The Forward Premium Puzzle and Transactions Costs

Yoshihiro Kitamura+ Ayano Sato+ Hiroya Akiba++

No. 0302

+Graduate School of Economics

++School of Political Science & Economics
Waseda University
Tokyo, Japan

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ABSTRACT

This paper presents new interpretation regarding the forward premium puzzle (FPP). We postulate that the negative correlation between the expected rate of depreciation and interest differential is due to transactions costs. We assume that deviations from uncovered interest parity (UIP) follow a random walk process within the transactions costs boundaries, but a mean-reverting process by arbitrage is expected once the deviations exceed the boundaries. Several non-linear processes for UIP residuals are tested for 16 high-income economies. A linear framework is rejected in favor of STAR processes for UIP deviations. This indicates that transactions costs provide an alternative explanation for FPP.

Address correspondence to:
Hiroya AKIBA
School of Political Science & Economics
Waseda University
1-6-1 Nishi-Waseda, Shinjuku-ward
Tokyo 169-8050, JAPAN
Phone & FAX: +81-3-5286-1767
E-mail: akiba@waseda.jp

Introduction[*]

This article focuses on one central question of whether transactions costs can account for the well-known empirical failure of the forward premium as an unbiased predictor of future exchange rate movements. This failure has been one of the unresolved issues in the forefront of modern international finance, known as the forward premium puzzle (FPP hereafter).[1] To be more specific, it refers to the empirical finding that the forward premium is not only a biased predictor, but also generally predicts the future spot exchange rate in the wrong direction.

The puzzle has been challenged in line with two main explanations: time varying risk premium and expectations errors. However, the puzzle still remains unresolved, and thus serious, in the sense that none of the subsequent empirical studies, using elaborated estimation techniques and different data sets, can account for the puzzle (Froot and Thaler, 1990; Wu and Zhang; 1996, Sarno and Taylor, 2002).

This empirical failure has led to two alternative, but not exclusive, approaches to the puzzle. One is an open economy macroeconomics approach to construct a plausible model that gives rise to the two characteristics of the FPP. One is that the risk premium is more volatile than the expected depreciation, and the other is that their covariation is negative (Fama, 1984; Engel, 1996; Obstfeld and Rogoff, 1996). Specifically, a consumption and/or money based general equilibrium model, or an intertemporal CAPM have been employed to examine the characteristics of the risk premium. However, most of the empirical studies either rejected the consumption Euler equations, or gave rise to implausibly large coefficient values of the relative risk aversion.[2] [3] A slightly more straightforward model-based explanation of the FPP has also

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[1] For an earlier survey on the FPP, see Hodrick (1987). For a recent comprehensive survey, see Engel (1996). For various unresolved puzzles in international finance, see Lewis (1995). For a brief exposition that the FPP contains a wide range of serious problems in international finance, see Obstfeld and Rogoff (1996, chapter 8).

[2] In this sense the FPP resembles, but is slightly more complicated than, the familiar equity premium puzzle by Mehra and Prescott (1985) (see Engel, 1996).

[3] See a comprehensive survey in Engel (1996, sections 3.1 and 3.2). Most of the studies that analyzed the FPP within a general equilibrium model employed a two-country model by Lucas

been offered by deriving the UIP condition from a rational expectations macro model that embodies monetary policy endogeneity (McCallum, 1994; Meledith and Ma, 2002). But none of these successfully accounts for the FPP in the short-run.

Another strand of tests has elaborated on estimations for the FPP by using new techniques and data sets. Several interesting attempts have been made along this line. Recent studies include, for example, consideration of structural changes (Wu, 1997), examination of the term structure of the risk premia (Clarida and Taylor, 1997) and interest rates (Bansal, 1997), non-parametric tests (Wu and Zhang, 1997), tests by the fully modified OLS (Goodhart, McMahon, and Ngama, 1997), tests by co-integration (Luintel and Paudyal, 1998), uses of the EMS data set (Flood and Rose, 1996) or a high-frequency, cross-country data set (Flood and Rose, 2002), and the random time effects panel data (Huisman, Koedijk, Kool, and Nissen, 1998) etc. Unfortunately, none of these could successfully account for the FPP, although some improvements have been made. A general conclusion is thus, that the FPP continues to remain unresolved.

Baillie and Bollerslev (2000) suggest that, by examining the monthly Deutsche mark (against the U.S.dollar) exchange rate for 1974-1991, the FPP is attributable to the slow speed of adjustment in the foreign exchange markets, and thus is a short-run phenomenon that will disappear in the longer time horizon. McDonald and Taylor (2001) conclude that, examining the monthly exchange rates of the EMS currencies against the Deutsche mark and the U.S. dollar for 1978-1994 by a vector error-correction model, the FPP is due to the forward exchange rates that are usually not weakly exogenous. Bansal and Dahlquist (2000, B-D hereafter), using monthly observations of 28 economies for the period 1976-1998, found that the FPP is confined only to high income economies, and that the forward premium responds asymmetrically to the interest rate differentials.

