Do spot return regressions convey useful information on return predictability?

Seongman Moon*
Universidad Carlos III de Madrid

Carlos Velasco†
Universidad Carlos III de Madrid

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Abstract

This paper shows that the estimates of the spot return regressions may not convey useful information about the predictability of excess returns if exchange rates are generated from a present value model where the discount factor is near unity and fundamentals are I(1) as in Engel and West (2005). The main reason is that the present value model induces a large bias in the estimation of the regressions accompanied by a high variability of the estimates. Nevertheless, the implications of the present value model for the volatility and persistence of the spot return and the forward premium are consistent with the data.

JEL Classification: F31, C13.

Keywords: discount factor, fundamentals, contemporaneous correlation, present value model.

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*Department of Economics, Universidad Carlos III de Madrid, Calle Madrid 126, 28903 Getafe Madrid, SPAIN. Email: smoon@eco.uc3m.es, Tel.: +34-91-624-8668, Fax: +34-91-624-9329. Research support from Spanish Secretary of Education (SEJ2007-63098) is gratefully acknowledged.

†Department of Economics, Universidad Carlos III de Madrid, Calle Madrid 126, 28903 Getafe Madrid, SPAIN. Email: Carlos.Velasco@uc3m.es, Tel.: +34-91-624-9646, Fax: +34-91-624-9329. Research support from the Spanish Plan Nacional de I+D+I (SEJ2007-62908) is gratefully acknowledged.
1 Introduction

Uncovered interest parity (UIP), one of major building blocks in many international macroeconomic models, states that the expected spot return on foreign relative to domestic currency must be equal to the interest rate difference between domestic and foreign countries under the assumptions of rational expectations and risk neutral preferences. Numerous studies, however, have persistently found the statistically significant negative estimates from the return regression of the log of the exchange rate change on the lagged interest rate difference (or the lagged forward premium) with the null value of one. These results imply that a higher interest rate currency tends to appreciate, rather than to depreciate. This is puzzling because it suggests not only the predictability of excess returns but also a profitable trading scheme.\(^1\)

Based on the regression test, Fama (1984) illustrated the degree of variability in the expected foreign excess return under the assumption of rational expectations. One popular relation he derived is that if the slope coefficient is less than one half in the return regression, the variance of the rational expectations risk premium (the expected excess return) should be greater than that of the expected change in the exchange rate. Since then, the risk premium literature has used this condition to judge the performance of international asset pricing models for explaining the observed behavior of forward and spot rates. However, most models miserably fail in generating a high volatility of the risk premium, which refers to the forward premium puzzle.\(^2\) The implicit presumption in the literature was that the estimated slope coefficient in the regression of the spot return on the forward premium would accurately convey information about the predictability of excess returns.

This paper questions this presumption on the information content of the regression.\(^3\) We show that the present value model of exchange rates with a near unity discount factor and unit root fundamentals generates a large magnitude of the relative variance between the spot return and the forward premium, consistent

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2. Recently, Alvarez, Atkeson, and Kehoe (2009) and Verdelhan (2010) are successful for relating the cause of the negative estimates to the rational expectations risk premium.
3. Baillie and Bollerslev (2000) and West (2008) also study a similar issue. In particular, our paper is close to West (2008) who considered the same assumptions as ours.
with the data. More importantly, this large relative variance appears simultaneously with a large bias in the estimation of the return regression, although the bias does not significantly affect the size of the t test for the estimated slope coefficient. Finally, we show that it should not be uncommon to find insignificant negative estimates frequently if the exchange rates would be generated from the typical present value model. This implies that the estimates from the return regression may not be informative and one should be careful to link the magnitude of the estimated slope coefficient to the presence of the risk premium (or a deviation from UIP) based on Fama’s volatility relation.

2 Spot Return Regressions in Foreign Exchange Markets

In this section, we first reproduce the estimation results from the standard test in the literature to motivate our study. Then, we present the present value model of exchange rates to set out explicitly our question.

