Time-Varying Risk Premium in the Forward Exchange Rate

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Numerous studies have found that an implausibly high degree of risk aversion is needed to account for the forward exchange risk premium in an intertemporal asset-pricing model. However, when central banks' demand for foreign-exchange reserves is introduced into the standard representative agent model, both generalized-method-of-moments estimates and simulation results suggest that the forward risk premium is consistent with some reasonable values of the coefficient of relative risk aversion. Hence, the forward risk premium depends not only on consumption growth but also on monetary authorities' demand for foreign-exchange reserves.

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Hansen and Hodrick (1980) and subsequent studies have demonstrated that the forward exchange rates are biased predictors of the future spot exchange rates. Fama (1984) shows that a time-varying risk premium is needed to account for the behavior of the forward exchange rates. However, as Engel (1996) points out, modeling the forward risk premium in an intertemporal utility maximization framework has proved to be challenging. In particular, a common finding is that estimates of the coefficient of relative risk aversion are implausibly large. For instance, Mark (1985) reports estimates ranging from 0 to 50.38. Hodrick (1989) reports an estimate of 60.9 with U.S. consumption data and 2.15 with U.K. data. Kaminsky and Peruga (1990) report an estimate of 372.4.

These findings echo with the numerous studies on the equity premium puzzle in that a high degree of risk aversion is needed to account for the equity premium and the forward risk premium. Previous studies such as Campbell (2003), Campbell and Cochrane (1999), and Constantinides (1990) demonstrate that habit formation is able to explain the behavior of the equity premium. However, habit formation has been less successful in accounting for the behavior of the forward risk premium. For instance, Backus, Gregory and Telmer (1993) report an estimated curvature parameter of 107.39 in their model of the forward risk premium with a utility function that exhibits intertemporal nonseparability. Hence, resolutions of the equity premium puzzle are not necessarily applicable to the forward risk premium.

This paper attempts to address the shortcoming of the consumption-based model by examining another factor that affects the forward risk premium. In particular, U.S. government bonds are held by many countries and international organizations as foreign-exchange reserves. Because intertemporal consumption smoothing is unlikely to characterize the behavior of this source of demand for government bonds, numerous studies have examined the precautionary demand for foreign-exchange reserves. For instance, Aizenman and Marion (2002) report that reserve holdings depend on the size and volatility of international transactions, exchange-rate arrangement, sovereign risk and government's fiscal liabilities. Dooley, Lizondo and Mathieson (1989) and Eichengreen (1998) find that the currency composition of a country's reserves depends on the denominations of external debts, currency peg, trade shares and output shares of the reserve-currency countries.

Because central banks hold sizeable reserves, central banks as a group constitute a substantial source of demand for government bonds. The combined reserves held by all IMF member countries amount to over 3 trillion dollars in 2003. Moreover, the U.S. dollar accounts for over 60% of world reserves. In particular, U.S. Treasury bonds and notes are widely held as reserves. According to the Treasury Department, foreign official holdings of Treasury bills, bonds and notes amount to 1.2 trillion U.S. dollars in April 2005. Thus, the demand for government bonds comes not only from consumption-smoothing households but also from reserves-holding central banks. Since there are two types of participants in the bond market, the forward risk premium depends not only on the marginal utility of consumption but also on the marginal utility of reserves. Anecdotal evidence suggests that the foreign-exchange markets react to news about changes in official reserves. With the depreciation of the U.S. dollar against the major currencies in 2004, the financial press frequently reported speculations about the reduction of the dollar in central banks' reserves.[1]

In order to determine the empirical significance of foreign-exchange reserves, this paper introduces a representative foreign bond-holder into an otherwise standard representative household model. While the representative household's utility exhibits habit formation, the representative foreign bond-holder has a time-separable constant-relative-risk-aversion (CRRA) utility. The preference parameters are estimated using the continuous-updating generalized method of moments (GMM). The estimation procedure is applied to the monthly dollar exchange rates against the British pound, the Deutsche mark and the Japanese yen. Weak-identification statistics are used to assess the quality of the estimates.

The empirical section of this paper reports the estimation results of twenty specifications with various orthogonality conditions and instrument sets. Among the twenty GMM estimates, nineteen estimates of the coefficient of relative risk aversion of the representative household fall between 0.0003 and 1.6600. The largest estimate is 4.7292. Estimates of the subjective discount factor range from 0.9861 to 1.2008. Eighteen estimates of the habit parameter fall between 0.6569 and 0.9857. The other two estimates are 0.2434 and 0.2932. These estimates suggest a high degree of habit persistence. As for the representative foreign bold-holder, the estimates of the coefficient of relative risk aversion range from 0.9387 to 5.5344. Thus, the foreign bond-holder tends

[1] For example, *Economist* 12/04/04, 2/26/05; *Financial Times* 1/24/05, 3/8/05, 5/19/05; and *New York Times* 3/11/05.

to be more risk-averse than the representative household. Given the foreign central banks' precautionary motive for holding foreign-exchange reserves, their higher degree of risk aversion is economically justifiable. Overall, the GMM estimates suggest that the forward risk premiums are consistent with reasonable values of the preference parameters.

Because of the possible effect of weak identification in the GMM estimations, a simulation exercise is conducted to ascertain if the model is capable of generating some important sample moments of the forward risk premiums with preference parameters that are close to the GMM estimates. In particular, the coefficient of relative risk aversion of the representative household is fixed at either 0.50 or 1.00. The subjective discount factor is fixed at 0.99. The habit parameter is constrained to be between 0.00 and 0.95. The representative foreign bold-holder's coefficient of relative risk aversion is constrained to be between 1.00 and 1.80. Some combinations of these parameter values are capable of generating simulated moments that are statistically close to the mean, variance and first-order autocorrelation of the risk premium as well as the covariance between the forward discount and risk premium. Overall, these two econometric exercises suggest that the forward risk premiums are consistent with some reasonable values of the preference parameters.

The rest of this paper is organized into five sections. Section I discusses the time-series properties of the forward risk premium and motivates the modeling strategy adopted in this study. Section II presents the model of risk premium and derives the key moment conditions. Section III reports the GMM estimates along with the weak-identification statistics. Section IV presents the results of the simulation exercise. Section V contains some concluding remarks.

I. The Forward Risk Premium

Hansen and Hodrick (1980) and subsequent studies have demonstrated that the forward exchange rates are biased predictors of the future spot rates. Fama (1984) shows that a time-varying risk premium is needed to account for the behavior of the forward exchange rates. However, as Engel (1996) points out, modeling the behavior of the forward risk premium has proved to be challenging. In the standard consumption-based asset-pricing model, asset returns are determined by the representative agent's marginal rate of substitution of intertemporal consumption. Under the assumptions of time separability and constant relative risk aversion, the expected return from currency speculation satisfies an Euler condition of the following form.

$$0 = E_t \left[\left(\frac{F_t - S_{t+1}}{S_t} \right) \frac{P_t}{P_{t+1}} \left(\frac{C_{t+1}}{C_t} \right)^{-\gamma} \right] \tag{1}$$

where γ is the coefficient of relative risk aversion, F and S are the forward and spot exchange rates, P is the price level, and C represents the amount of current consumption.

A common method of estimating equation (1) is the generalized method of moments (GMM). However, a widespread finding is that estimates of the coefficient of relative risk aversion γ are implausibly large. For instance, Mark (1985) reports estimates ranging from 0 to 50.38. Hodrick (1989) reports an estimate of 60.9 with U.S. consumption data and 2.15 with U.K. data. Kaminsky and Peruga (1990) report an estimate of 372.4. Backus, Gregory and Telmer (1993) find that the coefficient of relative risk aversion is even higher when the representative agent's utility function exhibits intertemporal nonseparability. In particular, they report a GMM estimate of 52.79 for a time-separable utility function and an estimate of 107.39 for a utility function that allows for durability and habit.