In view of these unsuccessful developments, further research is necessary in four general directions as mentioned by Engel (1996) to explain why the FPP occurs. They are extension of the risk premium analysis, the Peso problem, expectations using survey data, and inefficiency in the international financial markets arising from various frictions. This paper takes up the last direction, an analysis of frictions that include short-sale, borrowing, solvency constraints, and

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(1982) (see, e.g. Hodrick and Srivastava, 1984; Bekaert, 1996).

transactions costs. We focus on transactions costs that have attracted attention in the FPP literature (Cornell, 1989; Bekaert and Hodrick, 1993). Transactions costs are defined as the bid-ask spreads in foreign exchange dealings, as the spreads between bidding and asking prices constitutes costs incurred by investors in the foreign exchange markets (Hartmann, 1998). In fact, Curcio, Goodhart, Guillaume and Payne (1997, 2000) demonstrated that the foreign exchange markets are efficient in the sense that there are no excess profits from the foreign exchange dealings for the British pound, Deutsche mark, and Japanese yen against the U.S. dollar after adjusting transactions costs.

The purpose of this paper is to further examine two specific problems clarified in B-D (2000) by transactions costs. First of all, we examine whether transactions costs account for the FPP in 16 high income economies where B-D (2000) have demonstrated that the FPP is observed. We apply techniques employed by Michael, Nobay and Peel (1997) for analyzing transactions costs for deviations from the purchasing-power parity (PPP). The selected technique differs from theirs, because responses to the FPP could be asymmetric, while those to deviations from PPP would be symmetric.

Secondly, B-D (2000) have also demonstrated that the changes in the exchange rates respond asymmetrically to the interest rate differentials. Specifically, they showed that the FPP exists only in states when the U.S. interest rate exceeds the foreign rates. This empirical finding is, with the presumption that covered interest rate parity holds as a well-documented empirical regularity, equivalent to another empirical fact confirmed by Wu and Zhang (1996). We examine whether these asymmetric changes in the exchange rate are explained by transactions costs.

The rest of the paper is organized as follows: The next section briefly discusses the FPP. Section 3 is devoted to an empirical examination of the FPP. Our strategy for empirical analysis is to first assume that the residuals from the UIP condition reflect transactions costs in the foreign exchange markets, and then to examine whether the costs are approximated, and hence explained, by the Smooth Transition Auto Regressive (STAR hereafter) models. It is shown that our STAR models successfully account for the FPP in 15 of the 16 economies in which the FPP was detected by B-D (2000). Section 4 concludes the paper.

1 The Forward Premium Puzzle

This section briefly summarizes the forward premium puzzle, or FPP. If the exchange market uses all information available up to the present, the forward exchange rate at time $t+1$, F_t (in log form) contracted at t, is an unbiased predictor of the future spot exchange rate S_{t+1} (in log form). This so-called unbiasedness hypothesis is expressed in the following equation:

$$E\,[S_{t+1} \mid I_t] = F_t \qquad (1)$$

where E represents the mathematical expectations and I_t the information set available at time t. Subtracting the spot rate S_t (in log form) from both sides of equation (1) and transforming it into a regression equation yields:

$$S_{t+1} - S_t = \alpha + \beta\,(F_t - S_t) + u_{t+1} \qquad (2)$$

which has been empirically examined in the literature.[4] u_{t+1} denotes disturbance, which is a white noise error term with a zero mean and finite variance, and uncorrelated with information available at time t according to the efficient market hypothesis. F_t-S_t is called the forward premium. Equation (2) tells us whether the forward premium has the predictive power of future changes in the spot rate. The unbiasedness hypothesis maintains that the coefficients α and β should be equal to zero and one, respectively.

We assume, based on accumulated evidence (*e.g.*, Frankel and MacArthur, 1988), that the covered interest parity condition (CIP hereafter) holds, and this implies that the forward premium is equivalent to the interest rate differential. Thus we formally write equation (2) as follows:

$$S_{t+1} - S_t = \alpha + \beta\,(i_t - i_t^*) + u_{t+1} \qquad (3)$$

where i_t and i_t^* are the domestic and the foreign interest rates observed at t. Equation (3) is the so-called uncovered interest parity condition (UIP hereafter) in a regression form. The null hypothesis of UIP embodies the joint hypotheses of rational expectations, perfect capital movements, and perfect substitution among assets. The hypothesis still maintains that $\alpha = 0$ and $\beta = 1$. However, many preceding studies show that these joint hypotheses are strongly rejected. To be more specific, the estimate of the slope coefficient β is not only significantly

[4] Equation (1) is based on the Efficient Market Hypothesis that applies rational expectation hypothesis to the price formation of foreign exchange markets.

different from the theoretically prescribed value of unity, but also negative.[5] This negative empirical correlation between the interest rate differential (or the forward premium) and the expected rate of depreciation is puzzling, and thus has been called the forward premium puzzle (FPP).

2 The Empirical Analysis

2.1 The Data

We extend the analysis of FPP with the data of same 16 developed countries employed by B-D (2000): Switzerland, Hong Kong,[6] Singapore, Japan, Belgium, Austria, Denmark, Canada, France, Germany, the Netherlands, Italy, the U.K., Australia, Sweden, and Spain. We use monthly data of the spot exchange rates and the interest rates from the IFS CD-ROM (the 2001 edition) compiled by the International Monetary Fund (IMF), except for the interest rates of Hong Kong.