Consider the following bivariate regression model,

\[ s_t - s_{t-1} = \alpha + \beta(f_{t-1} - s_{t-1}) + u_t, \]  
\[ f_t - s_t = \delta + \varphi(f_{t-1} - s_{t-1}) + v_t, \]

where \( s_t - s_{t-1} \) denotes the spot return, \( f_{t-1} - s_{t-1} \) is the forward premium and assumed to be stationary, and \( \beta = 1 \) under UIP. The covariance matrix of the error terms, \( u_t \) and \( v_t \), is denoted as

\[ \Sigma = \begin{pmatrix} \sigma_u^2 & \sigma_{uv} \\ \sigma_{uv} & \sigma_v^2 \end{pmatrix}. \]

2.1 Motivation

Panel A of Table 1 reports the results from regression (1) using monthly spot and one-month forward exchange rates from the sample period of 1975-2009: Estimates of \( \beta \) are negative, their difference from the null value is very large, and the conventional t-test strongly rejects the UIP hypothesis. These results confirm well-known empirical regularities. In contrast, this strong deviation from UIP disappears in the subsample which excludes observations in 1980-87 from the entire sample. All currencies except for the Canadian dollar are not statistically sig-

\[^4\text{Engel and West (2005) use these two assumptions to study a link between fundamentals and near random walk behavior of exchange rates.}\]
significant as reported in Panel B. These results are consistent with the empirical
findings by Moon and Velasco (2011).

A peculiar phenomenon regarding the estimation results using the subsample
period is that although the estimated slope coefficients are not statistically signif-
icant, most estimates are negative, suggesting a large absolute difference between
the estimates and the null value of one. The main objective of the present paper
is to provide a reason why these insignificant negative estimates appear under the
null hypothesis, based on the typical present value model of exchange rates.

2.2 Robust test of UIP

Although regression (1) has been widely used to test the UIP hypothesis since
Fama (1984), it is now well known that nonzero correlations between $u_t$ and $v_t$
may generate an endogeneity problem. For example, Mankiw and Shapiro (1986)
illustrate that the $t$-test based on the conventional critical values tends to overreject
the null if the regressor exhibits strong persistence and the absolute value of the
contemporaneous correlation between disturbances to the dependent variable and
to the regressor is high [see also Stambaugh (1999)]. To investigate how much this
affects statistical inference, we conduct the conditional test by Jasson and Moreira
(2006) which is robust to the endogeneity problem and most powerful within the
class of unbiased tests, following Maynard (2006). As reported in third and fourth
columns in Table 1, the results from the conditional test confirm those from the
conventional $t$ test above in both samples. The main reason is that one of the two
conditions for the overrejection illustrated by Mankiw and Shapiro (1986) is not
satisfied in the data. The contemporaneous correlations between the innovations
to foreign excess returns and to the forward premium are close to zero as reported
in Table 2: the absolute values of the estimated correlations are less than 0.1 for
almost all currencies.

2.3 Spot Return Regressions in a Present Value Model of Exchange Rates

We now derive the covariance matrix of innovations, $\Sigma$, from the typical present
value model of exchange rates in order to study the information content in the
spot return regression. In this model, the spot exchange rate $s_t$ is expressed as a
discounted sum of current and expected future fundamentals,

\[ s_t = (1 - b) \sum_{i=0}^{\infty} b^i E_t[w_{t+i}], \quad (4) \]

where \( E_t(\cdot) \) is the mathematical expectation conditional on a time \( t \) information set, \( 0 < b < 1 \) represents the discount factor, and \( w_t \) represents the linear combination of logs of fundamental variables such as money and output [see, Engel and West (2005) for a more general framework]. This relation between the exchange rate and the fundamentals can be obtained from the typical monetary model of exchange rates under the assumptions of no bubbles and UIP.\(^5\)

