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Effect of Net Foreign Assets on Persistency of Time-Varying Risk Premium: Evidence from the Dollar-Yen Exchange Rate

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Abstract

In this paper, I focus on the time-varying and persistent exchange rate risk premiums in uncovered interest rate parity associated with changes in net foreign assets. The results of my analyses of the Dollar-Yen exchange rate provide evidence consistent with my risk premium formulation and the predictability of current account balances. I contend that the strong persistent effect causes nominal exchange rates to appear non-stationary in level. I also argue that the present value model of the level of exchange rates combined with the AR(1) approximation for interest rate differentials can reconcile a failure of uncovered interest rate parity.

JEL classification: F31, F32, F47, G12, G15

Keywords: Uncovered interest rate parity, Time-varying risk premium, Nominal exchange rate stationarity, Current account balance

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1. Introduction

Many researchers have sought to explore the failure of uncovered interest rate parity to hold by investigating time-varying exchange rate risk premiums.[2] Greater uncertainty in forecasting future exchange rates, which may be measured with the volatilities of uncovered interest rate parity regression residuals, may induce increases in risk premiums, which can lead investors to retain foreign currencies. Indeed, some studies have focused on the conditional variance of the regression residuals to investigate the variation in exchange rate risk premiums using ARCH-type models.[3] However, the theoretical relationship between volatilities of asset prices and risk premiums is somewhat vague because standard financial theory asserts that unsystematic risks are not priced. Additionally, in studies that have employed ARCH-type models, the sources of risk premium variation have not been identified. Therefore, attention should be given to the factor(s) of variation in risk premiums.

In this paper, I hypothesize that the exchange rate risk premium in uncovered interest rate parity depends on the amount of net foreign asset holdings. To hold more net foreign assets, risk-averse investors must receive a reward in the form of a larger risk premium associated with exchange rate variations. Therefore, changes in net foreign asset holdings cause exchange rate variations through time-varying exchange rate risk premiums.

[2] Engel (2014) reviewed this topic extensively. See also Alvarez et al. (2009).

[3] For instance, using the CGARCH-M model, which can decompose the permanent and transitory components of regression residual volatility, Li et al. (2012) showed that the time-varying risk premium (particularly, the permanent component) is statistically significant in the uncovered interest rate parity regression residuals. See also Aggarwal (2013).

Cochrane (2011) argued that expected excess returns on risky assets, or risk premiums, are generally predictable and time-varying. Relying on intertemporal budget constraints, Gourinchas and Rey (2007) argued that changes in exchange rates adjust external imbalances, not only through the trade channel, but also through the valuation channel.[4] The authors also stated that deviation of net foreign asset holdings from the trend helps predict future exchange rates. In this paper, I argue that if the amount of net foreign assets affects the risk premium, movements in exchange rates must be predictable.

I further argue that there are persistent effects of changes in net foreign assets, or risky asset holdings generally, on risk premiums. Some researchers have argued that uncovered interest rate parity holds better with a long horizon than with a short horizon.[5] Based on this, if I consider an infinitely long horizon, the uncovered interest rate parity equation becomes a present value model of the level of nominal exchange rates. On the other hand, researchers have come to a consensus that the actual movements of nominal exchange rates are non-stationary in level.[6] In contrast, I argue that the large persistent effects associated with changes in risky asset holdings cause nominal exchange rates to appear non-stationary.

Obstfeld (2006) investigated the Yen exchange rate and the Japanese current account balance (that roughly corresponds to the change of net foreign assets), in a more general manner than studies that used models primarily reliant on intertemporal budget constraints. Through his calibration, Obstfeld (2006) argued that there is a highly significant correlation

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[4] Like Gourinchas and Rey (2007), many studies have focused on the dynamic process of adjusting current account balances and exchange rates under intertemporal budget constraints. For instance, Blanchard et al. (2005) focused on the home bias effect. In this paper, I assume that the investors' aversion to the risk associated with a variation in the foreign exchange rate takes the role of home bias.

[5] See, for example, Chinn and Meredith (2005), Chinn (2006), and Chinn and Quayyum (2012). Despite the abundance of literature supporting this claim, it remains controversial in the empirical literature (e.g. Bekaert et al. (2007)).

[6] Recently, regarding international long-term bond arbitrage, Lustig et al. (2016) have argued in favor of the nominal exchange rate stationarity in level.

between appreciation in the Yen and the Japanese current account surplus. Consistent with Obstfeld's result, I present empirical evidence to show that changes in net foreign assets affect the exchange rate between the US Dollar and the Japanese Yen through the time-varying risk premium in uncovered interest rate parity. This evidence suggests that the Japanese current account balance can predict the Dollar-Yen exchange rate.

To explore this and other issues, I have organized this paper into a series of interrelated sections. In Section 2, I review the theoretical preliminaries for the empirical research on the effect of net foreign asset holdings in uncovered interest rate parity with the time-varying and persistent risk premium. Subsequently, in Section 3, I present the results of my empirical analyses using the exchange rate between the US Dollar and the Japanese Yen. Finally, I offer some concluding remarks in Section 4.

2. Theoretical Preliminaries

In this section, I review the theoretical issues associated with my empirical analyses. For the sake of simplicity, I incorporate only two countries into my analyses—one home country and one foreign country. I express the exchange rate as the price of the foreign currency in units of the home currency.

2.1. Longer-Horizon Uncovered Interest Rate Parity and the Present Value Model of Exchange Rate

I express uncovered interest rate parity (including risk premium) as:

$$E_t\left[\left(1 + R^{\$}_{t\to t+k}\right)\frac{S_{t+k}}{S_t}\right] = \left(1 + E_t[R_{t\to t+k}]\right)\left(1 + \rho_{t\to t+k}\right), \quad (1)$$

where S represents the spot exchange rate, R the nominal interest rate, and ρ the risk premium. Subscripts indicate the dates as usual and the subscript with an arrow (e.g., $t{\rightarrow}t{+}k$) denotes the interval over which the variable is applied. The superscript \$ indicates that the variable is foreign. By taking the logarithm of Eq. (1) and using the lower capital for the logarithm of the variables, the uncovered interest rate parity approximation is expressed as:

$$\begin{aligned} E_t[s_{t+k}] - s_t &= E_t\,[\log S_{t+k}] - \log S_t \\ &= E_t\big[\log(1 + R_{t\rightarrow t+k}) - \log\big(1 + R^{\$}_{t\rightarrow t+k}\big)\big] + \log(1 + \rho_{t\rightarrow t+k}) \\ &\approx E_t\big[R_{t\rightarrow t+k} - R^{\$}_{t\rightarrow t+k}\big] + \rho_{t\rightarrow t+k}. \end{aligned} \tag{2}$$

In many empirical studies that have applied equations similar to Eq. (2) (with or without a term for risk premium), $E_t[s_{t+k}]$ is replaced with $s_{t+k}+\theta_{t+k}$, where θ_{t+k} denotes forecast or measurement error at $t+k$ uncorrelated with the information available at t under the assumption of rational expectations.