In order to make our results comparable with B-D (2000), the sample period for the 15 countries, excluding Hong Kong, is from January 1976 to December 1998. The sample period for Hong Kong is from January 1982 to December 1998.[7] The maximum number of observations is 275. The interest rate used for 14 countries, excluding France and Hong Kong, is the end-of-period values of the money market rate. Due to missing observations in the interest rate data for France and Hong Kong, we use the end-of-period values of the Treasury bill rate from the IFS for France and, for Hong Kong, the overnight interbank rate from the Hong Kong Monetary Authority web site.

The IFS exchange rate data consist of end-of-period values of the market rate, defined by the home currency price per unit of the US dollar.

2.2 The OLS Result

[5] Froot and Thaler (1999) show that the average value of β across 75 previously published papers is -0.88.
[6] To be precise, Hong Kong is not a country, but we regard it as such.
[7] The interest rates before 1982 in Hong Kong are not available.

Equation (3) is estimated by the ordinary least squares with the US interest rate as i^*. Although some empirical studies on the FPP interpreted the estimated value of α as a risk premium,[8] we interpret it as the adjustment coefficient.[9] The reason for it is that we use the money market rate from the IFS, where there exist differences in the term-structure among them in each country.[10]

The estimated results are reported in Table 1. The regression coefficient β is fairly close to the theoretically prescribed value of unity only for Italy and Sweden, and it is statistically significant at the 5% level for Sweden. Therefore, we hesitate to conclude that the FPP is observed for these two countries. However, estimates of β is negative for the remaining 14 countries, and it is even significant at the 1% level for Japan and Canada, the 5% for UK, and the 10% for Switzerland and Hong Kong, respectively. Standard errors of the estimated coefficients are between 0.010 (Hong Kong) and 0.037 (Switzerland), which are relatively small.

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Insert Table 1 around here
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In short, our results confirm that the FPP is observed for 14 out of the total 16 countries, which is quite similar to the conclusions reached in B-D (2000). However, since the coefficient of determination is extremely low, the overall reliability of the regression results are low. This is confirmed visually in a scatter diagram (see Figure 1), in which we plot the depreciation rate of exchange rate along the vertical axis and the interest rate differential along the horizontal axis. As observed in the figure, most of the observations concentrate on the origin, implying that there seems to be no statistical correlation between the two variables. Thus, it is difficult, if not impossible, to infer the alternative hypothesis against the traditional UIP theory.

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[8] For instance, Baillie and Bollerslev (2000) or Meredith and Ma (2002).

[9] In the IFS data the money market rate is regarded as the short-term interbank rate, but it is unclear whether it is the rate of interest that has the same term-structure for all sample countries and periods. Therefore, there remains a possibility of serial correlation caused by a moving average process due to overlapping observations. In other words, since the observation is monthly, the interest rate should have the maturity of one month for the consistency of our estimation. Despite these possible shortcomings, we use the IFS data because of the high reliability and accuracy for using the same source of data.

[10] See Chen and Wu (2000), regarding the constant term in the regression equation as the

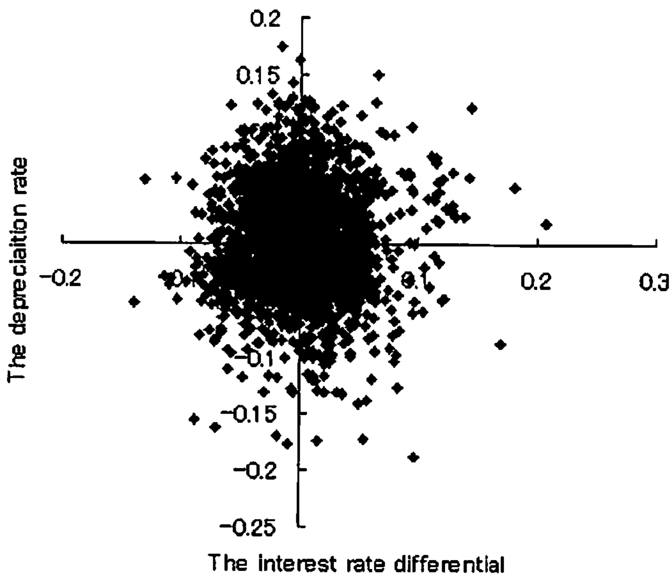

Figure 1: Interest rates and the depreciation rates of US dollar exchange rate

We then examine the validity of the restriction $\beta = 1$ imposed on equation (3) by the Wald Test (Table 1). From the table, the null hypothesis cannot be rejected for 5 countries at the significant level of 5%: Belgium, France, Germany, Italy, and Sweden. This suggests that UIP would possibly hold only in the one-third of the 16 countries. However, we proceed with our analysis, imposing the restriction $\beta = 1$ on equation (3),[11] because of the following reason. We put forth our hypothesis in the present investigation that, once transactions costs are appropriately considered, UIP should be vindicated statistically.

2.3 Estimation of STAR model

Testing of Transactions costs with STAR model

In their analysis of deviations from PPP, Michael, Nobay and Peel (1997) (MNP hereafter) consider the existence of transactions costs as a possible candidate for persistent deviations. MNP estimate a Smooth Transition Auto Regressive (STAR) Model, and show that the

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adjustment coefficient.