Following Engel and West (2005) and West (2008), we assume that the fundamental process evolves in the following way

\[ \Delta w_t = \Delta w_{1,t} + \eta_2,t, \quad (5) \]

\[ \Delta w_{1,t} = \phi \Delta w_{1,t-1} + \eta_1,t, \]

where \( 0 < \phi < 1 \) and both \( \eta_{1,t} \) and \( \eta_{2,t} \) are iid with zero mean normal distributions with variance \( \sigma_1^2 \) and \( \sigma_2^2 \), respectively.\(^6\) From equations (4)-(5), both the forecasting error and the forward premium are derived as

\[ u_t = s_t - E_{t-1}[s_t] = \frac{1}{1 - b\phi} (\eta_{1,t} + (1 - b\phi)\eta_{2,t}), \]

\[ x_t = f_t - s_t = \frac{(1 - b)}{1 - b\phi} \phi \Delta w_{1,t}, \quad (6) \]

where the persistence of the forward premium is governed by the parameter \( \phi \) in \( w_{1,t} \). Since \( \varphi = \phi \), the error term \( v_t \) in regression (2) is defined by

\[ v_t = \frac{1 - b}{1 - b\phi} \phi \eta_{1,t}. \quad (7) \]

\(^5\)In the ad hoc monetary models, a home money market relation is given by

\[ m_t - p_t = y_t - \kappa i_t, \]

where \( m_t - p_t \) is home real money demand, \( y_t \) is home output, \( i_t \) is home nominal interest rate, and \( \kappa \) is the interest elasticity of money demand. Foreign money demand can be defined analogously. From home and foreign money market relations, UIP, and covered interest parity, we can derive the present value relation (4) where \( w_t = (m_t - y_t) - (m_t^* - y_t^*) \) if purchasing power parity holds and \( w_t = (m_t - y_t) - (m_t^* - y_t^*) + s_t + p_t^* - p_t \) if it does not hold.

\(^6\)Our results are robust to other specifications of fundamental processes as long as the two assumptions (a near unity discount factor and unit root fundamentals) are satisfied. For example, see Section 4 for an alternative process for \( w_t \) where the fundamental process is the sum of random walk and stationary components.
Then, the covariance matrix $\Sigma = \Sigma(\phi, b)$ of the innovations $u_t$ and $v_t$ in the bivariate regressions (1)-(2) has elements given by

$$
\sigma_{uv}(\phi, b) = \left(\frac{1 - b}{1 - b\phi}\right)^2 \frac{\phi}{1 - b} \sigma_1^2,
$$

$$
\sigma_u^2(\phi, b) = \left(\frac{1}{1 - b\phi}\right)^2 \left(\sigma_1^2 + (1 - b\phi)^2 \sigma_2^2\right),
$$

$$
\sigma_v^2(\phi, b) = \left(\frac{1 - b}{1 - b\phi}\right)^2 \phi^2 \sigma_1^2,
$$

(8)

where the covariance is a function of the two economic parameters, the persistence parameter in the fundamental process and the discount factor. Note that $v_t$ is positively correlated with $u_t$, reflecting an endogenous feedback from forecasting errors to the future values of the regressor.

### 3 Information Content in the Spot Return Regression

In this section, we study the influence of the magnitude of two key quantities, $\frac{\sigma_{uv}}{\sigma_v^2}$ and $\frac{\sigma_u}{\sigma_v}$, on the information content of the spot regression. We show that the scaled correlation $\frac{\sigma_{uv}}{\sigma_v^2}$ affects the magnitude of the finite sample bias, while the noise-to-signal ratio $\frac{\sigma_u}{\sigma_v}$ mainly determines the variance of the estimated slope coefficient, $\tilde{\beta}$. Further, the relative quantity, $\frac{\sigma_{uv}}{\sigma_v^2} / \frac{\sigma_u}{\sigma_v}$, which is the contemporaneous correlation, mainly contributes to the over-rejections of the $t$-test along with the persistent regressor. Finally, we link them to the two economic parameters, $\phi$ and $b$, in the typical present value model of exchange rates.

