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ON TIME VARYING RISK PREMIA IN THE FOREIGN EXCHANGE MARKET
An Econometric Analysis

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A generic intertemporal asset pricing model is applied in an international setting to generate a (possibly time varying) risk premium in the market for forward foreign exchange. The model is fitted and statistical tests of its general specification are performed. These specification tests provide weak evidence against the model.

1. Introduction

There is, by now, a substantial empirical literature which has rejected the 'simple efficiency' hypothesis in the market for forward foreign exchange, i.e., the joint hypothesis of zero risk premia and market efficiency or rational expectations. Hansen and Hodrick (1980), Cumby and Obstfeld (1981), and Hakkio (1981), among others, tested and rejected restrictions imposed upon spot and forward exchange rate data by the simple efficiency hypothesis. In other studies, such as Bilson (1981), Bilson and Hseih (1983), and Hodrick and Srivastava (1984b), profitable trading strategies in forward foreign exchange, based on simple time series prediction formulas utilizing publically available information, were formulated in an apparent violation of the simple efficiency hypothesis. A number of economists have interpreted these results as evidence of the existence of risk premia in the foreign exchange market while maintaining the hypothesis of market efficiency.[1] This paper undertakes statistical specification tests of a model which is representative of this view.[2]

The theoretical justification for this view has been demonstrated by Kouri (1977), Fama and Farber (1979), Stulz (1980, 1981, 1984), and Hodrick (1981),

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[1] Statistical evidence that a risk premium exists, is presented in Fama (1984), Hodrick and Srivastava (1984), Domowitz and Hakkio (1984), and Korajczyk (1984).

[2] This view is not universal. Those studies which examine the time series behavior of spot and forward rates are subject to the usual qualifications when applying asymptotic distribution theory to finite samples, and the apparent profits generated by the speculative strategies might be reflective of transactions costs. It has also been argued [e.g., Bilson (1981)] that the market might be inefficient and that there is no risk premium.

among others,[3] who analyze international asset pricing problems by employing methods from the theory of finance. In these papers, nominal bonds denominated in the domestic and foreign currencies contribute to the overall riskiness of the representative investor's portfolio in a fashion which depends on the price level dynamics at home and abroad. To compensate investors for bearing these portfolio risks, a risk premium in forward foreign exchange emerges.

Empirically, the issue of the risk premium has been pursued in greater detail along two main lines. One group has advanced the 'portfolio balance' approach to explain the existence of a risk premium. This approach argues that portfolio balancers proportion their wealth among domestic money, domestic bonds, and foreign currency denominated bonds according to expected relative rates of return at home and abroad. Investors are assumed to view domestic and foreign bonds of identical default and political risk as less than perfect substitutes due to their different currency denominations. This imperfect substitutability drives a wedge between expected net rates of return on domestic and foreign bonds which gives rise to the risk premium. Structural estimation based on the portfolio balance model, however, has been largely unsuccessful.[4] Closely related, is an attempt by Frankel (1982a) to estimate a portfolio balancing model in which investors optimize over the mean and variance of returns. In this model, the risk premium in foreign exchange depends on asset supplies and investor risk aversion. Frankel could not reject the hypothesis of risk neutrality, in which case the risk premium disappears. However, he is careful to point out that the power of his test is extremely low, and that some of the assumptions he employs may not be econometrically justified.

The other line of empirical research on the risk premium in forward foreign exchange relies on discrete-time, intertemporal asset pricing models such as those of Lucas (1978, 1982) and Brock (1982).[5] In these models, investors are assumed to maximize discounted expected utility defined on consumption subject to sequential budget constraints. The first-order, necessary conditions require that the conditional first cross-moment of the marginal rate of substitution of consumption between periods t and $t+k$ and the k-period real return from investing in an asset equals zero. Exploiting this equilibrium asset pricing relationship in an international environment, the risk premium on forward foreign exchange is seen to be proportional to the conditional covariance of the k-period return on foreign exchange speculation and the marginal rate of substitution of money between periods t and $t+k$. Depending on the stochas-

[3] See also Driskill and McCafferty (1982), who adopt a market segmentation approach.