[11] Goodhart, McMahon and Ngama(1997) show that the estimate of β is 0.98 and close to 1. This

residuals of the PPP equation follow a nonlinear stochastic process.[12] [13]

Using monthly and annual data for the United States, the United Kingdom, France and Germany, MNP empirically show that transactions costs considered by Dumas (1992) successfully account for persistent deviations from PPP, because commodity trade is not frictionless. We apply the same idea as that used by MNP for analyzing deviations from PPP to disturbance u of equation (3). To be more specific, it is postulated that deviations from UIP follow a nonlinear process that is mean-reverting, with the speed of adjustment toward UIP varying directly with the extent of the deviation from UIP. As a result of the costs of foreign exchange trading, persistent deviations from UIP are left uncorrected as long as they are small relative to the cost of trading.

Let us define the transactions costs band, within which no arbitrage trading takes place. Within the band deviations from UIP follow a random walk process, so that exchange rate depreciation (left-hand side of equation (3)) spends most of the time away from the interest rate differential (right-hand side of equation (3)). When exchange rate change deviates from an interest rate differential, and the deviation exceeds the transactions costs band, the process for residual u is mean-reverting (see Figure 2) because of arbitrage trading. Thus, our strategy implies that if an estimated STAR model captures the motion of residuals statistically, the FPP is successfully accounted for by the transactions costs.

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result is consistent with our restriction $\beta=1$.

[12] See Terasvirta and Anderson (1992), Granger and Terasvirta (1993).

[13] STAR models have been applied to a number of time series analyses. See, *e.g.* Leybourne and Mizen (1997) for consumer prices, Lutkephl, Terasvirta and Wolters (1999) for money demand function, and Sarantis (1999), Taylor and Peel (1999) for exchange rates. But as far as we are aware, there has been no application of STAR models to UIP.

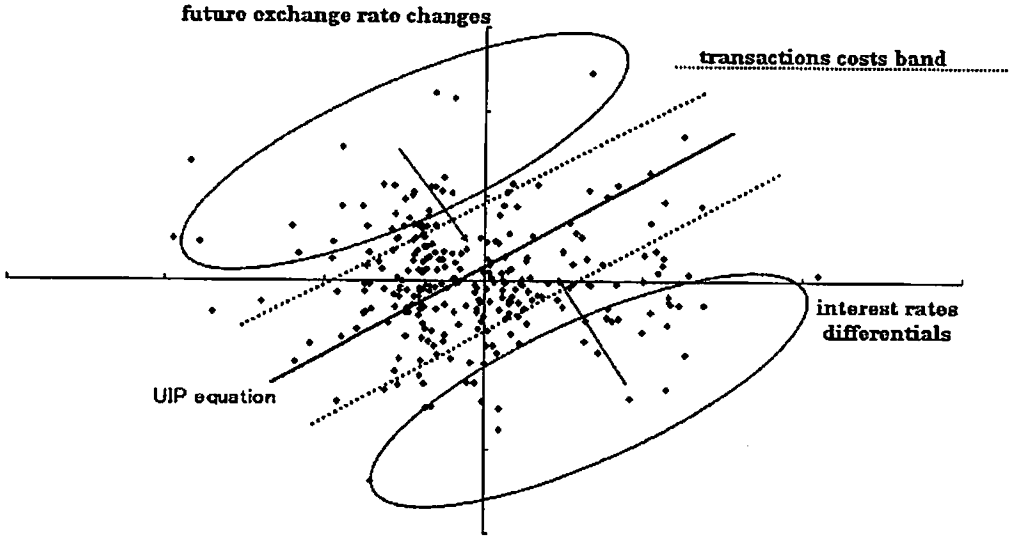

Figure 2: UIP and the transactions costs band

We assume that disturbance u is described by the following STAR model:

$$\Delta u_t = \lambda u_{t-1} + \sum_{j=1}^{p-1} \phi_j \Delta u_{t-j} + (\lambda^* u_{t-1} + \sum_{j=1}^{p-1} \phi_j^* \Delta u_{t-j}) F(u_{t-d}) + \xi_t \qquad (4)$$

where d is a delay parameter, and nonlinear function $F(u_{t-d})$ is assumed to be of two types. The first one is an asymmetrical logistic function:[14]

$$F(u_{t-d}) = \{1 + \exp[-\gamma(u_{t-d} - c)]\}^{-1} \quad \gamma > 0, \quad 0 \le F(u_{t-d}) \le 1 \qquad (5)$$

when $F(u_{t-d})$ is an asymmetric logistic function, the larger the deviation from c of u_{t-d} positively beyond the upper band of the transactions costs, the closer $F(u_{t-d})$ will be to 1. In this case, equation (4) becomes a linear model:

$$\Delta u_t = (\lambda + \lambda^*) u_{t-1} + \sum_{j=1}^{p-1} (\phi_j + \phi_j^*) \Delta u_{t-j} + \xi_t'$$

The necessary condition for deviations to converge into the band is $\lambda + \lambda^* < 0$ (u_t diverges

[14] Our discussion on the effect of the transactions costs suggests that the larger the deviation in equation (3) is, the stronger is the tendency to move back to the transactions costs band. γ measures the speed of transition for changing deviations, and c indicates the half-way point between the upper and the lower bounds of the transactions costs band.

for $\lambda + \lambda^{*} \geq 0$). Furthermore, as the magnitude of the negative deviation of u_{t-d} from c increases beyond the lower band of the transactions costs, $F(u_{t-d})$ approaches to 0. In this case, equation (4) becomes the following linear model:

$$\Delta u_t = \lambda u_{t-1} + \sum_{j=1}^{p-1} (\phi_{j+}\phi_j^{*}) \Delta u_{t-j} + \xi_t''$$

where $\lambda < 0$ is the necessary condition for deviations to converge into the band.