Consider the bivariate system of regressions (1)-(2) where $f_t$ and $s_t$ are generated from the present value model specified in the previous section. Following Stambaugh (1999), the expected sampling error of the estimated slope coefficient under normality is calculated by

$$
E[\hat{\beta} - \beta_0] = \frac{\sigma_{uv}}{\sigma_v^2} E[\hat{\phi} - \phi] \approx -\frac{1}{\phi(1 - b)} \frac{(1 + 3\phi)}{T},
$$

(9)

where the first order approximation to $E[\hat{\phi} - \phi]$ is defined by $-(1 + 3\phi)/T$ using the analysis of Kendall (1954, Eq.(20)). Equation (9) shows that the discount factor $b$ critically determines the feedback scale factor, $\frac{\sigma_{uv}}{\sigma_v^2}$, and thus strongly affects the magnitude of the bias: the value of the scale factor, $\frac{1}{\phi(1 - b)}$, significantly increases as $b$ goes to unity. For example, for $\phi = 0.3$, $E[\hat{\beta} - \beta_0]$ is 1.08 for $b = 0.99$, 0.22 for $b = 0.95$, and 0.11 for $b = 0.9$ for a currently available monthly sample of $T = 400$ in foreign exchange data. On the other hand,
\[ \frac{\sigma_u}{\sigma_v} = \frac{1}{\phi(1-b)} \sqrt{1 + (1 - b\phi)^2 \frac{\sigma_\eta^2}{\sigma_\xi^2}} \] also becomes larger as \( b \) is closer to one. This can be best illustrated in equation (6) where the error term and the regressor become

\[ u_t \approx \frac{1}{1 - \phi} \eta_{1,t} + \eta_{2,t}, \quad x_{t-1} \approx \frac{\phi(1-b)}{1 - \phi} \eta_{1,t-1}. \]  

(10)

Here, the variance of the regressor is close to zero for \( b \) sufficiently close to one.\(^7\) As a result, larger values of \( \frac{\sigma_u}{\sigma_v} \) implied by the near unity value of \( b \) make the distribution of \( \hat{\beta} \) wider and thus the estimated slope coefficient becomes less informative.\(^8\)

The contemporaneous correlation between \( u_t \) and \( v_t \) is the ratio between \( \frac{\sigma_{uv}}{\sigma_v^2} \) and \( \frac{\sigma_u}{\sigma_v} \), as defined in the following decomposition,

\[ \rho = \frac{\sigma_{uv}/\sigma_v^2}{\sigma_u/\sigma_v} = \frac{1}{\sqrt{1 + (1 - b\phi)^2 \frac{\sigma_\eta^2}{\sigma_\xi^2}}}. \]  

(11)

Equation (11) implies that a large magnitude of bias by itself does not necessarily lead to severe over-rejections of the \( t \)-test relative to its nominal size. Rather, what matters is the ratio between \( \frac{\sigma_{uv}}{\sigma_v^2} \) and \( \frac{\sigma_u}{\sigma_v} \), which determines the standardized bias of the \( t \)-test. Therefore, it is possible that this present value model can produce \textit{insignificant} values of the estimates very far away from the null value, including the negative values observed in the data, while the volatility of spot rates and persistence of the forward premium from the present value model are compatible with the data. For example, the large relative magnitude between the variances of the prediction error \( u_t \) and of the regressor implied by equation (10) is consistent with the foreign exchange data (reported in Table 2), as \( b \) is sufficiently close to unity.

### 4 Monte Carlo Experiments

In this section, we conduct Monte Carlo simulations to examine how severely the two economic parameters, \( \phi \) and \( b \), affect the information content of the slope coefficient in the spot return regression

We use as data generating process (4) for the spot exchange rate and the modified version of equation (5) for the fundamentals. To be compatible with the

\(^7\)Although we restrict our attention for the case of constant \( b \), its relaxation strengthens our results. For example, West (2008) assumes \( b = 1 - \frac{d}{\sqrt{T}} \) where \( d > 0 \) is constant and shows that the conventional \( t \)-test is not consistent.

\(^8\)Baillie and Bollerslev (2000) and West (2008) arrive at similar conclusions on the information content of the estimated slope coefficient using different assumptions.
evidence on the persistence of the forward premium in the data, we modify the fundamental process in (5) by

\[
\Delta w_t = \Delta w_{1,t} + \eta_{2,t},
\]

\[
\Delta w_{1,t} = \phi \Delta w_{1,t-1} + \eta_{1,t} + \theta \eta_{1,t-1},
\]

where \( \theta < 0 \) and \( 0 < \phi + \theta < 1 \). For robustness, we also consider an alternative fundamental process given by

\[
w_t = w_{1,t} + w_{2,t},
\]

\[
w_{1,t} = \phi w_{1,t-1} + \eta_{1,t},
\]

\[
w_{2,t} = w_{2,t-1} + \eta_{2,t},
\]

where \( 0 < \phi < 1 \) and both \( \eta_{1,t} \) and \( \eta_{2,t} \) are i.i.d. zero mean normal distributions with variance \( \sigma_1^2 \) and \( \sigma_2^2 \), respectively.