One often-voiced criticism of the consumption-based model is that consumption growth is too stable to account for the volatility of the forward risk premium. Table I reports some summary statistics of the percentage change in U.S. real consumption. Seasonally adjusted U.S. consumption data were obtained from the Commerce Department's *National Income and Products Accounts*. The sample period is January 1976 to June 2004. Consumption is defined as the expenditures on non-durables and services. Real consumption is calculated by deflating with U.S. consumer price index, which is obtained from the IMF's *International Financial Statistics*. It can be seen that the standard deviation of real consumption growth is only 0.1630 while the standard deviations of the risk premiums of the three exchange rates range from 3.1777 to 3.5371. Thus, Engel (1996) attributes the implausibly large estimates of the coefficient of relative risk aversion to the disparity between the variances of consumption growth and forward risk premium.

Table I also presents some descriptive statistics of the other time series used in this paper. The five time series are the realized forward risk premium, currency depreciation rate, forward discount, U.S. real consumption growth, and net purchase of U.S. Treasury bills and notes by foreign official institutions.

Monthly spot and forward exchange rates were obtained from Datastream. The sample period is January 1976 to December 1998 for the Deutsche mark. The sample period is January 1976 to June 2004 for the British pound and the Japanese yen. The exchange rates are expressed as the numbers of U.S. dollars per foreign currency unit. This unit of the exchange rates is chosen because the following model considers the U.S. as the home country and U.S. consumption is used to construct the pricing kernel. Uppercase symbols S_t and F_t denote the levels of the spot exchange rate and forward rate respectively. Lowercase symbols s_t and f_t denote the exchange rate and forward rate in logarithm and multiplied by 100. Thus, the first differences of the lowercase variables are approximately equal to the percentage changes over one month.

The realized risk premium is defined as the difference between the logarithm of the one-month forward rate and the logarithm of the spot exchange rate in the following month $f_t - s_{t+1}$. The currency depreciation rate is defined as the change in the logarithm of the spot exchange rate $s_{t+1} - s_t$. The forward discount is defined as the difference between the logarithm of the one-month forward rate and the logarithm of the spot exchange rate $f_t - s_t$. It can be seen from Table 1 that the standard deviations of the risk premiums are slightly higher than those of the currency depreciation rates, while the forward discounts exhibit skewness, kurtosis and high autocorrelation.

Data on foreign official purchases of U.S. government bonds were obtained from the Treasury Department's *Net Purchases of U.S. Treasury Bonds and Notes by Major Foreign Sector*. This series is deflated with U.S. consumer price index. While real consumption growth has the lowest standard deviation among all time series, net purchase of Treasury bills is the most volatile series. It can be seen from Table 1 that the standard deviation of net purchase is 7.1828, which is twice as large as the standard deviations of the forward risk premiums. Thus, the demand for foreign-exchange reserves could potentially account for the volatility of the forward risk premium.

Because intertemporal consumption smoothing is unlikely to characterize the demand for U.S. government bonds as foreign-exchange reserves, the following model aims to capture the interaction between consumption-smoothing households and reserves-holding central banks. The forward risk premium is therefore a function of the marginal utility of consumption and the marginal utility of reserves held by foreign official institutions.

II. Model of Time-Varying Forward Risk Premium

The basic set-up is a standard representative agent model with the U.S. as the home country. The representative household maximizes its utility subject to the budget constraint. As in Backus, Gregory and Telmer (1993), the representative agent's utility function exhibits time nonseparability. In particular, the representative agent's expected utility takes the following form.

$$\begin{aligned} U_t &= E_t \sum_{k=0}^{\infty} \beta^{t+k} u\left(d_{t+k}\right) \\ u\left(d_t\right) &= \frac{1}{1-\gamma}\left(C_t - \rho C_{t-1}\right)^{1-\gamma} \end{aligned} \tag{2}$$

where C_t is the representative agent's consumption at time t. The subjective discount factor is β. The curvature parameter of the utility function is γ. The degree of habit persistence is denoted by ρ. Positive values of ρ indicate habit persistence while negative values indicate durability. When ρ is equal to zero, utility function (2) becomes a time-separable constant-relative-risk-aversion utility function, which has been used in such previous studies as Mark (1985) and Hodrick (1989).

The household receives endowment Y_t. It holds three types of assets: domestic bonds b_t, foreign bonds $\tilde{b}_t$, and other securities z_t. The budget constraint can be written as follows.

$$P_t Y_t + b_{t-1}\left(1 + i_{t-1}\right) + \tilde{b}_{t-1}\left(1 + \tilde{\imath}_{t-1}\right) S_t + q_t z_{t-1} = P_t C_t + b_t + S_t \tilde{b}_t + q_t z_t \tag{3}$$

The exchange rate S_t is expressed as the U.S. dollar price of one unit of foreign currency. The domestic and foreign interest rates are denoted by i_t and $\tilde{\imath}_t$ respectively. The prices of other securities are denoted by q_t.

In addition to the consumption-smoothing household, foreign central banks also participate in the U.S. bond market. Out of precautionary motive, foreign central banks hold U.S. government bonds as reserves, which are denoted by R_t. Moreover, foreign central banks are assumed to hold U.S. government bonds only. This assumption can be justified by the fact that U.S. government bonds are the preferred vehicles of foreign-exchange reserves. According to the Treasury Department's *Net Purchases of U.S. Treasury Bonds and Notes by Major Foreign Sector*, the average size of foreign official net purchase of U.S. Treasury bonds and notes is sixty times larger than the average size of foreign official net purchase of U.S. equities. It is further assumed that foreign central banks can be modeled as a single representative bond-holder, whose utility depends on the amount of U.S. government bonds held as reserves in real terms. Moreover, the representative bond-holder's utility function is assumed to be time-separable.

$$v\left(X_t\right) = \frac{1}{1-\alpha} X_t^{1-\alpha} \tag{4}$$

where $X_t = \frac{R_t}{P_t}$ is the amount of reserves in real terms. The foreign bond-holder's coefficient of relative risk aversion is denoted by α.

The amount of reserves evolves according to the following equation.

$$R_{t+1} = \left(1 + i_t\right) R_t + \Delta R_{t+1} \tag{5}$$

Previous studies have shown that changes in reserves depend on a variety of factors. Aizenman and Marion (2002) report that reserve holdings depend on the size and volatility of international transactions, exchange-rate arrangement, sovereign risk and government's fiscal liabilities. Dooley, Lizondo and Mathieson (1989) and Eichengreen (1998) find that the currency composition of a country's reserves depends on the denominations of external debts, currency peg, trade shares and output shares of the reserve-currency countries. Given these findings, the change in reserves ΔR_{t+1} is assumed to be exogenous in this model.

Given the two sources of demands for domestic bonds, the bond-market equilibrium condition plays an important role in the determination of the forward risk premium.

$$\bar{B}_t = b_t + R_t \tag{6}$$

where $\bar{B}_t$ is the amount of outstanding domestic government bonds, b_t and R_t represent the demands from the representative household and the foreign bond-holder respectively.