The second type of $F(u_{t-d})$ is a symmetrical exponential function:

$$F(u_{t-d}) = \{1 - \exp[-\gamma(u_{t-d} - c)^2]\} \quad \gamma > 0, \quad 0 \leq F(u_{t-d}) \leq 1 \qquad (6)$$

In this case, as the deviation of u_{t-d} from c approaches to positive or negative infinity, $F(u_{t-d})$ approaches to 1. Then, the equation (4) becomes a linear model:

$$\Delta u_t = (\lambda + \lambda^{*}) u_{t-1} + \sum_{j=1}^{p-1} (\phi_{j+}\phi_j^{*}) \Delta u_{t-j} + \xi_t'''$$

For exponential $F(u_{t-d})$, the necessary condition for deviations to converge into the band is $\lambda + \lambda^{*} < 0$ (it diverges for $\lambda + \lambda^{*} \geq 0$).

In summary, $\lambda + \lambda^{*} < 0$ is the necessary condition for positive deviations to converge into the band and $\lambda < 0$ is the necessary condition for negative deviations to converge into the band for LSTAR. For ESTAR, we must have $\lambda + \lambda^{*} < 0$.

Estimation of STAR model

For estimation we first need to specify a linear autoregressive model AR (p), where p denotes the order of lag. The order is selected as p=1 for all countries except Hong Kong, for which the order is 4. [15]

Next, for each linear AR (p) model, we test linearity for different values of the delay parameter d, using the following equation (8). In order to specify d, the test is carried out for the range of values $1 \leq d \leq D$, where D is selected arbitrarily (see MNP). If linearity is rejected for more than one value of d, then d is determined as $\hat{d} = \arg\min p(d)$ for

[15] If the order is determined by the AIC for a finite sample, it is known that it has a positive probability of overfitting, but it is not consistent for a large sample. In contrast, it is known that the SBIC is consistent. Since the sample size of our data is relatively large, we adopt the SBIC. We also check the validity of the selected order of lag with the Ljung-Box (Q) statistics.

$1 \leq d \leq D$ where $p(d)$ is the p-value of the selected test (Terasvirta and Anderson, 1992: p.122). For our monthly data we consider $1 \leq d \leq 3$ as plausible values for the delay parameter (see Table 2). As observed from the table, the null hypothesis that residuals of the equation (3) follow a linear stochastic process is rejected at $\hat{d} = \arg\min p(d)$ for 15 countries, excepting the UK, for all $1 \leq d \leq 3$. With the exception of the UK, we conclude that residuals of the equation (3) follow a nonlinear stochastic process at the 10% level of significance for two countries, Switzerland and France, and at the 5% level for the remaining countries. With $\hat{d} = \arg\min p(d)$ for each country, we chose between logistic STAR (LSTAR hereafter) and exponential STAR (ESTAR hereafter) models for each country except the UK. Since the LSTAR model implies UIP as the threshold for limiting cases of positive or negative deviations, the transactions costs band (defined with ESTAR by MNP) is not well-defined. However, it is one of the salient features of the LSTAR model that it can allow for asymmetric adjustment.

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Insert Table 2 around here

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The artificial regression of equation (4) is as follows:

$$u_t = \beta_{00} + \sum_{j=1}^{p} (\beta_{1j} u_{t-j} + \beta_{2j} u_{t-j} u_{t-d} + \beta_{3j} u_{t-j} u_{t-d}^2 + \beta_{4j} u_{t-j} u_{t-d}^3) + \varepsilon_t' \qquad (8)$$

where d is determined for each country by $\hat{d} = \arg\min p(d)$ in Table 2. $\beta_{00} = 0$ is expected to hold since u_t is the deviation from the UIP. The model selection method advocated by Terasvirta (1994, p.212) is as follows.[16]

$$H_{03}: \ \beta_{4j} = 0$$
$$H_{02}: \ \beta_{3j} = 0 \quad given\, H_{03}$$
$$H_{01}: \ \beta_{2j} = 0 \quad given\, H_{03}\, and\, H_{02}$$

H_{03} reject,	$\Rightarrow$	LSTAR				
H_{03} accept,	H_{02} reject	$\Rightarrow$	ESTAR			
H_{03} accept,	H_{02} accept,			H_{01} reject	$\Rightarrow$	LSTAR

These hypotheses are tested with the Lagrange Multiplier (LM hereafter) statistics (Terasvirta

[16] Alternatively, Sarantis (1999, p.33) suggests that "one should compute the P-values for all F-tests...and make the choice of the STAR model on the basis of the lowest P-values". For more details, see Terasvirta (1994, p.212).

and Anderson, 1992). F_{03}, F_{02}, and F_{01} are the LM test statistics for H_{03}, H_{02}, and H_{01}, respectively (Table 3). It is concluded from the table that, for residuals of equation (3) for four countries (Switzerland, Japan, Denmark, and France), H_{02} is rejected at the 5 % level of significance, and thus the ESTAR models were chosen. For 2 countries, Canada and Australia, H_{01} is rejected and LSTAR models were chosen. For the other 9 countries, H_{03} is rejected at the 5 % level of significance and the LSTAR models were chosen.