It is often said that nominal exchange rates follow random walks and their levels are thus non-stationary. This non-stationarity further suggests that exchange rates are unpredictable. However, if nominal exchange rates follow random walks (unit root processes), the left-hand side of Eq. (2) should be a constant (the drift term). In this case, however, it is difficult to justify that the right-hand side of Eq. (2) is also a constant contrary to the time-varying risk premium. Therefore, if I assume that the process of the logarithm of the spot exchange rate, $\{s_t\}$, is stationary in Eq. (2), it is mean-reverting and, therefore, predictable.

Eq.(2) can take another form:

$$s_t \approx E_t\big[s_{t+k} + R^{\$}_{t\rightarrow t+k} - R_{t\rightarrow t+k}\big] - \rho_{t\rightarrow t+k}. \tag{3}$$

In Eq. (3), k can be made arbitrarily large. I define $X^* = \lim_{k\rightarrow+\infty} X_{t+k}$ and $X^*_t = \lim_{k\rightarrow+\infty} X_{t\rightarrow t+k}$,; iterating Eq. (3) forward, I obtain:

$$s_t \approx E_t\big[s^* + R^{\$*}_t - R^*_t\big] - \rho^*_t. \tag{4}$$

This is equivalent to a present value model of the spot exchange rate.[7] For Eq. (4) to converge, I must assume that $E_t[s^* + R_t^{\$*} - R_t^*] < +\infty$ in addition to assuming that the risk premium is finite. Accordingly, I define the real exchange rate, Q_t, and real interest rate, $V_{t\to t+k}$, such that

$$Q_t = \frac{S_t P_t^{\$}}{P_t},$$

$$1 + R_{t\to t+k} = (1 + V_{t\to t+k})\frac{P_{t+k}}{P_t},$$

where P represents the price level. Using these definitions in the logarithm, Eq. (3) becomes:

$$\begin{aligned} s_t &\approx E_t\left[q_{t+k} + p_{t+k} - p_{t+k}^{\$} + V_{t\to t+k}^{\$} + p_{t+k}^{\$} - p_t^{\$} - V_{t\to t+k} - p_{t+k} + p_t\right] - \rho_{t\to t+k} \\ &= E_t\left[q_{t+k} + V_{t\to t+k}^{\$} - V_{t\to t+k}\right] + p_t - p_t^{\$} - \rho_{t\to t+k}. \end{aligned}$$

Iterating this equation forward, as in Eq. (4), I obtain:

$$s_t \approx E_t\left[q^* + V_t^{\$*} - V_t^*\right] + p_t - p_t^{\$} - \rho_t^*.$$

For the convergence of this present value model of the spot exchange rate, I additionally assume that there is a steady state real spot exchange rate, q^*, and I assume that $\lim_{t\to+\infty}\left(V_t^{\$*} - V_t^*\right) = 0$. The latter condition can be interpreted that the global financial market will eventually be integrated in the real term.

I also assume that there are some risk-free rates for n years and define $k(n)$ to satisfy:

$$1 + R_t^* = \left(1 + R_{t+k(n)}^*\right)\left(1 + I_t(n)\right)^n, \tag{5}$$

where $I(n)$ denotes the annual risk-free rate for n years. Accordingly, taking the logarithm of Eq. (5) allows Eq. (3) to be expressed as:

$$s_t \approx E_t\left[s_{t+k(n)}\right] + n\left(I_t^{\$}(n) - I_t(n)\right) - \rho_{t\to t+k(n)} = E_t\left[s_{t+k(n)}\right] + nD_t(n) - \rho_{t\to t+k(n)}, \tag{6}$$

[7] See Eq. (4) in Engel (2014) for a discrete time model.

where the interest rate differential is defined as $D_t(n) = I_t^{\$}(n) - I_t(n)$. Similarly, Eq. (4) is expressed as:

$$s_t \approx E_t\left[s^* + R_{t+k(n)}^{*\$} - R_{t+k(n)}^*\right] + nD_t(n) - \rho_t^*. \qquad (7)$$

2.2. The Simple Portfolio Model of Foreign Asset Holdings

To get an idea of the effect of risk exposure on exchange rate variation, I consider a simple two-period portfolio problem for the representative home investor rather than deriving the risk premium from an intertemporal consumer problem or consumption growth model.[8] The representative home investor's future wealth, W^*, is:

$$\begin{aligned} W^*(F; S, S^*) &= (1+R)(W - SF) + \left(1 + R^{\$}\right)S^*F \\ &= (1+R)W - \left\{(1+R)S - \left(1+R^{\$}\right)S^*\right\}F, \end{aligned} \qquad (8)$$

where F denotes net foreign assets, which are assets based on the foreign currency, W represents initial wealth, and the symbol "*"[9] indicates future value. The representative home investor's problem is:

$$\max_F E\left[U\left(W^*(F; S, S^*)\right)\right].$$

This problem implies that the representative home investor can choose only the amount of net foreign assets that maximizes his or her expected utility of future wealth, and the risk comes only from the future spot exchange rate variation. I assume that the utility or value function is concave and at least twice differentiable. The utility function has

[8] Using US data, and consistent with Mehra and Prescott (1985), Lustig (2007) found problems with the exchange rate risk premium like those found in equity. Lustig (2007) argued that although consumption growth volatility is too small relative to exchange rate volatility in the consumption-based asset pricing model, the co-variance between consumption growth and the exchange rate remains important.

[9] In Section 2.1, I denote "*" as the steady state for a more general time horizon. In this two-period model, it is reasonable to consider the future period as a steady state because there is only one future period.

non-increasing absolute risk aversion if:

$$A(W^*) = -\frac{U''(W^*)}{U'(W^*)},$$

$$W_2^* > W_1^* \Longrightarrow A(W_2^*) \geq A(W_1^*).$$

Proposition: If the representative home investor's utility function has non-increasing absolute risk aversion, then the increase in net foreign asset holdings makes the spot exchange rate appreciate. That is,

$$\frac{dS}{dF} < 0.$$

Appendix A provides proof of this proposition. The broad types of the utility function often used in economic research have non-increasing or decreasing absolute risk aversion (DARA).

2.3. Time-Varying and Persistent Risk Premium

Let me now turn to the relationship between changes in net foreign asset holdings and risk premiums. I assume that the risk premium takes the following form:

$$\rho_t^* = g_t(F_t) + \omega_t, \tag{9}$$

where ω denotes variation in the risk premiums caused by factors other than changes in net foreign assets. Based on the previous section, the functions g are assumed to increase. The effect of changes in net foreign assets on risk premiums should be persistent under this specification.

I also assume some separability of the net foreign assets in the risk premiums. In Eq.

