_{\varepsilon,L}$	—	0.9509	0.9816	0.9264
$\hat{\sigma}_{\Delta S}/\hat{\sigma}_{\Delta\hat{S}}$	22.0067	2.8379	4.2139	2.3488
AIC	-8.2194	-8.3185	-8.2501	-8.3461
CAIC	-8.2223	-8.3546	-8.2798	-8.4320
Log likelihood	871.11	892.12	880.07	904.58
LR statistics	66.93***~$\chi^2(13)$	24.91***~$\chi^2(8)$	49.02***~$\chi^2(9)$	—

Notes: This table reports estimation results for Canada's linear, STR, TV-R and TV-STR models, and some comparative criteria among models.

1. Apply footnotes 1-6 in Table 3 here, except the degrees of freedom in the three LR tests (in footnote 5) are 13, 8, and 9, and the corresponding critical values at 1% of significance levels are $\chi^2(13) = 27.69$, $\chi^2(8) = 20.09$, $\chi^2(9) = 21.67$, respectively.

Table 5. Japan (Yen): Estimation results for linear, STR, TV-R and TV-STR models

parameter	Linear	STR	TV-R	TV-STR
ϕ_1	0.0015 [0.7421]	-0.0017 [-1.0117]	—	-0.0010 [-0.2218]
ϕ_2	0.1054 [0.3379]	0.0076 [0.0138]	—	-1.3737 [-1.5303]
ϕ_3	—	0.0092** [2.1975]	—	0.0043 [0.8667]
ϕ_4	—	0.2667 [0.3982]	—	2.6907*** [2.6841]
γ_g	—	20.9637 [—]	—	111.7142 [—]
c_g	—	-0.1484 [—]	—	-0.5581*** [-8.5336]
ϕ_6	—	—	—	1.8257 [1.5188]
ϕ_8	—	—	—	-4.2634*** [-3.1158]
γ_h	—	—	—	184.0660 [—]
c_h	—	—	—	0.5524*** [17.1830]
Adjusted R^2	-0.0028	0.0237	—	0.0335
$\hat{\sigma}_{\varepsilon}$	0.0330	0.0329	—	0.0324
$\hat{\sigma}_{\varepsilon,NL}/\hat{\sigma}_{\varepsilon,L}$	—	0.9997	—	0.9818
$\hat{\sigma}_{\Delta S}/\hat{\sigma}_{\Delta\hat{S}}$	52.9711	6.4833	—	4.0596
AIC	-7.0458	-7.0773	—	-7.1537
CAIC	-7.0487	-7.1134	—	-7.2272
Log likelihood	644.13	647.43	—	654.15
LR statistics	20.04** $\sim\chi^2(8)$	13.43*** $\sim\chi^2(4)$	—	—

Notes: This table reports estimation results for Japan's linear, STR, TV-R and TV-STR models, and some comparative criteria among models.

1. Apply footnotes 1-6 in Table 3 here, except the degrees of freedom in the first two LR tests (in footnote 5) are 8 and 4, and the corresponding critical values at 1% of significance levels are $\chi^2(8)=15.51$ and $\chi^2(4)$=9.49, respectively.