The following Lagrangian simultaneously maximizes the utilities of the representative household and the representative foreign bond-holder.

$$\mathcal{L}=E_0\sum_{t=0}^{\infty}\beta^t\left\{\begin{array}{c} u\left(d_t\right)+v\left(X_t\right)+\mu_t\left[\bar{B}_t-b_t-R_t\right] \\ +\lambda_t\left[\begin{array}{c} P_tY_t+b_{t-1}\left(1+i_{t-1}\right)+\tilde{b}_{t-1}\left(1+\tilde{\imath}_{t-1}\right)S_t \\ +q_tz_{t-1}-P_tC_t-b_t-S_t\tilde{b}_t-q_tz_t \end{array}\right] \end{array}\right\} \tag{7}$$

The first-order conditions with respect to C_t, X_t, b_t, and $\tilde{b}_t$ can be used to derive the pricing equations for the domestic and foreign interest rates as follows.

$$\frac{1}{1+i_t} = E_t\left[\beta\frac{P_t}{P_{t+1}}\frac{\partial U_{t+1}/\partial C_{t+1}}{\partial U_t/\partial C_t+v'\left(X_t\right)}\right] \tag{8}$$

$$\frac{1}{1+\tilde{\imath}_t} = E_t\left[\beta\frac{S_{t+1}}{S_t}\frac{P_t}{P_{t+1}}\frac{\partial U_{t+1}/\partial C_{t+1}}{\partial U_t/\partial C_t}\right] \tag{9}$$

where $\partial U_t/\partial C_t=\left(C_t-\rho C_{t-1}\right)^{-\gamma}-\beta\rho E_t\left(C_{t+1}-\rho C_t\right)^{-\gamma}$ is the marginal utility of consumption.

Because of the foreign central banks' participation in the domestic bond market, the marginal utility of reserves $v'\left(X_t\right)$ is one of the determinants of the domestic interest rate. In particular, the domestic interest rate is negatively related to the amount of reserves as $v''\left(X_t\right)$ is negative. In other words, the return on domestic bonds will decrease as the foreign central banks hold more reserves.

Given the domestic and foreign interest rates, the covered interest parity can be used to calculate the forward rate.

$$\frac{F_t}{S_t}\left(1+\tilde{\imath}_t\right)=\left(1+i_t\right) \tag{10}$$

In particular, substituting the expressions for the domestic and foreign interest rates in equations (8) and (9) into the covered interest parity (10) yields the following pricing equation for the forward risk premium.

$$E_t\left[\frac{P_t}{P_{t+1}}\frac{\partial U_{t+1}/\partial C_{t+1}}{\partial U_t/\partial C_t}\left(\frac{S_{t+1}}{F_t}-\frac{\partial U_t/\partial C_t}{\partial U_t/\partial C_t+v'\left(X_t\right)}\right)\right]=0 \tag{11}$$

Equation (11) shows that the forward risk premium depends on the marginal utility of consumption as well as the marginal utility of reserves. When the marginal utility of reserves is equal to zero, equation (11) reduces to a moment condition that can be derived from the standard representative agent model. In other words, reserve holdings are irrelevant for pricing the forward risk premium when the foreign bond-holder's coefficient of relative risk aversion α approaches infinity. The following GMM estimations and simulation exercises suggest that the marginal utility of reserves is positive because all estimates of α lie between 0.9387 and 5.5344. Hence, equation (11) provides an explanation of the volatility of the forward risk premium. Consumption smoothing implies that the marginal utility of consumption is relatively stable over time. As a result, consumption growth alone is unable to account for the volatility of the forward risk premium. Equation (11) shows that this shortcoming can be addressed by introducing the demand for official reserves into an otherwise standard representative agent model because changes in reserve holdings are much more volatile than consumption growth. Moreover, equation (11) is also consistent with the anecdotal evidence of the foreign-exchange markets' reactions to news about changes in official reserves.

Given the assumption that the representative foreign bond-holder does not participate in the stock market, the pricing equation for stock return is the same as in the standard representative agent model.

$$1 = E_t \left[\beta \left(1 + r_{t+1}\right) \frac{P_t}{P_{t+1}} \frac{\partial U_{t+1} / \partial C_{t+1}}{\partial U_t / \partial C_t} \right] \tag{12}$$

where r_{t+1} denotes the rate of return on stock realized at time $t + 1$. In the GMM estimations below, equation (12) will be included as a moment condition in some of the specifications. Stock return will be taken to be the value-weighted NYSE/AMEX index return obtained from the Center for Research in Security Prices (CRSP).

The rationale for including equation (12) as a moment condition is to shed light on the parallel between the equity premium and the forward exchange risk premium. Previous studies such as Campbell (2003), Campbell and Cochrane (1999), and Constantinides (1990) demonstrate that habit formation is able to explain the behavior of the equity premium. These studies provide support for the stock return equation (12). However, habit formation has been less successful in accounting for the behavior of the forward risk premium. Backus, Gregory and Telmer (1993) show that a large curvature parameter is needed even with habit formation. This paper provides an explanation of the difference between the equity premium and the forward exchange risk premium. That is, in addition to habit formation, the effect of foreign-exchange reserves is needed to account for the forward risk premium as shown in equation (11).

The following two econometric exercises aim to evaluate moment conditions (8), (11) and (12). First, the preference parameters are estimated using the continuous-updating generalized method of moments (GMM). Second, a simulation exercise ascertains whether the sample moments of the risk premiums can be generated from parameter values that are close to the GMM estimates.

III. GMM Estimation

The pricing equations derived in the previous section can be used to generate the orthogonality conditions for the GMM estimations of the preference parameters. There are four preference parameters. The representative household's coefficient of relative risk aversion is denoted by γ, the representative bond-holder's coefficient of relative risk aversion is denoted by α, the habit parameter is denoted by ρ, and the subjective discount factor is denoted by β. The parameters are constrained to be the same for the three exchange rates so that the parameter space is four dimensional.

In this paper, the GMM estimates minimize the robust continuous-updating GMM objective function:

$$S(\delta) = \left[\frac{1}{\sqrt{T}} \sum_{t=1}^{T} \phi_t(\delta) \right]' V(\delta)^{-1} \left[\frac{1}{\sqrt{T}} \sum_{t=1}^{T} \phi_t(\delta) \right] \tag{13}$$

where $\delta = (\gamma, \alpha, \rho, \beta)'$ is the vector of parameters to be estimated; $\phi_t(\delta) = h(Y_t, \delta) \otimes Z_t$, where $h(Y_t, \delta)$ is the vector of moment conditions and Z_t is the vector of instruments; and $V(\delta) = \frac{1}{T} \sum_{t=1}^{T} \left[\phi_t(\delta) - \bar{\phi}_t(\delta) \right] \left[\phi_t(\delta) - \bar{\phi}_t(\delta) \right]'$ is the robust covariance matrix. This choice of the weighting matrix facilitates the use of weak-identification statistics developed by Stock and Wright (2000).

Conventional GMM statistics such as the J-statistic and confidence interval assume that GMM solves a locally quadratic minimization problem, but this assumption does not always hold in practice. For instance, Ferson and Constantinides (1991) show that the GMM objective functions are clearly non-quadratic in the habit parameter in their estimation of the consumption-based asset-pricing model. Ma (2002) shows that the locally quadratic assumption is seriously violated in the GMM estimation of the new Keynesian Phillips curve. Yogo (2003) finds that weak identification is an important consideration in the estimation of the elasticity of intertemporal substitution. Monte Carlo studies such as Hansen, Heaton and Yaron (1996) have shown that asymptotic normality provides poor approximations to the finite-sample distributions of GMM estimators. Stock and Wright (2000) attribute the discrepancy to weak identification and develop test statistics that are applicable even when some of the parameters are weakly identified. Because conventional GMM statistics and weak-identification statistics often lead to different econometric inference, both sets of statistics are reported in this section.

This section will report the estimation results of twenty specifications with various instrument sets and moment conditions. Given the focus on the forward risk premium, equation (11) is the main moment condition of this estimation exercise. In addition, the interest rate equation (8) is also included as an orthogonality condition. Since the subjective discount factor β appears in equation (11) only multiplicatively, the inclusion of equation (8) helps to identify the parameters. Specifications that consist of these two moment conditions are referred to as the X1 specifications. There are eight X1 specifications with various instrument sets. The instrument sets differ in the included instruments as well as their lag lengths. A constant is included as an instrument in all specifications. Other instruments include lagged consumption growth $\frac{C_t}{C_{t-1}}$, risk premium $\frac{S_t}{F_{t-1}}$, and forward discount $\frac{S_t}{F_t}$. These instruments have been commonly used in previous studies such as Backus, Gregory and Telmer (1993), Hodrick (1989) and Mark (1985).