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Insert Table 3 around here
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Estimation Results

The estimation results of the STAR models are reported in Table 4.[17] Also shown is the ratio of the residual variance of a nonlinear model to that of the corresponding AR model chosen by the SBIC. The smaller the ratio is, the higher the explanatory power of the nonlinear model is (Terasvirta and Anderson, 1992).[18] In addition, Jarque-Bera test, Q statistics and ARCH statistics are reported in Table 4. With the exception of one country (the UK), it is shown that deviations from equation (3) follow a nonlinear process that is mean-reverting, with the speed of adjustment toward UIP varying directly with the extent of deviation from equation (3) for 15 countries.

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Insert Table 4 around here
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Table 5 reports the test of the hypotheses for transactions costs band defined by MNP. For 4 countries for which ESTAR models were chosen, the likelihood ratio statistics provide support for the presence of transactions costs for Switzerland and France. The countries for which

[17] The estimation result for Hong Kong with the order of lag being 4 is as follows:

$$u_t = .147(.001) - 3.908(.372)u_{t-1} + 1.640(.324)\Delta u_{t-1} - 1.116(.446)\Delta u_{t-2} - 3.060(.327)\Delta u_{t-3}$$
$$+3.732(.538)u_{t-1} - 1.947(.482)\Delta u_{t-1} + .845(.493)\Delta u_{t-2} + 3.011(.348)\Delta u_{t-3}]$$
$$\times\{1+exp\,[-1.713(.184)\times 90.909(u_{t-1} - 0.025(0.001))]\}^{-1}$$

(The numbers in parenthesis are standard errors.)

[18] For the ratio of the residual variances, the best ratio is 0.766 and the worst is 0.967 in MNP, 0.883 and 0.962 in Chen and Wu (2000), respectively.

LSTAR were chosen indicate that the tendency of converging into transactions costs band is asymmetrical. A plausible interpretation for the asymmetry is that reaction by international investors to appreciation or depreciation is asymmetrical, depending on whether the home currency appreciates or depreciates more than they expected. This asymmetrical reaction in turn leads to the difference in an adjustment speed into the transactions costs band.

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Several suggestive preceding studies are found in the literature on the difference in investors' behavior. In Huisman, Koedijk, Kool, and Nissen (1998), it was observed that the relationship between the forward premium and the future change in the spot exchange rate, as suggested by UIP, tends to be particularly strong in periods with large forward premiums. Wu and Zhang (1996), using the Deutschmark and the Japanese yen exchange rates against the US dollar for the period of March 1973 to May 1993, show that UIP holds in periods when the forward US dollar is quoted at a premium, but fails when it is quoted at a discount. The asymmetrical results observed by Wu and Zhang are reconfirmed with empirical results of Bansal (1997) and Bansal and Dalhquist (2000). A similar but slightly different interpretation is also possible. The speculators' asymmetrical behavior leads us to suggest that when dollar appreciation is expected, the speed of the adjustment of the exchange rate by speculator's dollar purchase is immediate, but its speed is slower when dollar depreciation is expected. This asymmetrical speed may be due to the imperfection of the international capital market, where the transactions costs for dollar purchase are smaller than for dollar sale. Moreover, Uesugi and Yamashiro (2002) find that: (1) asymmetric predictability of forward interest rates results from the sign of the forward premium and suggest that a time-varying term premium possibly explains this asymmetry. And, (2) another asymmetric predictability occurs between periods right after a monetary policy change and the following period, which leads up to the next policy change. These interpretations must be carefully examined in our future research.[19]

Our empirical results suggest that UIP holds when regression residuals follow a

[19] It seems to us that these findings are consistent with non-linearity found in recent empirical studies based on the microstructure approach. See, *e.g.*, Evans and Lyons (2001) and Lyons (2002). We are grateful to Ales Bulir for calling the literature to our attention.

mean-reverting process. In addition, whereas the preceding studies focus on the relationship between the expected rate of depreciation and the forward premium, our empirical results differ from those studies in that we focus on transactions costs.

Conclusions

This paper examined the forward premium puzzle, which is one of the most important unresolved puzzles in the modern theory of international finance, by focusing on transactions costs. Following a suggestion by Obstfeld and Rogoff (2000, p.383), we employ an idea put forward by Dumas (1992) and a technique by MNP for our analysis. We also use the LSTAR model, because of empirically observed asymmetric non-linearity in the FPP by Wu and Zhang (1996) and Huisman, Koedijk, Kool, and Nissen (1998).

Out of a total of 16 countries where the FPP has been observed in a recent study by B-D (2000), ESTAR models were chosen for 4 countries and LSTAR models were chosen for 11 countries. These results mean that the deviations from UIP follow a nonlinear process that is mean-reverting, with variable speed of adjustment toward UIP, and that the realized exchange rates return within the transactions costs band rapidly and converge toward UIP as deviations are further away from UIP. Our results also mean that the UIP condition does not necessarily hold within the transactions costs band, in particular where transactions costs are relatively small (Switzerland and France). Furthermore, our empirical results suggest that the transactions costs are a possible cause of FPP in our 11 sample countries for which asymmetric adjustment speed towards UIP is observed.