Table 6. Diagnostic tests for linear, STR, TV-R, and TV-STR models

		Linear	STR	TV-R	TV-STR
UK	LM(6)	0.2370 (0.9625)	0.6650 (0.6781)	0.4865 (0.8183)	0.5028 (0.8061)
	LM(12)	0.8442 (0.5947)	0.9245 (0.5227)	1.1049 (0.3557)	1.1237 (0.3401)
	Q(6)	1.4179 (0.9650)	2.5398 (0.2300)	11.4580 (0.4570)	3.0674 (0.800)
	Q(12)	9.7982 (0.6340)	15.2070 (0.8640)	0.9344 (0.988)	14.6710 (0.1981)
	ARCH-LM(6)	5.0895*** (0.0001)	3.0281*** (0.0069)	4.1453*** (0.0005)	1.96576* (0.0722)
	ARCH-LM(12)	2.9167*** (0.0007)	2.0457** (0.0205)	2.6374*** (0.0023)	1.3629 (0.1831)
Canada	LM(6)	0.4192 (0.8661)	0.8471 (0.5344)	0.4865 (0.8183)	9.8748 (0.1300)
	LM(12)	1.1454 (0.3226)	1.4729 (0.1334)	1.1049 (0.3557)	15.3676 (0.2219)
	Q(6)	2.4443 (0.8746)	4.5713 (0.5761)	4.0643 (0.6680)	2.1262* (0.0503)
	Q(12)	12.2397 (0.4266)	15.0740 (0.2375)	10.3982 (0.5811)	1.5962* (0.0917)
	ARCH-LM(6)	5.5479*** (0.0000)	7.6189 (0.2674)	10.5555*** (0.0051)	1.2016 (0.3052)
	ARCH-LM(12)	4.1021*** (0.0000)	30.3451*** (0.0012)	49.9477*** (0.0000)	2.1180** (0.0158)
Japan	LM(6)	1.6144 (0.1425)	1.5572 (0.1593)	—	1.8481* (0.0778)
	LM(12)	1.5747* (0.0977)	1.2558 (0.2445)	—	1.6986* (0.0550)
	Q(6)	10.0445 (0.1228)	8.5258 (0.2021)	—	11.5996 (0.1145)
	Q(12)	21.4714** (0.0439)	16.3175 (0.1771)	—	21.687* (0.0604)
	ARCH-LM(6)	1.4264 (0.2039)	16.0614 (0.1298)	—	1.4895 (0.1812)
	ARCH-LM(12)	0.9552** (0.04921)	0.2409 (0.2091)	—	0.9682 (0.4793)

Notes: This table reports various diagnostic tests for linear, STR, TV-R, and TV-STR models.

1. "*", "**", and "***" represent rejection of the null at the 10%, 5%, and 1% of significance levels, respectively.
2. Numbers in parentheses, ( ), represent p-values.
3. Without losing generality, we report only statistics for 6- and 12-lag of LM-tests for the null of no error autocorrelation, Ljung-Box Q tests for the null of no serial correlation, and ARCH-LM tests for the null of no residual heteroskedasticity.

Table 7. Parameter constancy tests for STR, TV-R and TV-STR models

		STR	TV-R	TV-STR
UK	H_0^1	4.8086*** (0.001)	1.7902 (0.1306)	1.6471 (0.9900)
	H_0^2	3.1253*** (0.0021)	1.3080 (0.2388)	11.3471 (0.7876)
	H_0^3	2.6429*** (0.0022)	1.9710** (0.0266)	—
Canada	H_0^1	0.7070 (0.5877)	3.3922** (0.0349)	0.1233 (0.9983)
	H_0^2	1.4477 (0.1760)	4.0604*** (0.0032)	0.3043 (0.9960)
	H_0^3	1.5835* (0.0952)	2.9109*** (0.0089)	—
Japan	H_0^1	2.3685* (0.0953)	—	0.1680 (0.9545)
	H_0^2	1.4278 (0.2246)	—	0.5213 (0.8403)
	H_0^3	1.0060 (0.4214)	—	—

Notes: This table reports the results of parameter constancy tests for STR, TV-R and TV-STR models.

1. "*", "**", and "***" represent rejection of the null at the 10%, 5%, and 1% of significance levels, respectively.
2. Numbers in parentheses, ( ), represent p-values.
3. Rejection of either one of the nulls, H_0^1, H_0^2 and H_0^3, will lead a conclusion favoring parameter nonconstancy, otherwise the parameters are time-invariant.

Table 8. No remaining nonlinearity tests for STR, TV-R and TV-STR models

	w_t	STR	TV-R	TV-STR
UK	Sharpe ratio $(t-1)$	—	0.0460**	0.6775
	Time trend (t^*)	0.0015***	—	0.8637
Canada	Sharpe ratio $(t-1)$	—	0.0053***	0.9294
	Time trend (t^*)	0.5572	—	0.9217
Japan	Sharpe ratio $(t-1)$	—	—	0.6236
	Time trend (t^*)	0.6195	—	0.4350

Notes: This table reports the p-values of no remaining nonlinearity tests for STR, TV-R and TV-STR models.