The first four specifications use the first lag of the included variables as instruments. Specifications that use the first lag as instrument are indicated by 'FL' in their labels. Thus, the first four specifications are referred to as specifications X1-FL-1 to X1-FL-4 in Panel A of Table II. Because of the possible effect of temporal aggregation bias, four other specifications use the second lag as instrument. These specifications are indicated by 'SL' in their labels. In particular, the four specifications are referred to as specifications X1-SL-1 to X1-SL-4 in Panel B of Table II.

Twelve additional specifications are generated from the interest rate equation (8), risk premium equation (11), and stock return equation (12). The label 'X2' indicates the use of these three moment conditions. In addition to instruments used for the X1 specifications, lagged stock return is also included as an instrument in some of specifications. Six of X2 specifications use the first lag as instrument. They are referred to as specifications X2-FL-1 to X2-FL-6 in Panel A of Table III. The other six X2 specifications use the second lag as instrument. They are referred to as specifications X2-SL-1 to X2-SL-6 in Panel B of Table III.

Overall, the GMM point estimates suggest that the moment conditions can be satisfied with reasonable values of the preference parameters. There is no systemic difference between the estimates of the X1 and X2 specifications. Among the twenty specifications, nineteen estimates of the coefficient of relative risk aversion of the representative household γ fall between 0.0003 and 1.6600. The largest estimate is 4.7292. Contrary to the finding of Backus, Gregory and Telmer (1993), there is no substantial increase in the relative risk aversion coefficient when time nonseparability is introduced. This is because all estimates of the habit parameter are positive and imply a high degree of habit persistence. In particular, eighteen estimates of the habit parameter ρ fall between 0.6569 and 0.9857. The other two estimates are 0.2434 and 0.2932. These estimates are very close to those reported in Ferson and Constantinides (1991). Hansen and Jagannathan (1991) demonstrate that habit persistence tends to increase the variability of the intertemporal marginal rate of substitution. It is therefore not surprising that a high degree of habit persistence is needed to account for the high volatility of the forward risk premium. Other related studies such as Fuhrer (2000) find that a habit parameter of 0.6 provides a good match between a sticky-price model and consumption data.

Estimates of the subjective discount factor β range from 0.9861 to 1.2008. Although some estimates are larger than one, the estimates are quite close to the *a priori* reasonable values. As for the representative foreign bold-holder, the estimates of the coefficient of relative risk aversion α range from 0.9387 to 5.5344. Both conventional standard errors and concentrated S-sets suggest that a majority of the GMM estimates of α are statistically non-zero. According to these point estimates, the foreign bondholder is slightly more risk-averse than the representative household. Given the foreign central banks' precautionary motive for holding foreign-exchange reserves, their higher degree of risk aversion is economically justifiable.

Because conventional confidence ellipse often mistakenly assumes that the objective function is locally quadratic around the GMM estimate, Tables II and III also report the concentrated S-sets of the preference parameters. The concentrated S-sets can be interpreted as confidence intervals that are robust to weak identification. According to Theorem 3 in Stock and Wright (2000), if a parameter θ is well-identified, then $S\left(\delta_0, \hat{\theta}\left(\delta_0\right)\right) \xrightarrow{D} \chi^2_{k-n}$, where $S\left(\delta_0, \hat{\alpha}\left(\delta_0\right)\right)$ denotes the concentrated objective function evaluated at δ_0, k is the dimension of the weighting matrix, and n is the dimension of θ. In other words, a concentrated S-set contains all parameter values such that the continuous-updating objective function is smaller than the χ^2_{k-n} critical value. It can be seen from Tables II and III that some of the concentrated S-sets are very wide. The large concentrated S-sets imply that the GMM estimates are not very informative of the precise values of the parameters because a large set of parameter values satisfies the moment conditions. Thus, the preference parameters are likely to be weakly identified.

With four preference parameters, the full S-set is four dimensional and hence cannot be graphically displayed. Instead, Figures 1 and 2 display the two-dimensional concentrated S-sets for the X1-FL specifications and the X2-SL specifications. The other two sets of figures for the X1-SL and X2-FL specifications are similar and are available upon request. The two dimensions of the concentrated S-sets correspond to the coefficients of relative risk aversion of the representative household γ and the representative foreign bond-holder α. The two-dimensional concentrated S-sets are also constructed according to Theorem 3 in Stock and Wright (2000). In other words, a 95% concentrated S-set contains values of α and γ such that the objective function is less than the 95% χ^2_{k-2} critical value, where k is equal to the dimension of the weighting matrix.

To illustrate the effect of weak identification, Figures 1 and 2 juxtapose the conventional 95% confidence ellipses with the 95% concentrated S-sets. As an indication of weak identification, the S-sets are much larger than the conventional confidence ellipses. In their estimation of the consumption-based capital-asset-pricing model, Stock and Wright (2000) present some S-sets that are very similar to the ones in Figures 1 and 2. Because preference parameters α and γ enter the risk premium equation (11) as exponents in a ratio, proportional increases in both α and γ may have little impact on the value of the continuous-updating objective function (13). As a result, some of the concentrated S-sets contain large sets of parameter values. It is in this sense that the parameters are weakly identified. Even though the point estimates are small, the orthogonality conditions can also be satisfied with large parameter values. This is a common feature of power utility. In his estimation of the Euler condition for the forward risk premium, Mark (1985, p.15) points out that the objective functions are flat around the GMM estimates. Thus, the S-sets would also have been large if applied to Mark's specifications. Ferson and Constantinides (1991) also show that the GMM objective functions are clearly non-quadratic in the habit parameter in their estimation of the consumption-based asset-pricing model.

Many of the concentrated S-sets of the representative foreign bond-holder's coefficient of relative risk aversion α diverge to infinity. Infinitely large S-sets are not uncommon for the consumption-based models. For instance, Yogo (2003) reports some infinitely large S-sets for the elasticity of intertemporal substitution. However, an infinitely large α implies that the marginal utility of reserves is equal to zero. That is, as α approaches infinity, the marginal utility of reserves becomes irrelevant and the model reduces to the standard representative agent model. To ascertain whether α is indeed infinitely large, the twenty specifications are re-estimated with the marginal utility of reserves set to zero. That is, the preference parameters are estimated with the following three moment conditions.

$$\frac{1}{1+i_t} = E_t\left[\beta\frac{P_t}{P_{t+1}}\frac{\partial U_{t+1}/\partial C_{t+1}}{\partial U_t/\partial C_t + v'\left(X_t\right)}\right] \tag{14}$$

$$0 = E_t\left[\frac{P_t}{P_{t+1}}\frac{\partial U_{t+1}/\partial C_{t+1}}{\partial U_t/\partial C_t}\left(\frac{S_{t+1}-F_t}{F_t}\right)\right] \tag{15}$$

$$1 = E_t\left[\beta\left(1+r_{t+1}\right)\frac{P_t}{P_{t+1}}\frac{\partial U_{t+1}/\partial C_{t+1}}{\partial U_t/\partial C_t}\right] \tag{16}$$

Table IV reports the estimation results. The point estimates of the representative household's coefficient of relative risk aversion become much larger. Seventeen of the twenty estimates of γ lie between 10.8825 and 204.9652. The other three estimates are 0.8155, 3.2991 and 5.4532. This finding should not be surprising as previous studies such as Mark (1985), Hodrick (1989), and Kaminsky and Peruga (1990) have shown that the forward risk premium equation (15) can be satisfied with large values of γ. Except for specification X2-SL-4, the lower end point of the concentrated S-set of γ is larger than 1.4. This suggests that moment conditions (14), (15) and (16) cannot be satisfied with small values of γ when the effect of reserves is omitted from the model.