(9), the functions g are as follows:

$$g_{t+j}(F_{t+j}) = g_{t+j}(F_t + \Delta^F_{t\to t+j}) = \phi_1 g_t(F_t) + h_{t+j}(\Delta^F_{t\to t+j}), \quad (10)$$

where Δ^F denotes the increments in which the net foreign asset changes, ϕ_1 is a positive constant, and the functions h increase with $h(0) = 0$. Eq.s (9) and (10) indicate that the risk premium at $t+j$ is:

$$\begin{aligned}\rho^*_{t+j} &= g_{t+j}(F_{t+j}) + \omega_{t+j} = \phi_1 g_t(F_t) + h_{t+j}(\Delta^F_{t\to t+j}) + \omega_{t+j} \\ &= \phi_1(\rho^*_t - \omega_t) + h_{t+j}(\Delta^F_{t\to t+j}) + \omega_{t+j} = \phi_1 \rho^*_t + h_{t+j}(\Delta^F_{t\to t+j}) - \phi_1\omega_t + \omega_{t+j}. \quad (11)\end{aligned}$$

If the risk premium is stationary, then $\phi_1 < 1$.[10] A backward iteration of Eq. (10) indicates that the risk premium at t is:

$$\rho^*_t = g_t(F_t) + \omega_t = {\phi_1}^t g_0(F_0) + \sum_{l=1}^{t} {\phi_1}^{t-l} h_l(\Delta^F_{l-1\to l}) + \omega_t. \quad (12)$$

From Eq. (12), when ϕ_1 is closer to one, changes in the net foreign assets have a more persistent effect on the risk premium and, thus, on the exchange rate. In this way, ϕ_1 can be seen as a parameter of persistency. When ϕ_1 is less than one but sufficiently large, exchange rates may appear non-stationary in level.

Given this formula for risk premiums, changes in net foreign assets can predict future exchange rates in the sense that expected future exchange rates depend on past changes in net foreign assets. Indeed, from Eq.s (4) and (12), the expected future exchange rates are:

$$\begin{aligned}E_t[s_{t+k}] &\approx E_t\left[E_{t+k}[s^* + R^{\$*}_{t+k} - R^*_{t+k}] - \rho^*_{t+k}\right] \\ &= E_t[s^* + R^{\$*}_{t+k} - R^*_{t+k}] - E_t\left[{\phi_1}^{t+k} g_0(F_0) + \sum_{l=1}^{t+k} {\phi_1}^{t+k-l} h_l(\Delta^F_{l-1\to l}) + \omega_{t+k}\right]\end{aligned}$$

[10] The growth of total wealth in Japan may explain why $\phi_1 < 1$. That said, I do not consider the effect of change in initial wealth in the model from Section 2.2. However, increases in initial wealth induce depreciation in the spot exchange rate under the presumption that Japan has positive net foreign assets during the period in question. See the comment after the proof in Appendix A.

$$= E_t[s^* + R^{\$*}_{t+k} - R^*_{t+k}] - E_t\left[\sum_{l=1}^{k} {\phi_1}^{k-l} h_{t+l}\left(\Delta^F_{t+l-1\to t+l}\right) + \omega_{t+k}\right]$$
$$+{\phi_1}^{t+k} g_0(F_0) + \sum_{l=1}^{t} {\phi_1}^{t+k-l} h_l\left(\Delta^F_{l-1\to l}\right).$$

Thus, the expected future exchange rate, $E_t[s_{t+k}]$, depends on past changes in net foreign assets $\{\Delta^F_{l-1\to l}\}$.

3. Empirical Evidence from the Dollar-Yen Exchange Rate

In this section, I empirically investigate the present value model of exchange rates derived in Section 2.1. Specifically, I focus on the effect of net foreign assets on risk premiums discussed in Sections 2.2 and 2.3. I use data related to the exchange rate between the US Dollar and the Japanese Yen to perform multiple empirical analyses.[11] For the purpose of my analyses, Japan is considered the home country.

3.1. Effect of Net Foreign Assets on Risk Premium and Exchange Rate Levels in Longer-Horizon Interest Rate Parity

By setting j = 1, Eq. (11) becomes:

$$\rho^*_t = \phi_1\rho^*_{t-1} + h_t(\Delta^F_{t-1\to t}) - \phi_1\omega_{t-1} + \omega_t.$$

For the sake of simplicity, I assume that the variation in the risk premium attributable to factors other than net foreign assets is random with a mean of zero. Moreover, for the OLS regression, I also assume this risk premium formula to be linear:

$$\rho^*_t = \phi_1\rho^*_{t-1} + \phi_2\Delta^F_{t-1\to t} + \tilde{\omega}_t, \tag{13}$$

[11] See Appendix B for details related to my data.

where ϕ_2 is a positive constant. The symbol "~" denotes a mean-zero variable, (i.e., $\tilde{\omega}_t = \omega_t - \phi_1 \omega_{t-1}$). Using the risk premium as in Eq. (13) and the iteration of the uncovered interest rate parity as in Eq. (7), I obtain:

$$\begin{aligned} s_t &\approx E_t\left[s^* + R^{*\$}_{t+k(n)} - R^*_{t+k(n)}\right] + nD_t(n) - \phi_1 \rho^*_{t-1} - \phi_2 \Delta^F_{t-1\to t} - \tilde{\omega}_t \\ &= E_t\left[s^* + R^{*\$}_{t+k(n)} - R^*_{t+k(n)}\right] + nD_t(n) \\ &\quad -\phi_1\left(E_{t-1}\left[s^* + R^{*\$}_{t-1+k(n)} - R^*_{t-1+k(n)}\right] + nD_{t-1}(n) - s_{t-1}\right) - \phi_2 \Delta^F_{t-1\to t} - \tilde{\omega}_t \\ &= E_t\left[s^* + R^{*\$}_{t+k(n)} - R^*_{t+k(n)}\right] - \phi_1 E_{t-1}\left[s^* + R^{*\$}_{t-1+k(n)} - R^*_{t-1+k(n)}\right] \\ &\quad +\phi_1 s_{t-1} + n\left(D_t(n) - \phi_1 D_{t-1}(n)\right) - \phi_2 \Delta^F_{t-1\to t} - \tilde{\omega}_t. \end{aligned} \tag{14}$$

Because there are no quarterly data on Dollar-based net foreign assets in Japan, I instead use the normalized quarterly current account balance,[12] C_{t-1}, as a proxy for $\Delta^F_{t-1\to t}$. Note that the current account balance does not include the unrealized gains and losses on risky assets based on the US Dollar (e.g., US stock capital gain). On the other hand, it must include Yen-based trades. There are two points to note about this. First, the share of Yen-based foreign trade in Japan is small. The Japanese Ministry of Economy, Trade, and Industry reported that the share of Yen-based trade is about 35% of exports and 25% of imports in 2002. Second, in Japan, the current account balance is highly correlated with the change in the net foreign assets. The Japanese Ministry of Finance reported annual net foreign asset data based on the IMF Balance of Payments Manual's fifth edition (BPM5), which is based on the Dollar, up to 1994. In my data, the estimated correlation coefficient between changes in the annual net foreign assets and the annual current account balance from 1977 to 1994 is 0.8851.[13]

While I argue, in Section 2.1, that the exchange rate process should be stationary from a theoretical viewpoint, my data do not allow me to reject the null hypothesis of

[12] See Appendix B.