1. "*", "**", and "***" represent rejection of the null at the 10%, 5%, and 1% of significance levels, respectively.
2. Numbers in this table are p-values.
3. w_t denotes the transition variable of other additive nonlinearity in equation (11).

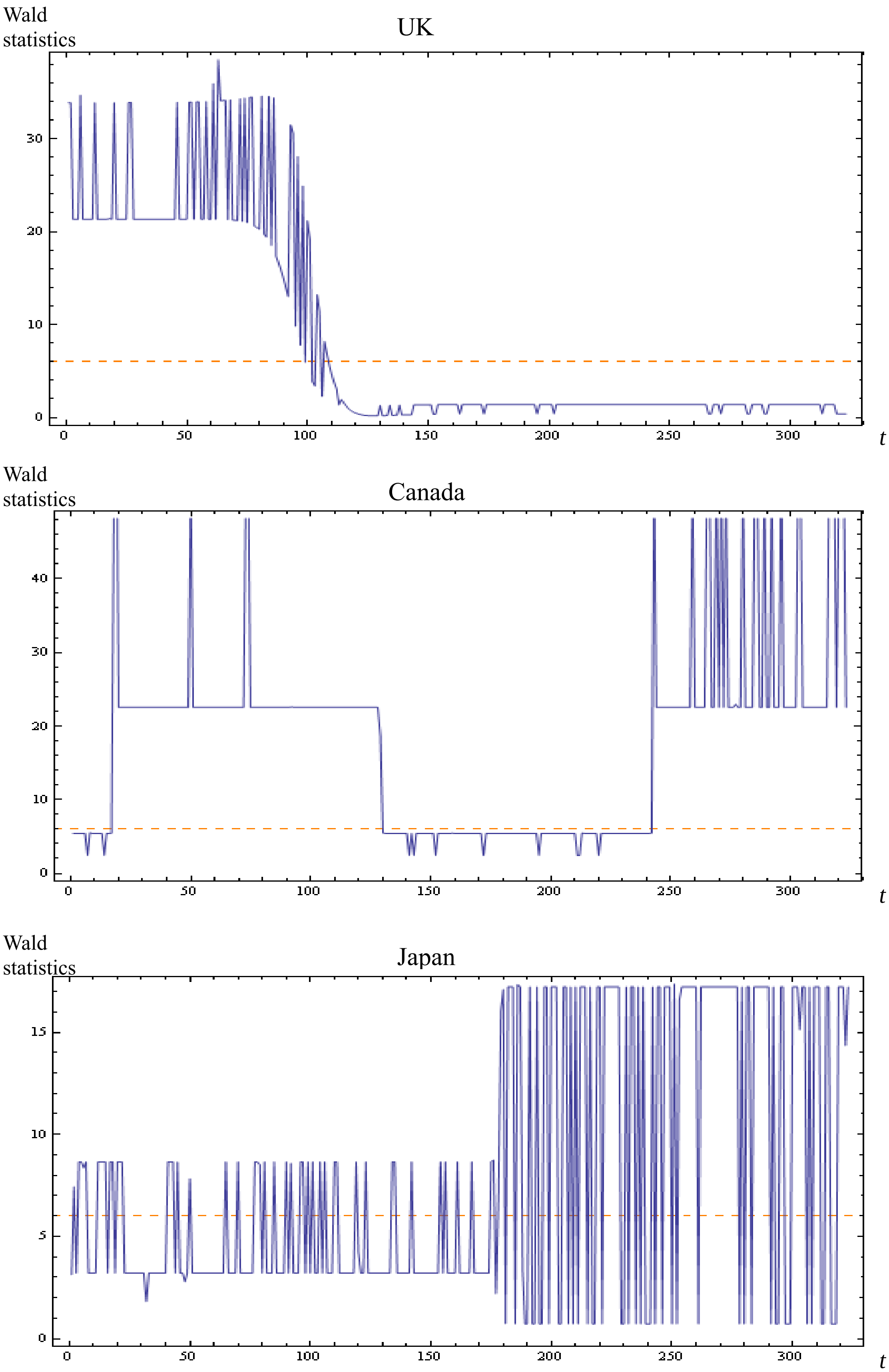

Figure 1. The dynamics of Wald statistics against time for the UK, Canada and Japan