Therefore, the implausibly large estimates of γ can be attributed to the omission of the marginal utility of foreign-exchange reserves.

Even though the parameters are weakly identified, these estimation results suggest that the forward risk premium is consistent with reasonable values of the coefficient of relative risk aversion once the effect of foreign-exchange reserves is taken into account. Figures 1 and 2 show that the concentrated S-sets contain economically plausible values of the representative household's relative risk aversion coefficient, although weak identification implies that large parameter values cannot be ruled out. In other words, a large set of parameter values satisfies the moment conditions.

Given the possibility of weak identification, the J-statistic does not necessarily provide an accurate assessment of the moment conditions. In particular, Hansen, Heaton and Yaron (1996) have shown that asymptotic normality provides poor approximations to the finite-sample distributions of GMM estimators. Among the forty specifications reported, only specifications X1-FL-3 in Table II and X2-FL-3 in Table III produce J-statistics with p-values that are less than 10%. Interestingly, these two specifications use the same instrument set, which consists of a constant, the first lag of consumption growth and the first lag of forward discount. As Engel (1996) points out, previous studies often reject specifications that include lagged forward discount as an instrument. This feature could simply reflect the time-series properties of the forward discount. It can be seen from Table I that the forward discount exhibits high skewness, kurtosis and autocorrelation. Because of the high autocorrelation, lagged forward discount is highly correlated with contemporaneous forward discount, which is used to derive the risk premium equation (11). Thus, the low p-values may simply reflect the persistence of the forward discount.

IV. Simulation

Because the parameters are not strongly identified in the above GMM estimations, a simulation exercise is conducted to evaluate the model. The design of the simulation exercise is similar to those in Backus, Gregory and Telmer (1993), Constantinides (1990), Heaton (1993), and Mehra and Prescott (1985). The aim is to ascertain whether the GMM estimates obtained above are capable of generating some key sample properties of the forward risk premiums. While the above GMM estimations constrain the parameters to be the same across the three exchange rates, this simulation exercise is done separately for each of the three exchange rates.

Rearranging equation (11) shows that the forward rate can be generated according to the following equation.

$$F_t = \frac{E_t\left[S_{t+1}\frac{P_t}{P_{t+1}}\frac{\partial U_{t+1}/\partial C_{t+1}}{\partial U_t/\partial C_t}\right]}{E_t\left[\frac{P_t}{P_{t+1}}\frac{\partial U_{t+1}/\partial C_{t+1}}{\partial U_t/\partial C_t+v'(X_t)}\right]} \tag{17}$$

Equation (17) shows that the forward rate is equal to a ratio of expectations of two functions of the joint stochastic process of consumption growth, inflation and exchange-rate movement, which are taken to be the three state variables in this simulation exercise. The three-variable joint stochastic process is approximated by an eight-state Markov chain. In particular, each of the three state variables may take on two values: high and low. A state variable is in the high state when it is above its sample mean, and it is in the low state otherwise. The transition probabilities are estimated from the transition frequencies in the data. The eight states of the Markov chain correspond to the eight combinations of the values that the state variables may take on. With habit formation, the marginal utility of consumption at time $t+1$ depends on consumptions at time t, $t+1$ and $t+2$. As a result, each of the two expected values in equation (17) is equal to a probability-weighted average of the 16 possible combinations of the values that the state variables may take on. The simulated forward rate is equal to a ratio of the two expected values.

The simulated forward risk premium is generated from the simulated forward rate and the simulated exchange rate, as the change in exchange rate is one of three state variables. Four simulated moments are calculated from 2000 replications of the sample lengths: 246 for the Deutsche mark and 324 for the pound and the yen. The four simulated moments are the covariance between the risk premium and forward discount, the mean, variance and coefficient of first-order autocorrelation of the forward risk premium. The standard deviations of the simulated moments are also calculated from the 2000 replications. In Table V, asterisks beside the simulated moments indicate that the sample moments are within two standard deviations of the simulated moments.

Because the possible effect of weak identification in the above GMM estimations, the aim of this simulation exercise is to ascertain whether the sample moments can be generated from parameter values that are close to the GMM estimates reported above. Since nineteen of the twenty estimates of the coefficient of relative risk aversion of the representative household fall between 0.0003 and 1.6600, the representative household's curvature parameter γ is fixed at either 0.5 or 1. The subjective discount factor β is fixed at 0.99 as the GMM estimates range from 0.9861 to 1.2008. Because all twenty GMM estimates of the habit parameter are positive, the habit parameter ρ is chosen from the unit interval [0, 1]. Since the GMM estimations suggest that the representative foreign bond-holder is slightly more risk-averse than the representative household, the representative foreign bond-holder's coefficient of relative risk aversion α is chosen from the interval [1.0, 1.8].

Table V reports the simulation results of various combinations of the preference parameters. Different parameter values are needed to match the sample moments of the three exchange rates. For the Japanese yen, even when the habit parameter is fixed at zero, all four sample moments lie within two standard deviations of the respective simulated moments when $\gamma = 1.00$ and $\alpha = 1.20$. For the other two exchange rates, a high degree of habit persistence is needed to match the sample moments. This is consistent with the fact that eighteen of the twenty estimates of the habit parameter fall between 0.6569 and 0.9857. These simulation results provide support for the GMM point estimates reported above. Hence, although the GMM estimates are weakly identified, they are capable of generating some important sample properties of the forward risk premiums.

V. Conclusion

This paper presents evidence that the forward exchange risk premium is consistent with reasonable values of the coefficient of relative risk aversion. This result is quite encouraging in light of the previous estimates of the coefficient of relative risk aversion. Once the demand for foreign-exchange reserves is introduced into an otherwise standard representative agent model, the key sample properties of the risk premium can be generated by assuming that the representative household has a coefficient of relative risk aversion of no more than one and that the representative foreign bond-holder is slightly more risk-averse. Hence, this paper demonstrates that the risk premium depends not only on consumption growth but also on monetary authorities' demand for foreign-exchange reserves.

One possible direction of future research is to introduce the demand for foreign-exchange reserves into a general equilibrium model of exchange rate. Chari, Kehoe and McGrattan (2002) find that a high degree of risk aversion is necessary to account for the volatility and persistence of the real exchange rates. Perhaps one could investigate whether the demand for foreign-exchange reserves can improve the goodness of fit of a quantitative general equilibrium model of exchange rate. Future work could also attempt to address the weak identification of the preference parameters. To deal with the weak identification of the GMM estimates, an identification scheme is necessary to produce more precise estimates of the preference parameters.

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Table I
Summary statistics of the main time series

Monthly spot exchange rates and forward rates, s_t and f_t, are in logarithm and multiplied by 100. The exchange rates are expressed as the numbers of U.S. dollars per currency unit. The sample period is January 1976 to December 1998 for the Deutsche mark. The sample period is January 1976 to June 2004 for all other time series. Consumption and net purchase of Treasury bills are deflated by U.S. consumer price index.

time series	mean	standard deviation	skewness	kurtosis	coefficient of first-order autocorrelation
Realized forward risk premium $f_t - s_{t+1}$					
British pound	-0.1819	3.1777	0.0722	4.5958	0.0744
Deutsche mark	0.1056	3.5371	0.2924	4.0051	0.0251
Japanese yen	0.0978	3.5715	-0.3918	4.2343	0.0572
Monthly change in the exchange rate $s_{t+1} - s_t$					
British pound	-0.0081	3.1349	-0.1235	4.7805	0.0492
Deutsche mark	0.0887	3.4239	-0.0982	3.6775	0.0074
Japanese yen	0.2006	3.5268	0.4571	4.3394	0.0253
Forward discount $f_t - s_t$					
British pound	-0.1900	0.3356	-7.3923	100.5739	0.3187
Deutsche mark	0.1943	0.6063	10.2962	142.6883	0.2631
Japanese yen	0.2984	0.3109	-1.1720	17.6489	0.4486
Percentage change in real consumption of non-durables and services					
U.S. consumption	0.2159	0.1630	-0.3099	3.7862	-0.0456
Percentage change in net purchase of U.S. Treasury bills and notes by foreign official institutions					
Foreign-exchange reserves	0.7683	7.1828	0.0410	4.3460	-0.0069