For both specifications, we consider nine combinations of the following parameter values \( \phi = [0.9, 0.95, 0.99] \) and \( b = [0.9, 0.95, 0.97] \). For the first specification, we set \( (\phi, \theta) = [(0.9, -0.8), (0.95, -0.85), (0.99, -0.9)] \). We obtain the absolute value of \( \rho \) around \([0.2, 0.25]\), which covers its upper bound in the data by varying the magnitude of the relative variance \( \sigma_2^2 / \sigma_1^2 \).

4.1 Results

Table 3 presents simulation results using 10,000 repetitions. Panel A reports the results from the case in which \( s_t \) is generated using equations (4) and (12) and Panel B presents the results from the case in which \( s_t \) is generated using (4) and (13). The conventional \( t \)-test is conducted for 1, 5, 10% significant levels against left-tail and right-tail alternatives, respectively. To conserve space, we only report the results at the 5% significant level for left- and right-tail alternatives in the fourth and fifth columns. We also report the five different percentiles of the distribution of \( \hat{\beta} \), the estimates of the contemporaneous correlation, \( \hat{\rho} \), and the estimates of the first order persistence parameter of the regressor, \( \hat{\varphi} \).

Overall, we find that the distribution of \( \hat{\beta} \) is very broad in that we observe negative values of the estimates up to 40th percentiles under the null hypothesis of \( \beta = 1 \). Nevertheless, the size of the \( t \)-test is close to its nominal value.

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\(^9\)We omitted our analysis with this generalization in Section 3 for the sake of simplicity but it is available upon request.
In both specifications, size distortions do not appear despite of strong persistence in the regressor. This is mainly because we designed the simulations so that the contemporaneous correlation between the disturbance to the dependent variable and to the regressor is very low, following the evidence in the data. The absolute values of the estimate of ρ in our simulations are close to the maximum value obtained from the data (see fifth column in Table 2), implying that the overrejection of the conventional t-test is not likely to occur in the data.

Despite of the small magnitude of the contemporaneous correlation, the magnitude of the bias for the estimation turns out to be large. In the first specification with \((b = 0.9, \phi = 0.9, \text{ and } \theta = -0.8)\), the estimated slope coefficient, \(\hat{\beta}\), is -1.41 at the 20th percentile where the null value is 1, implying that its distribution is very broad and affected by the small sample bias. We find similar results from the second specification. So in both cases, we often observe insignificant negative values of the estimated slope coefficient.

5 Conclusions

This paper shows that the estimates of the spot return regression may not convey useful information about the predictability of excess returns if exchange rates are generated from the present value model where the discount factor is near unity and fundamentals follow unit root processes. The main reason is that the model induces a large bias in the estimation of the regression accompanied by a large magnitude of the variance of the estimate. Empirical evidence on the near unity value of the discount factor and on the large magnitude of the relative variance between the spot return and the forward premium supports our assumptions on the present value model.
Reference