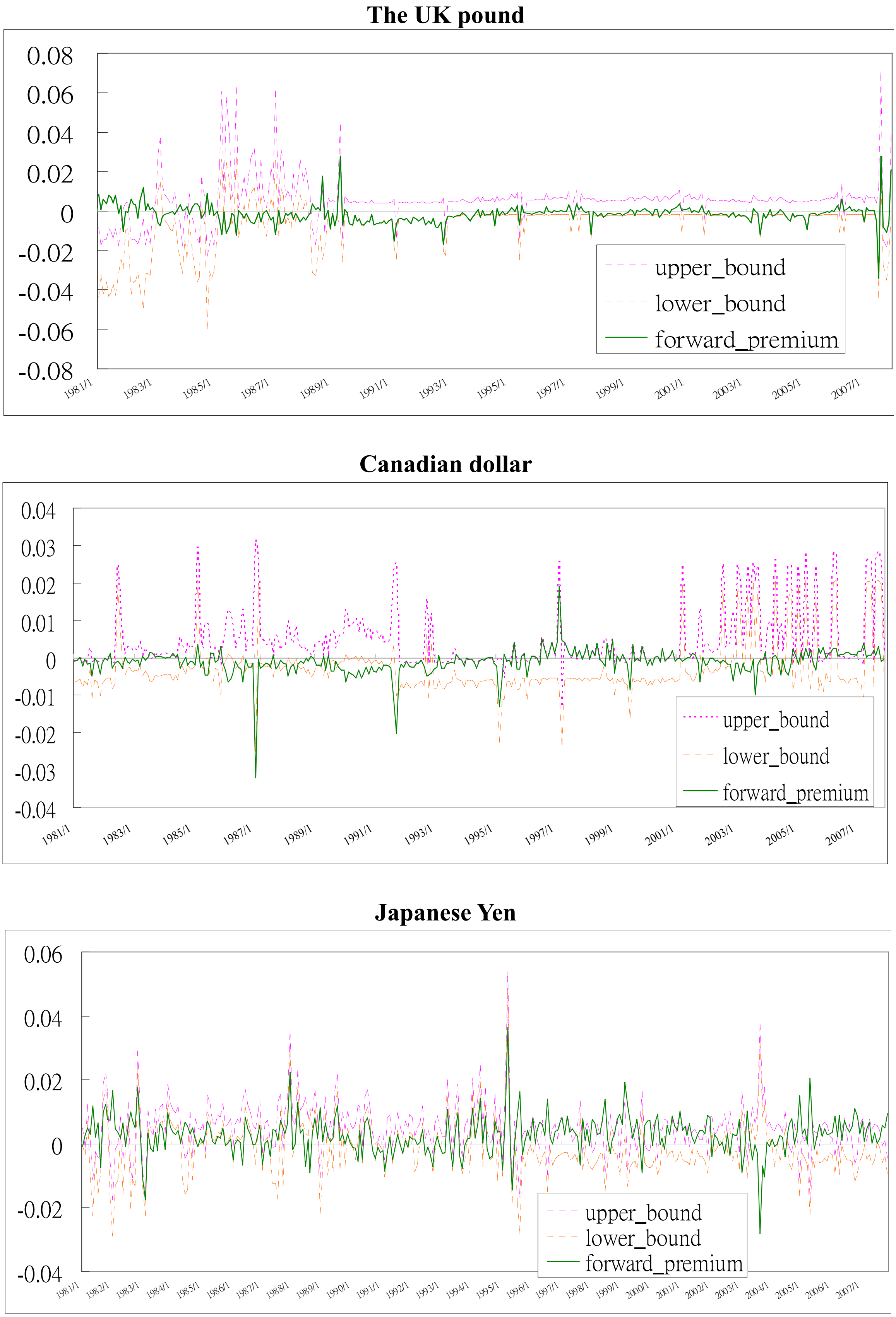

Figure 2: Actual forward premiums and the TV-STR 95% CI for spot rate changes