It should be mentioned that, for all countries that the LSTAR model was chosen, an asymmetric adjustment process is suggested. This gives rise to at least two further problems to be addressed in future. One is that whether the speed of adjustment differs between inside and outside of the transactions costs band appropriately defined. Furthermore, we should deliberately consider from where such non-linearity comes from. It has been suggested that the reaction by international investors is asymmetrical to dollar appreciation or depreciation. This seems to be a promising avenue to the answer and to confirm the results of preceding studies.

Still certain other issues remain to be addressed. First, the CIP condition does not necessarily hold for all observed countries and periods,[20] although we assumed that it does. Therefore, in order to properly examine the existence of transactions costs as a determinant of FPP, the forward premium itself should be used rather than the interest rate differential. Second, we assume that deviations from UIP represent only transactions costs. However, as many preceding studies point out, the deviations may include various elements, such as the deviation generated from the failure of rational expectations and the risk premium. Finally, we neglect to consider the possibility of central bank's intervention. Krugman (1991) shows exchange rates reverse within a band by intervention. If the exchange rate markets for our sample countries have been frequently intervened, we would have different results.

It is interesting to note that Obstfeld and Rogoff (2000, p.351) argue that trade in securities is "free and costless" (see also Obstfeld and Rogoff, 1996, p.586). Nonetheless, even without taking the above issues into consideration, the results of this study suggest that the existence of transactions costs should be treated properly in an analysis of the determinants of the forward premium puzzles.

[20] See Ito (1986) concerning the factors leading to deviation from the CIP condition.

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no	country name	α	β	R_2	D.W	S.E	Wald statistics
1	SWITZERLAND	-0.005*	-1.007*	0.010	1.903	0.037	11.456***
		(-0.003)	(-0.593)				
2	HONG KONG	0.001*	-0.780*	0.017	1.685	0.010	18.464***
		(0.000)	(0.414)				
3	SINGAPORE	-0.002	-0.983	0.005	1.871	0.016	5.254**
		(0.001)	(0.918)				
4	JAPAN	-0.009***	-2.508***	0.029	1.904	0.034	16.480***
		(0.002)	(0.864)				
5	BELGIUM	0.000	-0.526	0.001	1.972	0.033	3.661*
		(0.002)	(0.797)				
6	AUSTRIA	-0.002	-0.525	0.001	1.997	0.032	4.014**
		(0.002)	(0.761)				
7	DENMARK	0.000	-0.205	0.000	2.005	0.032	4.040**
		(0.002)	(0.599)				
8	CANADA	0.002***	-1.281***	0.025	2.190	0.013	22.312***
		(0.000)	(0.483)				
9	FRANCE	0.001	-0.534	0.001	2.010	0.032	3.213*
		(0.002)	(0.856)				
10	GERMANY	-0.002	-0.389	0.000	1.991	0.032	3.344*
		(0.002)	(0.759)				
11	NETHERLANDS	-0.002	-0.896	0.005	2.026	0.033	6.338**
		(0.002)	(0.753)				
12	ITALY	-0.001	1.024	0.008	1.821	0.030	0.001
		(0.003)	(0.689)				
13	UK	0.003	-1.655**	0.015	1.830	0.033	11.326***
		(0.002)	(0.789)				
14	AUSTRALIA	0.003*	-0.472	0.002	2.030	0.029	6.771**
		(0.002)	(0.565)				
15	SWEDEN	0.000	0.946**	0.020	1.877	0.030	0.018
		(0.002)	(0.393)				
16	SPAIN	0.003	-0.176	0.000	1.850	0.033	6.308**
		(0.002)	(0.468)				

Table1: OLS Results and Wald Test

Standard errors appear in parentheses,

***, **, * indicate significance at 1, 5, 10% level, respectively.

no	country name	p	Q[4]	Q[12]	d=1	d=2	d=3
1	SWITZERLAND	1	1.24	12.20	0.242	0.066	0.868
2	HONG KONG	4	30.5**	66.4**	0.000	0.000	0.000
3	SINGAPORE	1	4.41	11.60	0.262	0.003	0.004
4	JAPAN	1	1.26	9.42	0.099	0.010	0.819
5	BELGIUM	1	4.15	14.30	0.000	0.040	0.968
6	AUSTRIA	1	2.03	11.80	0.000	0.129	0.996
7	DENMARK	1	4.49	16.20	0.000	0.061	0.920
8	CANADA	1	.946	20.60	0.025	0.683	0.875
9	FRANCE	1	4.49	11.10	0.066	0.209	0.746
10	GERMANY	1	2.14	11.70	0.000	0.153	0.998
11	NETHERLANDS	1	2.94	11.80	0.000	0.032	0.875
12	ITALY	1	1.96	15.80	0.031	0.230	0.286
13	UK	-	-	-	0.300	0.758	0.193
14	AUSTRALIA	1	8.00	17.70	0.002	0.441	0.770
15	SWEDEN	1	2.12	15.80	0.006	0.315	0.169
16	SPAIN	1	9.17	11.80	0.009	0.454	0.025

Table 2: P-value for the linearity test for various *d*

Q[4] and Q[12] are Q statistics for residual auto-correlation up to order 4 and 12, respectively; ** indicate significance at 5% level.