Table II
GMM estimates of the preference parameters

Preference parameters are estimated using the continuous-updating GMM with the interest rate equation (8) and risk premium equation (11) as the moment conditions. A constant is included as an instrument in all specifications. The second column indicates the other included instruments: CG stands for consumption growth $\frac{C_t}{C_{t-1}}$, RP risk premium $\frac{S_t}{F_{t-1}}$, and FD forward discount $\frac{S_t}{F_t}$. In parentheses below the GMM estimates are the standard errors calculated from the robust covariance matrix in the continuous-updating objective function. Open intervals below the standard errors are the concentrated S-sets, which contain parameter values such that the continuous-updating objective functions are smaller than the 95% χ^2_{k-3} critical values. The degree of freedom k is equal to the dimension of the weighting matrix, which is equal to the number of orthogonality restrictions. The J-statistic is the value of the continuous-updating objective function at the GMM estimate. The corresponding p-value is reported in parentheses below the J-statistic.

Panel A: Instrument sets consist of a constant and the first lag of the included variables

specification	instrument set includes of the first lag of the following variables	GMM estimates, standard errors and 95% concentrated S-sets				J-statistic	number of orthogonality restrictions
		$\hat{\gamma}_{GMM}$	$\hat{\alpha}_{GMM}$	$\hat{\rho}_{GMM}$	$\hat{\beta}_{GMM}$		
X1-FL-1	CG	0.0089	1.1304	0.6569	0.9974	7.5334	8
		(0.1617)	(1.0011)	(2.6167)	(0.0019)	(0.4803)	
		(-4.5977, 0.5996)	(1.0033, ∞)	(-∞, 0.9695)	(0.9971, 1.5224)		
X1-FL-2	CG, RP	1.0593	1.2083	0.8132	1.0460	19.4177	20
		(1.1133)	(0.2672)	(0.0807)	(0.0441)	(0.4948)	
		(0.5413, 1.8096)	(1.0226, 1.8096)	(0.7513, 0.9129)	(0.9729, 1.3386)		
X1-FL-3	CG, FD	0.0003	1.4280	0.9857	0.9978	35.2067	20
		(0.0007)	(0.1989)	(0.0060)	(0.0008)	(0.0190)	
		(-0.9154, 0.0946)	$\varnothing$	$\varnothing$	$\varnothing$		
X1-FL-4	CG, RP, FD	1.6600	1.6216	0.9224	0.9331	24.2624	32
		(0.0094)	(0.0064)	(0.0041)	(0.0115)	(0.8348)	
		(0.9140, 5.0190)	(0.9476, ∞)	(0.8894, 0.9749)	(0.8405, 1.0245)		

Panel B: Instrument sets consist of a constant and the second lag of the included variables

specification	instrument set includes of the second lag of the following variables	GMM estimates, standard errors and 95% concentrated S-sets				J-statistic	number of orthogonality restrictions
		$\hat{\gamma}_{GMM}$	$\hat{\alpha}_{GMM}$	$\hat{\rho}_{GMM}$	$\hat{\beta}_{GMM}$		
X1-SL-1	CG	0.5152	1.1672	0.8560	1.0519	6.4425	8
		(3.0241)	(0.4388)	(0.3212)	(0.0944)	(0.5978)	
		(-0.3629, 1.3095)	(0.9795, ∞)	(0.8325, 0.8959)	(0.9895, 1.1400)		
X1-SL-2	CG, RP	0.9765	1.9299	0.8005	1.2008	14.1342	20
		(0.1365)	(1.0129)	(0.0458)	(0.0784)	(0.8236)	
		(0.4728, 1.9948)	(1.1227, ∞)	(0.7752, 0.9595)	(0.9873, 1.2239)		
X1-SL-3	CG, FD	0.8704	1.3183	0.9213	0.9861	19.5117	20
		(0.2485)	(0.0962)	(0.0204)	(0.0383)	(0.4888)	
		(0.6193, 2.3124)	(0.8904, ∞)	(0.9063, 0.9533)	(0.9524, 1.0411)		
X1-SL-4	CG, RP, FD	0.3647	0.9387	0.9588	1.0013	28.7170	32
		(0.1429)	(0.0790)	(0.0014)	(0.0094)	(0.6335)	
		(0.1942, 0.4551)	(0.8242, ∞)	(0.9497, ∞)	(0.9517, 1.1876)		

Table III
GMM estimates of the preference parameters

Preference parameters are estimated using the continuous-updating GMM with the interest rate equation (8), risk premium equation (11) and stock return equation (12) as the moment conditions. Instrument SR denotes stock return. Other notes on methodology can be found in Table II.

Panel A: Instrument sets consist of a constant and the first lag of the included variables

specification	instrument set includes of the first lag of the following variables	GMM estimates, standard errors and 95% concentrated S-sets $\hat{\gamma}_{GMM}$	$\hat{\alpha}_{GMM}$	$\hat{\rho}_{GMM}$	$\hat{\beta}_{GMM}$	J-statistic	number of orthogonality restrictions
X2-FL-1	CG	0.0594	1.8807	0.2434	0.9974	13.9759	10
		(0.2643)	(1.3608)	(2.1555)	(0.0005)	(0.1741)	
		(0.0370, 0.0814)	(1.2444, ∞)	(-∞, 0.4472)	(0.9973, 0.9974)		
X2-FL-2	CG, RP	0.2778	5.3011	0.8897	1.0323	28.4920	25
		(0.0163)	(3.0631)	(0.0095)	(0.0145)	(0.2857)	
		(0.1599, 0.8192)	(1.0252, ∞)	(0.8617, 0.9516)	(0.9854, 1.0806)		
X2-FL-3	CG, FD	0.0003	1.5303	0.9854	0.9973	50.5458	25
		(0.0005)	(0.1710)	(0.0035)	(0.0003)	(0.0018)	
		(-0.1721, 0.1075)	$\varnothing$	$\varnothing$	$\varnothing$		
X2-FL-4	CG, RP, FD	1.1627	4.7618	0.9119	1.0003	38.7871	40
		(0.0164)	(0.2062)	(0.0031)	(0.0021)	(0.5248)	
		(1.0601, 2.4053)	(1.1248, ∞)	(0.8931, ∞)	(0.9913, 1.0313)		
X2-FL-5	SR	0.2176	1.8742	0.2932	0.9977	10.8461	10
		(0.3826)	(5.6436)	(0.0558)	(0.0008)	(0.3696)	
		(0.0722, 0.4498)	(1.0536, ∞)	(-∞, 0.8738)	(0.9973, 0.9980)		
X2-FL-6	CG, RP, FD, SR	4.7292	5.5344	0.7820	1.0245	56.0277	45
		(0.0062)	(0.0068)	(0.0027)	(0.0058)	(0.1255)	
		(4.4861, 5.0201)	(4.4861, ∞)	(0.7874, 0.8286)	(1.0126, 1.0378)		