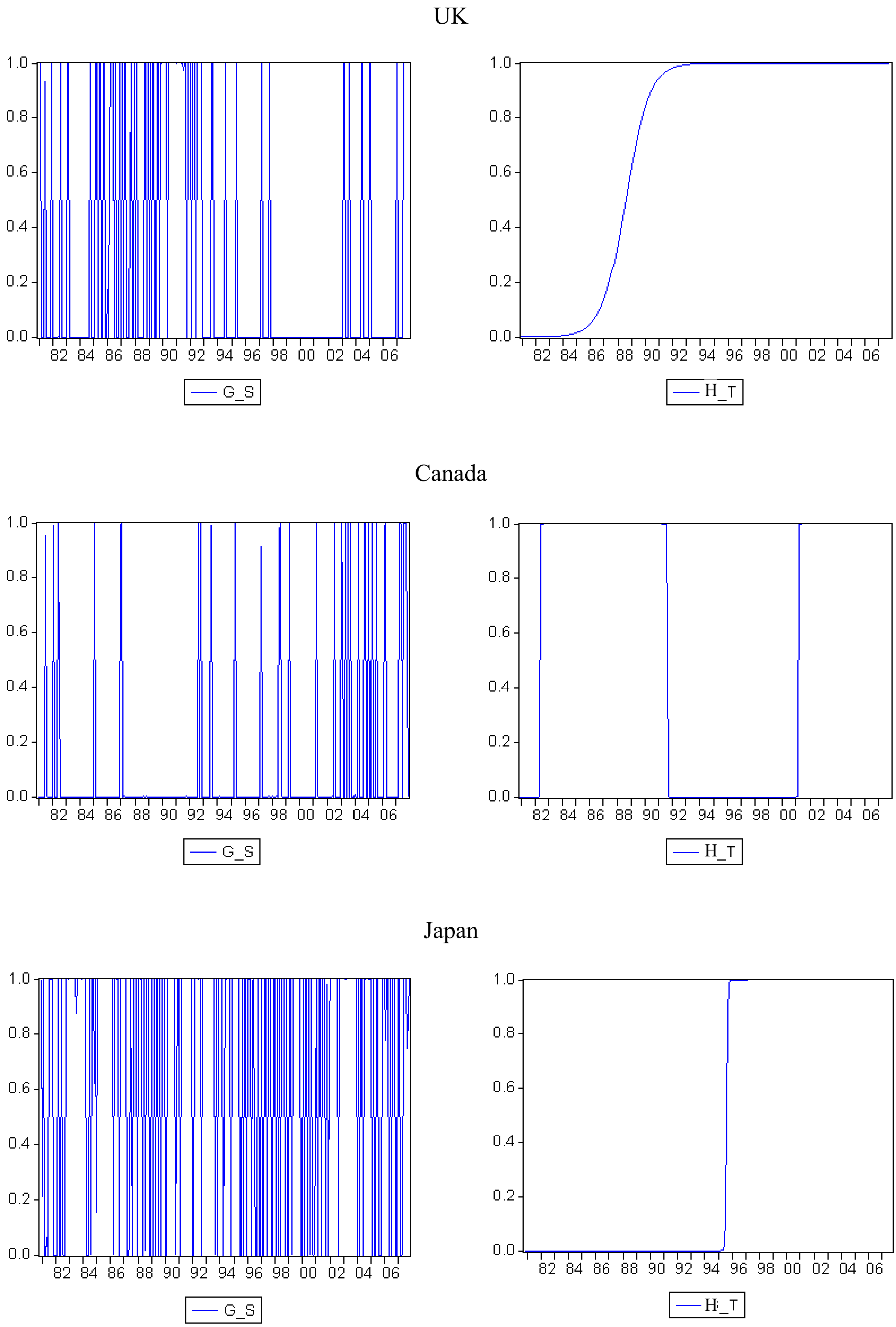

Figure 3. The movements of transition functions, $G(\cdot)$ and $H(\cdot)$, along with time line