Table 1 Results from the conventional $t$ test and the conditional test

<table>
<thead>
<tr>
<th>Series</th>
<th>$\beta$</th>
<th>t-sta</th>
<th>[2.5, 97.5]</th>
<th>$R^2$</th>
<th>$T$</th>
</tr>
</thead>
<tbody>
<tr>
<td>Panel A. The entire sample period: 1975-2009</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>GDM</td>
<td>-0.92</td>
<td>$-2.47$</td>
<td>[-1.92, 1.96]</td>
<td>0.01</td>
<td>418</td>
</tr>
<tr>
<td>BRP</td>
<td>-1.14</td>
<td>$-3.01$</td>
<td>[-1.92, 2.01]</td>
<td>0.01</td>
<td>418</td>
</tr>
<tr>
<td>JPY</td>
<td>-2.07</td>
<td>$-4.49$</td>
<td>[-1.92, 1.99]</td>
<td>0.02</td>
<td>418</td>
</tr>
<tr>
<td>CAD</td>
<td>-0.90</td>
<td>$-3.11$</td>
<td>[-2.02, 1.90]</td>
<td>0.00</td>
<td>418</td>
</tr>
<tr>
<td>SWF</td>
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<td>$-2.97$</td>
<td>[-1.93, 1.97]</td>
<td>0.01</td>
<td>418</td>
</tr>
<tr>
<td>DNK</td>
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<td>$-2.41$</td>
<td>[-1.94, 2.02]</td>
<td>0.00</td>
<td>364</td>
</tr>
<tr>
<td>AUD</td>
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<td>$-2.91$</td>
<td>[-2.03, 1.86]</td>
<td>0.01</td>
<td>277</td>
</tr>
</tbody>
</table>

| Panel B. The subsample period |
| GDM    | -0.26   | $-1.44$ | [-1.92, 1.96] | 0.00  | 322 |
| BRP    | 0.56    | $-0.51$ | [-1.97, 1.95] | 0.00  | 322 |
| JPY    | -0.86   | $-1.48$ | [-1.90, 1.98] | 0.00  | 322 |
| CAD    | -0.70   | $-2.29$ | [-2.10, 1.73] | 0.00  | 322 |
| SWF    | -0.23   | $-1.17$ | [-1.97, 1.92] | 0.00  | 322 |
| DNK    | -0.17   | $-1.63$ | [-2.03, 1.93] | 0.00  | 262 |
| AUD    | -0.70   | $-1.81$ | [-1.88, 2.01] | 0.00  | 262 |

Note: The subsample excludes observations in 1980-87 out of the entire sample period. $\beta$ is estimated from the regression (1) using monthly foreign excess returns and forward premiums. The t-statistic of the estimated slope coefficient $\hat{\beta}$ (t-sta) in bold means that the null hypothesis, $\beta = 1$, is rejected at the 5% level based on the conventional critical value. We conduct the conditional test by Jasson and Moreira (2006) based on the bivariate regressions (1)-(2). The fourth column reports the 2.5 and 97.5% quantiles of the $t$ distribution obtained using the conditional test. Polk, Thompson, and Vuolteenaho (2006) develop the algorithms for the conditional test which are available on Polk’s homepage (http://personal.lse.ac.uk/POLK/research/work.htm). The interval of the quantiles in bold means that the null hypothesis is rejected at the 5% level. Data are collected from London close bid and ask prices and obtained from the database of Global Insight. Our sample includes spot and one-month U.S. dollar (USD) prices of the German mark (GDM), the British pound (BRP), the Japanese yen (JPY), the Canadian dollar (CAD), the Swiss franc (SWF), the Danish Krona (DNK) and the Australian dollar (AUD). We use monthly observations from 1975:1 to 2009:12 and select closing prices at the end of each month.
Table 2 Estimates of Contemporaneous Correlations ($\hat{\rho}$) and Regressor Persistence ($\hat{\varphi}$) in Foreign Exchange Markets