[13] Moreover, according to Obstfeld (2006), the IMF's Coordinated Portfolio Investment Survey 2001 suggested that 75% of Japan's external assets are in non-Yen currency.

non-stationarity associated with the logarithm of the spot exchange.[14] I performed Dickey-Fuller tests with constant terms on each of the three data series (i.e., the logarithm of the spot exchange rate, the interest rate differential, and the normalized current account balance). By evaluating the t-statistics, I can only reject the null hypothesis of unit root associated with the normalized current account balance (at the 1% level). I also performed augmented Dickey-Fuller tests with constant terms and without time trends. I can reject the null hypothesis of unit root associated with the interest rate differential (with a lag, which is determined by Akaike's information criterion) at the 5% level. I also performed Phillips-Perron tests without time trends. These tests allowed me to reject the null hypothesis of unit root associated with the interest rate differential, but only at the 10% level. However, the difference between the test statistic (−2.811) and the 5% critical value (−2.882) is only marginal. I can reject the null hypothesis of unit root associated with the normalized current account balance at the 5% level. From these results, I conclude that at least two data series, the interest rate differential and the current account balance, are stationary. Table 1 summarizes these results.

[14] On the other hand, using recent data at quarterly frequency (from 2002:Q1 to 2014:Q1), Kurihara and Fukushima (2015) report that the null hypothesis of unit root associated with the level of the nominal Dollar-Yen exchange rate can be rejected with an augmented Dickey-Fuller test significant at 5% and a Phillips-Perron test significant at 10%. This reflects that the movement of the nominal Dollar-Yen exchange rate appears to be mean-reverting in the recent period. From the viewpoint of this paper, this could be caused by the fact that the Japanese quarterly current account surplus peaked and then marked deficits at times in that period.

Table 1 Unit root tests for the logarithm of the spot exchange rate, the interest rate differential, and the normalized current account balance

	Dickey-Fuller	Augmented Dickey-Fuller	Phillips-Perron
s	−1.660	−1.895	−1.730
D	−2.474	−3.072**	−2.811*
C	−3.706***	−2.413	−3.516**

Notes: This table reports the results of unit root tests for the three data series under analysis. The test statistics, Dickey-Fuller, or Z-tau are reported in this table. The significance levels to reject the null hypothesis of unit root process are denoted by the symbol "***," "**," and "*" for 1%, 5%, and 10%, respectively.

To investigate the effect of the interest rate differential and the current account balance on the exchange rate without the problem of non-stationarity of the exchange rate, I run the following VAR(1) regression for the exchange rate with quarterly data (1981Q1−2016Q2):

$$s_t = \beta_0 + \beta_1 s_{t-1} + \beta_2 D_{t-1}(10) + \beta_3 C_{t-1} + \tilde{\varepsilon}_t, \qquad (15)$$

where $\tilde{\varepsilon}_t$ can include the change in the investors' conditional expectations for the steady state spot exchange rate and for the future interest rates in the home and foreign countries, and measurement errors (in addition to the change in the risk premium attributable to factors other than the net foreign asset). This regression Eq. (15) corresponds to Eq. (14). To obtain this regression Eq. (15), I assume that $\{D_t\}$ can be approximated as an AR(1) process[15] to eliminate D_t in Eq. (14) and I replace $\Delta^F_{t-1\to t}$ with C_{t-1}. Table 2 lists the results of the regression analysis associated with Eq. (15).

[15] See Table A2 in Appendix B.

Table 2 Regression results for Eq. (15)

β_0	$t(\beta_0)$	β_1	$t(\beta_1)$	β_2	$t(\beta_2)$	β_3	$t(\beta_3)$	R^2
0.4506	4.346***	0.9062	41.659***	1.1719	2.606***	−0.0360	−3.851***	0.9665

Notes: This table reports the estimated coefficients for Eq. (15). The numbers in the columns marked as $t(\cdot)$ are t-statistics. All t-statistics are significant at the 1% level, as denoted by the symbol "***."

This result indicates that the interest rate differential[16] and the current account balance (i.e., the proxy for the increment of net foreign assets) have a significant influence on the Dollar-Yen exchange rate.

3.2. Estimation for Risk Premium

I now turn to estimating the coefficients of the risk premium in Eq. (13). I define $\bar{\rho}$ as the (sample) mean of the risk premium. Given this, the risk premium can be expressed as:

$$\rho_t^* = \bar{\rho} + \tilde{\rho}_t, \qquad (16)$$

From Eq.s (7) and (16), I obtain:

$$\begin{aligned} s_t &\approx E_t\left[s^* + R^{*\$}_{t+k(n)} - R^*_{t+k(n)}\right] + nD_t(n) - (\bar{\rho} + \tilde{\rho}_t) \\ &= E_t\left[s^* + R^{*\$}_{t+k(n)} - R^*_{t+k(n)}\right] - \bar{\rho} + nD_t(n) - \tilde{\rho}_t. \end{aligned} \qquad (17)$$

In line with Eq. (17), I run the following OLS regression:

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[16] However, the estimated value of β_2, 1.1719, is surprisingly large. Suppose $\{D_t\}$ is an AR(1) and the AR(1) coefficient is ψ_1. From Eq. (14), the coefficient for D_{t-1} in Eq. (15) becomes $(n(\psi_1 - \beta_1)$. Using the result of the regression for β_1 as the point estimator of ϕ_1 and the point estimator of ψ_1 (0.9119) from my data (see Table A2 in Appendix B), $(n(\psi_1 - \beta_1)$ is just 0.057.

$$s_t = \beta_0' + \beta_1' D_t(10) + \tilde{\varepsilon}_t'. \quad (18)$$

Since I use ten-year government bond yields as the annual interest rates, β_1' is my estimator for $n = 10$.

A major purpose of conducting this regression analysis is to obtain the regression residuals to estimate the coefficients of the risk premium in Eq. (13). Even if $\{s_t\}$ is a unit root process, as I have mentioned in the previous section, this regression is not spurious as long as $\{D_t\}$ is not a unit root process. In this case, however, the results of the t-tests become invalid, and I proceed as if the regression residuals were data series to be regressed to estimate the risk premium.[17] I will argue in favor of the stationarity in the level of the nominal exchange rate in Section 3.4. Table 3 lists the results of the regression analysis associated with Eq. (18).

Table 3 Regression results for Eq. (18)

β_0'	$t(\beta_0')$	β_1'	$t(\beta_1')$	R^2
4.4155	30.155***	13.9353	2.767***	0.3461

Notes: This table reports the estimated coefficients for Eq. (18). The numbers in the columns marked as $t(\cdot)$ are the Newey-West corrected t-statistics. All t-statistics are significant at the 1% level, as denoted by "***."