Panel B: Instrument sets consist of a constant and the second lag of the included variables

specification	instrument set includes of the second lag of the following variables	GMM estimates, standard errors and 95% concentrated S-sets				J-statistic	number of orthogonality restrictions
		$\hat{\gamma}_{GMM}$	$\hat{\alpha}_{GMM}$	$\hat{\rho}_{GMM}$	$\hat{\beta}_{GMM}$		
X2-SL-1	CG	0.4473	3.3999	0.8597	1.0231	13.8989	10
		(0.2893)	(0.5386)	(0.0214)	(0.0558)	(0.1777)	
		(0.3990, 0.5014)	(1.3917, ∞)	(0.8525, 0.9455)	(1.0136, 1.0379)		
X2-SL-2	CG, RP	0.3784	2.0885	0.9445	0.9989	29.3849	25
		(0.0099)	(0.0164)	(0.0019)	(0.0021)	(0.2481)	
		(0.3082, 0.5489)	(1.1183, ∞)	(0.9158, ∞)	(0.9874, 1.1049)		
X2-SL-3	CG, FD	0.2980	1.5358	0.9484	1.0024	29.8648	25
		(0.0041)	(0.0032)	(0.0040)	(0.0093)	(0.2294)	
		(0.5944, 0.7130)	(1.1320, ∞)	(0.9507, 0.9620)	(0.9942, 1.0217)		
X2-SL-4	CG, RP, FD	0.4841	1.9397	0.9492	0.9993	48.5920	40
		(0.0047)	(0.0170)	(0.0034)	(0.0065)	(0.1653)	
		(0.2338, 0.6153)	(1.1869, ∞)	(0.9286, ∞)	(0.9791, 1.1483)		
X2-SL-5	SR	0.1295	1.5408	0.9613	1.0198	10.3252	10
		(0.0064)	(0.0312)	(0.0014)	(0.0032)	(0.4124)	
		(0.0091, 0.2601)	(1.0177, ∞)	(0.9378, ∞)	(1.0060, 1.0640)		
X2-SL-6	CG, RP, FD, SR	0.0199	2.3851	0.9850	0.9984	50.8719	45
		(0.0488)	(0.0567)	(0.0033)	(0.0371)	(0.2535)	
		(0.0113, 0.0286)	(1.2069, ∞)	(0.9808, ∞)	(0.9867, 1.0305)		

Table IV
GMM estimates of the preference parameters
with the marginal utility of reserves set to zero

The twenty specifications reported in Tables II and III are re-estimated with the marginal utility of reserves set to zero. That is, preference parameters are estimated using continuous-updating GMM with equations (14), (15) and (16) as moment conditions. Instruments used in each specification are the same as those used in the respective specification reported in Tables II and III.

Panel A: X1 specifications

	GMM estimates, standard errors and 95% concentrated S-sets			
specification	$\hat{\gamma}_{GMM}$	$\hat{\rho}_{GMM}$	$\hat{\beta}_{GMM}$	J-statistic
X1-FL-1	37.5629	0.7746	0.6831	2.9872
	(11.4614)	(0.0167)	(0.1283)	(0.9352)
	(15.5412, ∞)	(0.6609, ∞)	(0.5678, 0.9920)	
X1-FL-2	10.8825	0.5730	1.4508	16.0866
	(4.9421)	(0.0539)	(0.1301)	(0.7112)
	(8.5865, 38.1163)	(0.5531, 0.6059)	(1.0070, 1.5818)	
X1-FL-3	18.1799	0.6978	0.9615	11.9919
	(5.9117)	(0.0352)	(0.0987)	(0.9164)
	(13.5471, ∞)	(0.6689, 0.9531)	(0.8817, 1.2307)	
X1-FL-4	92.5203	0.6231	0.8375	20.7560
	(6.4704)	(0.0167)	(0.0099)	(0.9369)
	(30.5042, ∞)	(0.4438, ∞)	(0.4684, 1.4891)	
X1-SL-1	61.6556	0.5292	1.1791	4.1101
	(26.5496)	(0.0295)	(0.2238)	(0.8471)
	(10.7480, ∞)	(0.2344, 0.7085)	(0.6485, 1.4298)	
X1-SL-2	11.4379	0.7494	1.0470	14.6113
	(8.1137)	(0.0498)	(0.1591)	(0.7982)
	(1.4292, ∞)	(0.3733, 0.8539)	(0.8689, 1.1968)	
X1-SL-3	25.6797	0.6336	1.0241	17.3921
	(8.4231)	(0.0390)	(0.1233)	(0.6274)
	(14.0417, ∞)	(0.5425, 0.8383)	(0.8564, 1.3606)	
X1-SL-4	70.5302	0.3300	1.6189	24.9451
	(7.8428)	(0.0317)	(0.0850)	(0.8082)
	(58.8029, ∞)	(0.2975, 0.9094)	(1.1078, 1.8574)	

Panel B: X2 specifications

specification	GMM estimates, standard errors and 95% concentrated S-sets $\hat{\gamma}_{GMM}$	$\hat{\rho}_{GMM}$	$\hat{\beta}_{GMM}$	J-statistic
X2-FL-1	13.3628	0.6841	1.1549	10.0164
	(15.1805)	(0.0263)	(0.3161)	(0.4391)
	(10.0908, 16.3311)	(0.6717, 0.8435)	(1.1189, 1.2098)	
X2-FL-2	135.3660	0.4486	0.8839	24.7796
	(3.1852)	(1.2125)	(0.0367)	(0.4748)
	(114.7162, ∞)	(0.4262, ∞)	(0.7610, 1.8247)	
X2-FL-3	16.4893	0.7145	0.9564	22.7029
	(3.0949)	(0.0194)	(0.0602)	(0.5949)
	(14.1778, 19.5669)	(0.6955, 0.8387)	(0.9130, 1.0162)	
X2-FL-4	3.2991	0.8920	0.9076	38.2074
	(2.7750)	(0.0846)	(0.0109)	(0.5512)
	(2.0647, 8.6288)	(0.8619, 0.9190)	(0.8838, 0.9787)	
X2-FL-5	100.2614	0.4948	0.9613	11.1191
	(30.2864)	(0.0181)	(0.2450)	(0.3483)
	(101.6800, 137.9680)	(0.0401, 0.7520)	(0.5766, 1.5117)	
X2-FL-6	9.2985	0.8272	0.8523	46.4781
	(0.0802)	(0.0049)	(0.0002)	(0.4113)
	(5.8336, 104.0991)	(0.3262, 0.8565)	(0.8168, 1.1386)	
X2-SL-1	21.0711	0.6545	0.9864	2.3456
	(0.9283)	(0.0125)	(0.0074)	(0.9929)
	(19.9589, 30.8458)	(0.6332, 0.6798)	(0.8939, 1.0709)	
X2-SL-2	5.4532	0.6871	1.2767	26.8770
	(6.3438)	(0.0446)	(0.1973)	(0.3621)
	(4.1601, 124.1895)	(0.6544, 0.8451)	(1.2075, 1.6002)	
X2-SL-3	24.8451	0.6382	1.0304	27.0797
	(5.7846)	(0.0343)	(0.0796)	(0.3519)
	(14.7014, 142.8230)	(0.5545, 0.8099)	(0.8931, 1.1958)	
X2-SL-4	0.8155	0.9290	1.0108	48.7067
	(4.1427)	(0.3228)	(0.0474)	(0.1626)
	(0.6058, 1.0460)	(0.8976, ∞)	(0.9549, 1.0245)	
X2-SL-5	52.4439	0.6651	0.5999	10.7180
	(22.5589)	(0.0282)	(0.2215)	(0.3799)
	(43.7877, 108.5778)	(0.5412, 0.7454)	(0.4994, 0.8368)	
X2-SL-6	192.4886	0.1330	1.9835	43.3634
	(0.2427)	(0.0126)	(0.0302)	(0.5414)
	(164.5044, ∞)	(-∞, 0.2327)	(1.4421, 2.6271)	

Table V
Simulation results

The subjective discount factor β is fixed at 0.99. Simulation results are reported for various combinations of the other three parameters: the representative household's curvature parameter γ, the relative risk aversion coefficient of the foreign bond-holder α, and the habit parameter ρ. The simulated moments are calculated from 2000 replications of the sample lengths: 246 for the Deutsche mark and 324 for the pound and the yen. In parentheses are the standard deviations of the simulated moments calculated from the 2000 replications. The moment 'auto' denotes the coefficient of first-order autocorrelation. Asterisks indicate that the sample moments are within two standard deviations of the corresponding simulated moments.