no	country name	p	d	F03	F02	F01	model
1	SWITZERLAND	1	2	0.596	0.009	0.678	ESTAR
2	HONG KONG	4	1	0.000	0.000	0.000	LSTAR
3	SINGAPORE	1	2	0.006	0.009	0.956	LSTAR
4	JAPAN	1	2	0.091	0.018	0.082	ESTAR
5	BELGIUM	1	1	0.002	0.001	0.404	LSTAR
6	AUSTRIA	1	1	0.002	0.000	0.911	LSTAR
7	DENMARK	1	1	0.064	0.000	0.375	ESTAR
8	CANADA	1	1	0.470	0.104	0.012	LSTAR
9	FRANCE	1	1	0.131	0.028	0.736	ESTAR
10	GERMANY	1	1	0.003	0.001	0.969	LSTAR
11	NETHERLANDS	1	1	0.016	0.000	0.296	LSTAR
12	ITALY	1	1	0.004	0.661	0.613	LSTAR
13	UK	-	-	-	-	-	-
14	AUSTRALIA	1	1	0.091	0.067	0.003	LSTAR
15	SWEDEN	1	1	0.001	0.302	0.850	LSTAR
16	SPAIN	1	1	0.002	0.496	0.164	LSTAR

Table 3: P-value for the model selection tests

The test statistics associated with H_{03}, H_{02}, and H_{01} are denoted by F_{03}, F_{02}, and F_{01}, respectively.

no	country name	λ	λ^*	γ	C	ARCH[1]	ARCH[4]	Q[1]	Q[4]	JB	adj R2	σ 2NL/σ 2L
1	SWITZERLAND	-.444	-.577	5.840×693.021	.016	0.427	1.400	0.885	11.133	14.932***	.468	.978
		(.217)	(.223)	(4.498)	(.004)							
2	HONG KONG	-.704	-1.403	7.465×88.791	.013	12.37***	78.421***	17.953***	64.886***	1498.899***	.847	.538
		(.085)	(.195)	(1.780)	(.000)							
3	SINGAPORE	-.994	.065	3.357×62.778		11.771***	26.707***	3.860	10.586	308.367***	.475	1.003
		(.127)	(.196)	(33.549)								
4	JAPAN	-.687	-.580	.677×776.163		1.088	2.539	3.460	11.326	20.448***	.460	.971
		(.100)	(.239)	(.650)								
5	BELGIUM	-2.716	1.832	2.333×29.554	-.108	0.190	1.202	4.586	18.187	9.441***	.503	.978
		(6.095)	(6.120)	(4.129)	(.107)							
6	AUSTRIA	-.922	-.261	2.411×30.128	.065	3.258*	4.360	2.655	13.396	9.571***	.487	.999
		(.085)	(.300)	(7.866)	(.062)							
7	DENMARK	-.407	-1.037	.296×938.375		2.326	3.349	3.588	11.652	.818	.522	.930
		(.189)	(.228)	(.178)								
8	CANADA	-.903	-.759	1.126×69.843	.047	0.053	2.193	0.714	20.19*	56.635***	.527	.991
		(.107)	(.964)	(1.894)	(.039)							
9	FRANCE	-.562	-.591	.591×955.907	.004	0.378	2.011	3.311	8.633	10.836***	.503	.985
		(.022)	(.014)	(.242)	(.008)							
10	GERMANY	-.919	-.254	2.304×30.111	.062	3.233*	4.558	2.751	13.010	10.238***	.489	1.000
		(.087)	(.289)	(7.536)	(.056)							
11	NETHERLANDS	-3.308	-2.419	1.815×29.543	-.124	0.011	0.982	3.754	15.263	6.983**	.505	.980
		(13.186)	(13.214)	(2.993)	(.180)							
12	ITALY	-.916	-.037	5.693×32.926		1.905	4.569	2.873	15.673	32.113***	.467	1.003
		(.099)	(.129)	(218.405)								
13	AUSTRALIA	-.823	-.273	.688×30.128		0.712	1.014	4.747	14.366	741.459***	.501	.993
		(.248)	(.404)	(1.916)								
14	SWEDEN	-.908	-.052	2.401×32.276		0.068	1.049	2.460	13.377	329.083***	.466	1.003
		(.127)	(.179)	(32.432)								
15	SPAIN	-.806	-.180	2.342×30.184		0.250	0.317	2.265	10.915	304.184***	.457	1.000
		(-.124)	(181)	(34.555)								

Table 4: Estimation Results

Note: The number of parentheses are standard error, σ^{2NL} and σ^{2L} are the residual variance from the nonlinear (ESTAR or LSTAR) and linear models, respectively.

no	country name	*LR_a*		*LR_b*	
1	SWITZERLAND	0.073		0.352	
2	JAPAN	2.480		25.556	***
3	DENMARK	8.198	***	2.514	
4	FRANCE	2.478		0.416	

Table 5: The transactions costs hypotheses.

Note: The test statistics associated with H_0^a *and* H_0^b are denoted by LR_a and LR_b, respectively.

***, **, * indicate significance at 1, 5, 10% level, respectively.

$$H_0^a : 1+\lambda = -\lambda^*, H_0^b : \lambda = 0$$

H_0^a : the deviation outside of the band is white noise.

H_0^b : the deviation inside of the band has a unit root.