Because the Durbin-Watson statistics for this regression are quite low (0.1416), which suggests that the regression residuals are auto-correlated, I apply the Newey-West correction to the t-statistics. However, these results are *not* inconsistent with my

[17] Alternatively, defining $\hat{\varepsilon}_t = s_t - 10D_t(10)$, I run the regression Eq. (21) below, replacing the regression residuals of Eq. (18) with this $\hat{\varepsilon}$. As a result, both t-statistics for the regression coefficients are slightly lower but still hold similar significance.

assumption of the risk premium formulation Eq. (13).[18] The difference between $\beta_1' \approx 13.9353$, which is our estimator for $n = 10$, and ten is rather large in terms of standard error before the Newey-West correction (1.6023), but not after it (5.0369).[19]

Next, I use the residuals from the regression Eq. (18), $\tilde{\varepsilon}_t'$, which include $\tilde{\rho}_t$, to estimate the coefficients of the risk premium. Comparing Eq.s (17) and (18), I assume:

$$\tilde{\varepsilon}_t' = -\tilde{\rho}_t + \tilde{\theta}_t, \tag{19}$$

where $\tilde{\theta}$ includes changes in the investors' expectations for the steady state spot exchange rate and the future interest rate in both home and foreign countries, and the measurement error.

Using Eq.s (13), (16), and (19) repeatedly, I obtain:

$$\begin{aligned}\tilde{\varepsilon}_t' &= -\tilde{\rho}_t + \tilde{\theta}_t = \bar{\rho} - \rho_t^* + \tilde{\theta}_t = \bar{\rho} - (\phi_1 \rho_{t-1}^* + \phi_2 \Delta_{t-1\to t}^F + \tilde{\omega}_t) + \tilde{\theta}_t \\ &= \bar{\rho} - \phi_1(\bar{\rho} + \tilde{\rho}_{t-1}) - \phi_2 \Delta_{t-1\to t}^F - \tilde{\omega}_t + \tilde{\theta}_t \\ &= (1 - \phi_1)\bar{\rho} + \phi_1\left(-\tilde{\rho}_{t-1} + \tilde{\theta}_{t-1}\right) - \phi_1 \tilde{\theta}_{t-1} - \phi_2 \Delta_{t-1\to t}^F - \tilde{\omega}_t + \tilde{\theta}_t \\ &= (1 - \phi_1)\bar{\rho} + \phi_1 \tilde{\varepsilon}_{t-1}' - \phi_2 \Delta_{t-1\to t}^F - \phi_1 \tilde{\theta}_{t-1} - \tilde{\omega}_t + \tilde{\theta}_t \,.\end{aligned} \tag{20}$$

I then run the following OLS regression, which corresponds to Eq. (20):

$$\tilde{\varepsilon}_t' = \gamma_0 + \gamma_1 \tilde{\varepsilon}_{t-1}' + \gamma_2 C_{t-1} + \tilde{\eta}_t. \tag{21}$$

Table 4 summarizes the results for the regression Eq. (21). The signs of the regression coefficients are as expected as γ_1 and $-\gamma_2$ estimate ϕ_1 and ϕ_2 respectively. Comparing Eq.s (20) and (21), I can estimate the average risk premium in the sample period, $\bar{\rho}$, using $\frac{\gamma_0}{1-\gamma_1} = \frac{0.056}{1-0.8479} \approx 0.368$. This result implies that during the sample period (from the second quarter of 1981 to the second quarter of 2016), the Japanese Yen appreciated about

[18] Indeed, the null hypothesis of no higher order auto-correlation in residuals cannot be rejected at 1% significance level by the Breusch-Godfrey test. Similarly, the null hypothesis of no heteroskedasticity in residuals cannot be rejected at the 5% level by the Breusch-Pagan test.

[19] Because only coupon bond yields are available, I use them to represent Japanese interest rates, therefore β_1' could be *upper* biased.

26.9% relative to the US Dollar via the high risk premium.[20]

Table 4 Regression results for Eq. (21)

γ_0	$t(\gamma_0)$	γ_1	$t(\gamma_1)$	γ_2	$t(\gamma_2)$	R^2
0.0560	3.200***	0.8479	19.685***	−0.0555	−3.699***	0.741

Notes: This table reports the estimated coefficients for Eq. (21). The numbers in the columns marked as t(・) are t-statistics. All t-statistics are significant at the 1% level, as denoted by the symbol "***."

3.3. Short-Horizon Uncovered Interest Rate Parity

In many empirical studies of the uncovered interest rate parity using the equation similar to Eq. (2), researchers have used the realized spot exchange rate rather than the expected future spot exchange rate. By contrast, from Eq.s (4) and (7), the difference between the current and future realized spot exchange rates can be expressed as:

$$\begin{aligned} s_{t+k(n)} - s_t &= E_{t+k(n)}\left[s^* + R^{*\$}_{t+k(n)} - R^*_{t+k(n)}\right] - \rho^*_{t+k(n)} \\ &\quad - \left(E_t\left[s^* + R^{*\$}_{t+k(n)} - R^*_{t+k(n)}\right] + nD_t(n) - \rho^*_t\right) \\ &= \left(E_{t+k(n)} - E_t\right)\left[s^* + R^{*\$}_{t+k(n)} - R^*_{t+k(n)}\right] + n\left(-D_t(n)\right) - \left(\rho^*_{t+k(n)} - \rho^*_t\right), \end{aligned} \quad (22)$$

where $(E_{t+k(n)} - E_t)$ is an operator representing the difference in expected values between the information sets at t and $t+k(n)$. From Eq. (13), Eq. (22) becomes:

$$\begin{aligned} s_{t+k(n)} - s_t = \left(E_{t+k(n)} - E_t\right)\left[s^* + R^{*\$}_{t+k(n)} - R^*_{t+k(n)}\right] + n\left(-D_t(n)\right) \\ -\left(\phi_1\rho^*_t + \phi_2\Delta^F_{t\to t+k(n)} + \widetilde{\omega}_{t+k(n)} - \rho^*_t\right) \end{aligned}$$

[20] With the higher inflation rate in the US, this estimation is consistent with the calibration proposed by Obstfeld (2006). The author estimated that, in 2004, the effect of net foreign assets on the Yen could be a *real* appreciation of 30%.

$$= \left(E_{t+k(n)} - E_t\right)\left[s^* + R^{*\$}_{t+k(n)} - R^*_{t+k(n)}\right] + n\left(-D_t(n)\right) - \phi_2 \Delta^F_{t \to t+k(n)} + (1-\phi_1)\rho^*_t - \tilde{\omega}_{t+k(n)}$$

$$= (1-\phi_1)\bar{\rho} + n\left(-D_t(n)\right) - \phi_2 \Delta^F_{t \to t+k(n)} + (1-\phi_1)\tilde{\rho}_t - \tilde{\omega}_{t+k(n)} + \left(E_{t+k(n)} - E_t\right)\left[s^* + R^{*\$}_{t+k(n)} - R^*_{t+k(n)}\right]. \quad (23)$$

I used 3-month LIBOR rates (available from the data in 1986), therefore, $k(0.25) = 1$ in this empirical model. By setting $n = 0.25$ and treating the $(E_{t+1} - E_t)$ term as noise included in the regression residual, I run the following OLS regression, which corresponds to Eq. (23):

$$s_{t+1} - s_t = \lambda_0 + \lambda_1\left(-D_t(0.25)\right) + \lambda_2 C_t + \tilde{\delta}_{t+1}. \quad (24)$$

Note that, in this short-horizon uncovered interest rate parity regression, the current account balance does not predict the difference in the exchange rate because both variables are used simultaneously in the regression. Table 5 reports the results of the regression Eq. (24). As expected, the sign for λ_2 is negative and significant, indicating that the current account balance also has a significant effect on short-horizon uncovered interest rate parity.