<table>
<thead>
<tr>
<th>Series</th>
<th>$\frac{\text{var}(y_t)}{\text{var}(x_t)}$</th>
<th>$\hat{\varphi}$</th>
<th>s.e.</th>
<th>$\hat{\rho}$</th>
<th>s.e.</th>
<th>$T$</th>
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<tr>
<td><strong>Panel A. The entire sample period : 1975-2009</strong></td>
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<tr>
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<td>0.06</td>
<td>418</td>
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<td>CAD</td>
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<td>0.03</td>
<td>0.09</td>
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<td>418</td>
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<td>-0.04</td>
<td>0.06</td>
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<td>0.08</td>
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<tr>
<td>AUD</td>
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<td>0.03</td>
<td>-0.04</td>
<td>0.09</td>
<td>277</td>
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<tr>
<td><strong>Panel B. The subsample period</strong></td>
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<td></td>
<td></td>
</tr>
<tr>
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<td>174.8</td>
<td>0.97</td>
<td>0.02</td>
<td>0.00</td>
<td>0.10</td>
<td>322</td>
</tr>
<tr>
<td>BRP</td>
<td>222.7</td>
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<td>322</td>
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<tr>
<td>JPY</td>
<td>439.0</td>
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<tr>
<td>CAD</td>
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<td>0.09</td>
<td>322</td>
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<td>SWF</td>
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<td>DNK</td>
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<td>0.12</td>
<td>0.01</td>
<td>0.07</td>
<td>262</td>
</tr>
<tr>
<td>AUD</td>
<td>286.0</td>
<td>0.97</td>
<td>0.02</td>
<td>-0.15</td>
<td>0.10</td>
<td>262</td>
</tr>
</tbody>
</table>

Note: $\rho$ and $\varphi$ are estimated from the bivariate regressions (1)-(2) using monthly foreign excess returns and forward premiums. See also Note in Table 1.
Table 3 Simulation Results: T=400

<table>
<thead>
<tr>
<th>parameter values</th>
<th>size distribution of $\beta$</th>
<th>$\hat{\rho}$</th>
<th>$\hat{\varphi}$</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>$5L$</td>
<td>$5R$</td>
<td>10%</td>
</tr>
<tr>
<td>Panel A. First specification of fundamentals</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>0.90 0.90 -0.80</td>
<td>0.06 0.04</td>
<td>-2.51 -1.41</td>
<td>0.79 2.81</td>
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<tr>
<td>0.90 0.95 -0.85</td>
<td>0.07 0.04</td>
<td>-1.42 -0.63</td>
<td>0.83 2.17</td>
</tr>
<tr>
<td>0.90 0.99 -0.90</td>
<td>0.08 0.03</td>
<td>-0.49 0.00</td>
<td>0.82 1.55</td>
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<tr>
<td>0.95 0.90 -0.80</td>
<td>0.06 0.04</td>
<td>-5.70 -3.57</td>
<td>0.63 4.55</td>
</tr>
<tr>
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<td>0.07 0.04</td>
<td>-3.22 -1.82</td>
<td>0.72 3.06</td>
</tr>
<tr>
<td>0.95 0.99 -0.90</td>
<td>0.07 0.03</td>
<td>-1.53 -0.68</td>
<td>0.72 2.00</td>
</tr>
<tr>
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<td>0.07 0.04</td>
<td>-6.51 -4.09</td>
<td>0.45 4.63</td>
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<tr>
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<td>0.07 0.04</td>
<td>-5.05 -3.08</td>
<td>0.58 3.92</td>
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<tr>
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<td>0.59 2.14</td>
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<tr>
<td>Panel B. Second specification of fundamentals</td>
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<td></td>
<td></td>
</tr>
<tr>
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<td>0.04 0.06</td>
<td>-0.57 -0.01</td>
<td>1.07 2.22</td>
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<tr>
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<td>0.04 0.06</td>
<td>-1.03 -0.30</td>
<td>1.13 2.68</td>
</tr>
<tr>
<td>0.90 0.99 0.00</td>
<td>0.03 0.08</td>
<td>-3.50 -1.65</td>
<td>1.86 5.78</td>
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<tr>
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<td>1.07 2.24</td>
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<tr>
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<td>0.04 0.07</td>
<td>-4.68 -2.40</td>
<td>1.88 6.49</td>
</tr>
</tbody>
</table>

Note: To generate the simulated data, the first specification uses equations (4) and (12) and the second specification uses (4) and (13), respectively. The $t$-test is conducted for the 5% significant level against either left-tail ($5L$) or right-tail ($5R$) alternatives. 10%, ..., 90% denote percentiles of the distribution of $\hat{\beta}$: for example, 10% means that the 10th percentile. $\hat{\rho}$ is the estimate of the contemporaneous correlation between $u_t$ and $v_t$ in the bivariate regressions (1)-(2) and $\hat{\varphi}$ is the estimated slope coefficient in regression (2).