Panel A: Japanese yen

	parameters	parameter values				
	γ	1.00	1.00	1.00	0.50	0.50
	ρ	0.00	0.65	0.90	0.90	0.95
	α	1.20	1.50	1.50	1.00	1.50
moments	sample moments	simulated moments				
mean($f_t - s_{t+1}$)	0.0978	0.0450*	-0.0622*	-0.2406*	0.2341*	-0.3005*
		(0.1965)	(0.1968)	(0.2016)	(0.2005)	(0.2073)
var($f_t - s_{t+1}$)	12.7553	12.1203*	12.1127*	12.2333*	12.1414*	12.2138*
		(0.3386)	(0.3270)	(0.3377)	(0.3120)	(0.3103)
cov($f_t - s_{t+1}, f_t - s_t$)	0.2069	0.0010*	-0.0035*	0.0571*	-0.0212*	-0.0065
		(0.1273)	(0.1220)	(0.1271)	(0.1162)	(0.1058)
auto($f_t - s_{t+1}$)	0.0572	-0.0022*	-0.0021*	-0.0061*	-0.0021*	0.0084*
		0.0571	(0.0564)	(0.0564)	(0.0586)	(0.0550)

Panel B: British pound

	parameters	parameter values				
	γ	1.00	1.00	1.00	0.50	0.50
	ρ	0.00	0.85	0.90	0.94	0.95
	α	1.20	1.50	1.50	1.77	1.10
moments	sample moments	simulated moments				
mean($f_t - s_{t+1}$)	-0.1819	0.1189*	-0.0696*	-0.1389*	-0.1956*	-0.0757*
		(0.1770)	(0.1776)	(0.1782)	(0.1805)	(0.1743)
var($f_t - s_{t+1}$)	10.0980	9.6114	9.6231	9.6922	9.7648*	9.7226
		(0.1572)	(0.1515)	(0.1728)	(0.2028)	(0.1725)
cov($f_t - s_{t+1}, f_t - s_t$)	0.1915	0.0005	-0.0014	0.0547*	0.1294*	0.0660*
		(0.0730)	(0.0691)	(0.0803)	(0.0957)	(0.0801)
auto($f_t - s_{t+1}$)	0.0744	-0.0030*	0.0003*	0.0190*	0.0319*	-0.0050*
		(0.0578)	(0.0558)	(0.0560)	(0.0559)	(0.0570)

Panel C: Deutsche mark

moments	parameters	parameter values				
	γ	1.00	1.00	0.50	0.50	0.50
	ρ	0.00	0.90	0.90	0.95	0.96
	α	1.20	1.50	1.50	1.50	1.52
moments	sample moments	simulated moments				
mean$(f_t - s_{t+1})$	0.1056	0.1086*	-0.3251*	-0.1793*	-0.4881	0.4342*
		(0.2145)	(0.2240)	(0.2188)	(0.2278)	(0.1989)
var$(f_t - s_{t+1})$	12.5110	11.4176	11.5281	11.4529	11.8281*	12.1855*
		(0.2300)	(0.3024)	(0.2585)	(0.4026)	(0.4296)
cov$(f_t - s_{t+1}, f_t - s_t)$	0.5776	0.0011	0.1640	0.0608	0.5008*	0.7467*
		(0.1082)	(0.1477)	(0.1238)	(0.2064)	(0.2132)
auto$(f_t - s_{t+1})$	0.0251	0.0006*	0.0077*	-0.0012*	-0.0021*	-0.1129
		(0.0637)	(0.0639)	(0.0634)	(0.0640)	(0.0610)

Figure 1

Joint 95% confidence ellipses and 95% concentrated S-sets with respect to the coefficients of relative risk aversion

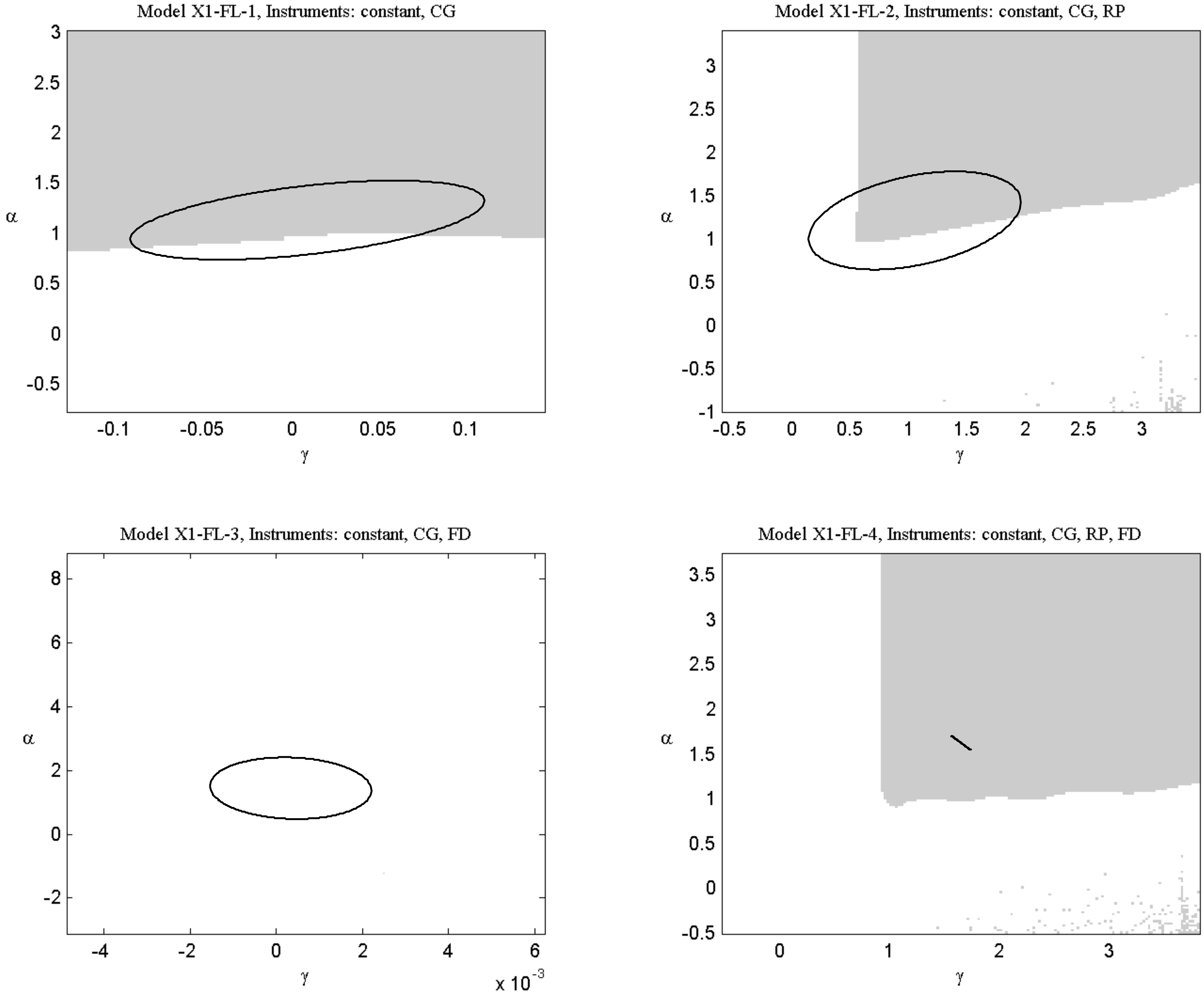

Figure 2

Joint 95% confidence ellipses and 95% concentrated S-sets with respect to the coefficients of relative risk aversion