Table 5 Regression results for Eq. (24)

λ_0	$t(\lambda_0)$	λ_1	$t(\lambda_1)$	λ_2	$t(\lambda_2)$	R^2
0.0152	1.160	−0.2839	−1.172	−0.0218	−2.214**	0.031

Notes: This table reports the estimated coefficients in Eq. (24). The numbers in the columns marked as $t(\cdot)$ are t-statistics. Only the estimator λ_2 is significant at the 5% level as denoted by "**."

If, as above, the risk premium has a non-zero mean, λ_0 may not be zero. However, the estimated λ_0 is not significantly different from zero in this regression. This is likely to depend on the fact that ϕ_1 is close to one. Indeed, the result of regression Eq. (21) shows that the estimator for ϕ_1 is $\gamma_1 = 0.8479$, making $(1-\phi_1)$ close to zero. Because $\tilde{\delta}$ in regression Eq. (24) is meant to include $(1-\phi_1)\tilde{\rho}$ from Eq. (23), I expected $\tilde{\delta}$ to be auto-correlated. Similarly, the Durbin-Watson statistic for this regression is close to two (1.9017), indicating a lack of auto-correlation in the residuals.

The coefficient for the short-term interest rate differential, λ_1, which is the estimator for $n = 0.25$, is insignificant and has a sign opposite of that expected, which shows a failure of uncovered interest rate parity. On the other hand, if $D_t(n)$ can be approximated as an AR(1), from Eq. (7), the expected effect of $-D_t(n)$ on $S_{t+k(n)} - S_t$ is:

$$-n\left(\psi(n)^{k(n)} - 1\right)$$

where $\psi(n)$ is the AR(1) coefficient for $D_t(n)$. In the short-horizon case of $n = 0.25$, using the point estimator for $\psi(0.25) = 0.9673$ from my data,[21]

$$-0.25\left(\psi(0.25)^{k(0.25)} - 1\right) = 0.25 \times (1 - (0.9673)^1) \approx 0.008.$$

This result is slightly positive but only marginal. In contrast, in the long-horizon case when $n = 10$ ($\psi(10) = 0.9119$),

$$-10\left(\psi(10)^{k(10)} - 1\right) = 10 \times (1 - (0.9119)^{40}) \approx 9.750.$$

Some studies report that uncovered interest rate parity holds better with a long horizon than with a short horizon, and the coefficient of the interest rate differential in uncovered interest rate parity is often lower than n even when significant with a longer horizon. These results suggest that if the interest rate differential is approximated as an AR(1), it can explain that uncovered interest rate parity holds better with a long horizon than with a short horizon.[22] In addition to n being smaller, the tendency for the AR(1) coefficient in the

[21] See Table A2 in Appendix B.

[22] A more serious failure, which is called the uncovered interest rate parity puzzle, is that the coefficient of the interest rate differential in the uncovered interest rate parity regression is negative as the result of Eq. (24) in this paper. This means that the low-interest-rate currencies depreciate relative to high-interest-rate currencies. This issue remains open for future research.

shorter-term interest rate differential to approach one can worsen the results for uncovered interest rate parity regressions.

3.4. Annual Frequency

To this point, I have used quarterly data for my analyses. If exchange rates are stationary in level, the persistency parameters should be smaller when using annual data rather than quarterly data. To verify this, I repeated all previous longer-horizon interest rate parity regressions with annual data and compared the point estimators with those for quarterly data. From the quarterly data, I generated four annual data series that respectively begin in January, April, July, and October.[23] I ran the regressions using each of these annual data series separately. First, for the AR(1) model of the logarithm of the spot exchange rate, the coefficient fell from 0.9728 for quarterly data (see α_1 in Table A1 in Appendix B) into the range between 0.8912 and 0.8571 for annual data (the average of four estimators is 0.8722). Secondly, the coefficient β_1 in Eq. (15) of the VAR(1) regression fell from 0.9062 for quarterly data into the range between 0.6936 and 0.5959 for annual data (the average of four estimators is 0.6466). Lastly, the coefficient γ_1 in Eq. (21) of the risk premium estimation regression fell from 0.8479 for quarterly data into the range between 0.4310 and 0.2927 for annual data (the average of four estimators is 0.3730). These results indicate that, as expected, the use of annual data reduces the persistency parameters.

All the values of the Japanese annual current balances during the sample period were positive. Therefore, I estimated the regression defined by Eq. (21) by replacing the

[23] While the quarterly data allows for a sample of 141, the annual data series allow for a sample of 36 (beginning in January and April) or 35 (beginning in July and October).

normalized current account balance with the logarithm of the current account balance. This regression analysis produced higher t-statistics (all of which increased by one or more) for the coefficient of current account balances, γ_2. This suggests that changes in net foreign assets may have non-linear effects on exchange rate risk premiums. However, using the logarithm of the current account balance decreases the t-statistics for the persistency parameters, γ_1, for all four series.

4. Summary and Concluding Remarks

In this paper, I discussed the mechanism through which net foreign asset holdings affect exchange rates from the standpoint of asset pricing without explicitly using intertemporal budget constraints. Because risk premiums are the central issue in the current research on asset pricing, I have focused on the time-varying and persistent exchange rate risk premium related to uncovered interest rate parity. To do so, I have argued that spot exchange rate risk premiums vary through changes in net foreign asset holdings and, especially, increase with the accumulation of net foreign asset holdings.

In the case of an infinitely long horizon, especially under the stationarity assumption, the uncovered interest rate parity equation is equivalent to a present value model of the level of exchange rates. My empirical results provide evidence consistent with my formulation of the time-varying and persistent risk premiums associated with changes in net foreign assets. The results also suggest that, in the longer-horizon case, the Japanese current account balance predicts the Dollar-Yen exchange rate level. While academics agree that nominal exchange rates are non-stationary in level generally, I argue that the strong persistent effect associated with changes in net foreign asset holdings causes exchange rates to appear non-stationary in level. The results of the regression replications with annual data show that the persistent effect shrinks in magnitude as the time interval becomes longer. These results indicate the stationarity in the level of nominal exchange rates.

Although my risk premium formulation does not solve the short horizon uncovered interest rate parity puzzle, the change in the current account balance has a significant effect on short-horizon uncovered interest rate parity, too. Moreover, using the present value model of the level of exchange rates combined with the AR(1) approximation for interest rate differentials, I demonstrated that it can explain the failure of uncovered interest rate parity (i.e., uncovered interest rate parity holds better with a long horizon than with a short horizon).

To confirm the arguments presented here, future empirical work based on the analysis of other currencies would be useful. However, there is a caveat. I used the Dollar-based current account balance as the proxy for change in Japanese net foreign assets. As stated in Section 3.1, the share of Yen-based trade in foreign trade is small in Japan. The share of Dollar-based trade in Japan is quite large even in Asia. Therefore, the Dollar-Yen exchange rate is advantageous for the research in this paper. If I use other currencies such as the Euro against the Yen, the results may be different. Similarly, if there is only a small share of Yen based assets in the US and I use the current account balance for the US, the results may be different. Another future line of research that would shed light on the issues discussed in this paper relates to the identification of factors other than changes in net foreign asset holdings that may explain time-varying exchange rate risk premiums.

Appendix A: Proof of Proposition

The first order condition of the representative investor's problem in Section 2.2 is:

$$\mathrm{E}\left[\left\{\left(1+\mathrm{R}^{\$}\right)\mathrm{S}^{*}-(1+\mathrm{R})\mathrm{S}\right\}\mathrm{U}'(\cdot)\right]=0. \tag{A1}$$

By applying the implicit function theorem to the first order condition, Eq. (A1), I obtain:

$$\frac{\mathrm{dS}}{\mathrm{dF}}=-\frac{\mathrm{E}\left[\left\{\left(1+\mathrm{R}^{\$}\right)\mathrm{S}^{*}-(1+\mathrm{R})\mathrm{S}\right\}^{2}\mathrm{U}''(\cdot)\right]}{\mathrm{E}[-(1+\mathrm{R})\mathrm{U}'(\cdot)-(1+\mathrm{R})\mathrm{F}\{(1+\mathrm{R}^{\$})\mathrm{S}^{*}-(1+\mathrm{R})\mathrm{S}\}\mathrm{U}''(\cdot)]}$$

$$=\frac{\mathrm{E}\left[\left\{\left(1+\mathrm{R}^{\$}\right)\mathrm{S}^{*}-(1+\mathrm{i})\mathrm{S}\right\}^{2}\mathrm{U}''(\cdot)\right]}{(1+\mathrm{R})\mathrm{E}[\mathrm{U}'(\cdot)+\mathrm{F}\{(1+\mathrm{R}^{\$})\mathrm{S}^{*}-(1+\mathrm{R})\mathrm{S}\}\mathrm{U}''(\cdot)]}.$$

I can fix Rs and S at an arbitrary equilibrium value. If F=0 is optimal, it becomes trivial to render $\frac{\mathrm{dS}}{\mathrm{dF}}<0$ as $\mathrm{U}'(\cdot)>0$ and $\mathrm{U}''(\cdot)<0$. If $\mathrm{F}\neq 0$ is optimal, I only need to demonstrate:

$$\mathrm{E}\left[\mathrm{F}\left\{\left(1+\mathrm{R}^{\$}\right)\mathrm{S}^{*}-(1+\mathrm{R})\mathrm{S}\right\}\mathrm{U}''(\cdot)\right]\geq 0.$$

First, suppose that $\mathrm{F}>0$. Given this, I will show:

$$\mathrm{E}\left[\left\{\left(1+\mathrm{R}^{\$}\right)\mathrm{S}^{*}-(1+\mathrm{R})\mathrm{S}\right\}\mathrm{U}''(\cdot)\right]\geq 0.$$

From the first order condition, Eq. (A1), and $\mathrm{U}'(\cdot)>0$,

$$\text{For }\ \forall\bar{\mathrm{S}}\geq\frac{1+\mathrm{R}}{1+\mathrm{R}^{\$}}\mathrm{S},\ \ 0<\int_{\bar{\mathrm{S}}}^{+\infty}\left\{\left(1+\mathrm{R}^{\$}\right)\mathrm{S}^{*}-(1+\mathrm{R})\mathrm{S}\right\}\mathrm{U}'\left(\mathrm{W}^{*}(\mathrm{F};\mathrm{S},\mathrm{S}^{*})\right)\mathrm{f}(\mathrm{S}^{*})\mathrm{dS}^{*}$$

$$=-\int_{0}^{\bar{\mathrm{S}}}\left\{\left(1+\mathrm{R}^{\$}\right)\mathrm{S}^{*}-(1+\mathrm{R})\mathrm{S}\right\}\mathrm{U}'\left(\mathrm{W}^{*}(\mathrm{F};\mathrm{S},\mathrm{S}^{*})\right)\mathrm{f}(\mathrm{S}^{*})\mathrm{dS}^{*}$$

where $\mathrm{f}(\mathrm{S}^{*})$ is the probability density function of S^{*}. Using the definition of $\mathrm{A}(\mathrm{W}^{*})$ from the text, I can express this inequality as:

$$0<\int_{0}^{\bar{\mathrm{S}}}\left\{\left(1+\mathrm{R}^{\$}\right)\mathrm{S}^{*}-(1+\mathrm{R})\mathrm{S}\right\}\frac{\mathrm{U}''(\cdot)}{\mathrm{A}(\cdot)}\mathrm{f}(\mathrm{S}^{*})\mathrm{dS}^{*}$$

$$=-\int_{\bar{\mathrm{S}}}^{+\infty}\left\{\left(1+\mathrm{R}^{\$}\right)\mathrm{S}^{*}-(1+\mathrm{R})\mathrm{S}\right\}\frac{\mathrm{U}''(\cdot)}{\mathrm{A}(\cdot)}\mathrm{f}(\mathrm{S}^{*})\mathrm{dS}^{*}. \tag{A2}$$

From Eq. (8), future wealth W^{*} increases in S^{*} if $\mathrm{F}>0$. Because $\mathrm{A}(\mathrm{W}^{*})$ is positive and non-increasing in W^{*}, it also is non-increasing in S^{*}. Accordingly, I obtain:

$$\int_0^{\bar{S}} \{(1+R^{\$})S^* - (1+R)S\}U''(\cdot)\, f(S^*)dS^*$$

$$\geq -\int_{\bar{S}}^{+\infty} \{(1+R^{\$})S^* - (1+R)S\}U''(\cdot)\, f(S^*)dS^* \Leftrightarrow$$

$$0 \leq \int_0^{\bar{S}} \{(1+R^{\$})S^* - (1+R)S\}U''(\cdot)\, f(S^*)dS^*$$

$$+ \int_{\bar{S}}^{+\infty} \{(1+R^{\$})S^* - (1+R)S\}U''(\cdot)\, f(S^*)dS^*$$

$$= E[\{(1+R^{\$})S^* - (1+R)S\}U''(\cdot)].$$

Now suppose that $F<0$. Given this condition, I will show that:

$$E[\{(1+R^{\$})S^* - (1+R)S\}U''(\cdot)] \leq 0.$$

If $F < 0$, future wealth W^* decreases in S^*. Similarly, $A(\cdot)$ is non-decreasing in S^*. Therefore, from Eq. (A2), I obtain:

$$\int_0^{\bar{S}} \{(1+R^{\$})S^* - (1+R)S\}U''(\cdot)\, f(S^*)dS^*$$

$$\leq -\int_{\bar{S}}^{+\infty} \{(1+R^{\$})S^* - (1+R)S\}U''(\cdot)\, f(S^*)dS^* \Leftrightarrow$$

$$0 \geq \int_0^{\bar{S}} \{(1+R^{\$})S^* - (1+R)S\}U''(\cdot)\, f(S^*)dS^*$$

$$+ \int_{\bar{S}}^{+\infty} \{(1+R^{\$})S^* - (1+R)S\}U''(\cdot)\, f(S^*)dS^*$$

$$= E[\{(1+R^{\$})S^* - (1+R)S\}U''(\cdot)].$$

Now define the ith investor's net foreign asset demand as $F_i(S)$ such that

$$F_i(S) \equiv \mathrm{argmax}_{F_i}\, W^*(F_i; S, S^*),$$

and modify F to be the initial net foreign assets in the country. The market clearing condition is then $F = \sum_i F_i(S)$ and $dF = \sum_i dF_i(S)$. Since all the investors are non-increasing absolute risk averse, the above result, combined with the market clearing condition, implies that the increase in net foreign asset holdings makes the spot exchange rate appreciate. ■

I have shown that, if $F > 0$, $E\left[\left\{\left(1+R^{\$}\right)S^{*}-(1+R)S\right\}U''(\cdot)\right] \geq 0$. From this, I can show that, if $F > 0$, an increase in the investor's initial wealth will be consistent with the spot exchange rate depreciation. Indeed, suppose $F > 0$ and

$E\left[\left\{\left(1+R^{\$}\right)S^{*}-(1+R)S\right\}U''(\cdot)\right] \geq 0$. Then,

$$\frac{dS}{dW} = -\frac{E\left[(1+R)\left\{\left(1+R^{\$}\right)S^{*}-(1+R)S\right\}U''(\cdot)\right]}{E[-(1+R)U'(\cdot)-(1+R)F\{(1+R^{\$})S^{*}-(1+R)S\}U''(\cdot)]}$$

$$= \frac{E\left[\left\{\left(1+R^{\$}\right)S^{*}-(1+R)S\right\}U''(\cdot)\right]}{E[U'(\cdot)+F\{(1+R^{\$})S^{*}-(1+R)S\}U''(\cdot)]} > 0.$$

Appendix B: Data

I obtained interest rate data from the Organization for Economic Co-operation and Development (OECD). For my analyses, I used ten-year government bond yields as the ten-year annual interest rate. However, only monthly averages of government bond yields are available, and Japanese government bonds are coupon bonds, not discounted bonds. In addition, the data related to Japanese government bond yields are not continuous, so I merged two different data series (January, 1981—February, 1992 and January, 1989—April, 2016). I obtained older Japanese government bond yields, which are sourced from OECD, from the Bank of Japan's Statistics Monthly. When the periods overlapped, resulting in two values in a single period, I took the mean of the two values for the overlapping periods.

Differences between data in the overlapping periods were minimal. For the OLS regression analysis on short-horizon uncovered interest rate parity in Section 3.3, I used the closing value of the 3-month LIBOR rate (based on the Dollar and Yen) for the previous period to represent a given quarter. These data were available from January 1986 to March 2016.

I used the monthly average of the spot exchange rate in the Tokyo Foreign Exchange; these data were available from the Bank of Japan. I used the first-month average spot exchange rates and the interest rates as the values to represent a given quarter. Because capital control was largely deregulated in Japan in 1980, I used data from 1981 onward to avoid the constraints associated with capital control and the structural changes that accompany it. I used the current account balance as a proxy to indicate an incremental change in net foreign assets. I obtained the Japanese quarterly Dollar-based current account balance data from the IMF for the period from the fourth quarter of 1980 to the fourth quarter of 2008. For the period between the first quarter of 1985 and the first quarter of 2016, I collected data from the OECD. The IMF Balance of Payments Manual was recently renewed. As such, I obtained data related to the Japanese current account balances until 2013 from the manual's fifth edition (BPM5) and data from the first quarter of 1996 to the first quarter of 2016 from the sixth edition (BPM6). When periods overlapped, I again averaged the overlapping values. The differences in the overlapping periods were relatively small except for the fourth quarter of 2012, where the BPM5 reports a slight surplus, but the BPM6 reports a slight deficit. This occurred because the Japanese current account surplus has recently begun to shrink. I ran the same regressions as those described in the paper using data from the smaller sample without the BPM6's current account balance; the results did not vary significantly.

Because of these decisions, my data series start with the monthly average spot exchange rate and the interest rates of January of 1981, and the current account balance of the fourth quarter of 1980. I, then, took the logarithm of the spot exchange rate, used decimal fractions to represent interest rate differentials (e.g., 0.05 instead of 5[%]), and normalized the current account balances dividing them by the sample median of the absolute values of current account balances.

Table A1 reports the result of the estimation of $s_{t+1} = \alpha_0 + \alpha_1 s_t + \tilde{\zeta}$ where I assume that the logarithm of the nominal Dollar-Yen exchange rate, s_t, is AR(1).

Table A1 Estimation for AR(1) model of the logarithm of the nominal Dollar-Yen exchange rate

α_0	$t(\alpha_0)$	α_1	$t(\alpha_1)$	R^2
0.1282	1.324	0.9728	48.029***	0.9625

Notes: This table reports the estimated coefficients for $s_{t+1} = \alpha_0 + \alpha_1 s_t + \tilde{\zeta}$. The numbers in the columns marked as $t(\cdot)$ are the Newey-West corrected t-statistics. The symbol "***" denotes a significance level at 1%.

Even if the value of the coefficient in this estimation ($\alpha_1 = 0.9732$) is correct, it is difficult to reject the null hypothesis of unit root. I also use the AR(1) approximation for the interest rate differentials of the ten-year government bond yields denoted by $D(10)$, and of the 3-month LIBOR rates denoted by $D(0.25)$. Table A2 reports the results of the estimation.

Table A2 AR(1) approximation for the interest rate differentials

	ψ_0	$t(\psi_0)$	ψ_1	$t(\psi_1)$	R^2
$D(0.25)$	−0.001	−0.815	0.9673	36.813***	0.9272

D(10)	0.003	2.079**	0.9119	21.009***	0.8273

Notes: This table reports the estimated coefficients for $D_{t+1} = \psi_0 + \psi_1 D_t + \tilde{\xi}$. The numbers in the columns marked as $t(\cdot)$ are the Newey-West corrected t-statistics. The symbols "***" and "**" denote significance levels at 1% and 5% respectively.

Even though these variables can follow other processes, due to the high coefficients of determination, the AR(1) is a good approximation for the research in this paper.

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