[4] For empirical research along these lines, see Frankel (1982b) and Dooley and Isard (1984).

[5] These are applications of the intertemporal substitution hypothesis which has been so prominent in the business cycle literature, e.g., Lucas and Rapping (1969), Barro (1976), and Long and Plosser (1983).

tic structure of the economy, this conditional covariance (and the factor of proportionality) can depend on elements in economic agents' information sets. Thus, time variation of the risk premium can be explained through the time variation of this conditional covariance. As Frankel (1982a) has argued, a successful inquiry into the nature of the risk premium must explain how it can fluctuate from positive to negative,. since the forward premium appears to be an unconditionally unbiased predictor of exchange rate changes.

Recently, Hansen and Hodrick (1983) formulated and estimated three econometric models of the risk premium based on this conditional covariance decomposition.[6] In each case, a joint hypothesis that the model was correct and a particular statistical assumption (i.e., a constant conditional moment or cross-moment) was tested. The ultimate rejection of all three formulations led Hodrick and Srivastava (1984b) to investigate whether or not 'some other model of risk and return is necessary to describe the risk premium'.[7] However, these authors frequently mention that these rejections are not necessarily rejections of the underlying model generating the risk premium since all tests were on a joint hypothesis of a particular statistical assumption and of the underlying economic model.

This paper inquires further in the issues of forward foreign exchange risk premia motivated by the intertemporal asset pricing models considered by Hansen and Hodrick (1983). In particular, specification tests of a model which generates the time varying risk premium are directly undertaken.

The remainder of this paper is organized as follows. In the next section, a more detailed exposition of the hypothesis to be tested is developed. The estimation strategy employed here draws on the procedures outlined in Hansen and Singleton (1982) and Hansen (1982), and a brief description of this procedure is presented in section 3. The data and instruments used in estimation are discussed in section 4. Section 5 reports the empirical results, and some concluding remarks are contained in section 6.

2. The theoretical motivation

The theoretical models of intertemporal asset pricing, such as those of Lucas (1978, 1982) and Brock (1982), relied upon by Hansen and Hodrick as well as in the present paper, assume that investors maximize expected discounted utility defined on future levels of consumption subject to sequential budget constraints. In equilibrium, asset prices are set such that the marginal utility of

[6] These authors imposed additional structure on the problem so that the risk premium was seen to be proportional to returns on various financial assets. The factors of proportionality included conditional moments or cross-moments of certain economic variables.

[7] Two of these formulations were rejected by Hansen and Hodrick (1983). In a closer examination of the third model, which uses a latent variables statistical procedure, Hodrick and Srivastava (1984b) were able to reject the restrictions imposed by that formulation also.

a unit of current consumption foregone equals the expected discounted utility of the return from investing that unit of the consumption good. That is,

$$U'(C_t) = \beta \mathrm{E}\big(U'(C_{t+1}) r_{t+1}^j \mid I_t\big), \qquad j = 1, 2, \ldots, N, \tag{1}$$

where I_t is the information set available to economic agents at time t, $U'(\cdot)$ is the marginal utility of consumption, C_t, $0 < \beta < 1$ is a subjective discount factor, and r_{t+1}^j is the one-period real return on asset j, indexed in terms of the consumption good. Agents are assumed to be 'rational', so $\mathrm{E}(\cdot \mid I_t)$ is interpreted to be both the mathematical expectation conditioned on information available at t, as well as the expectation of economic agents. Since C_t is known at time t, the equilibrium condition can be rearranged as

$$1 = \beta \mathrm{E}\left[\left.\frac{U'(C_{t+1}) r_{t+1}^j}{U'(C_t)}\right| I_t\right], \qquad j = 1, 2, \ldots, N. \tag{2}$$

Assuming that the utility function is concave, it is seen from (2), that asset prices are set to yield high returns during periods in which consumption is low, relative to levels which are expected to prevail in the future, and vice versa. Empirically, this model of investor behavior has been met with varying degrees of success in the pricing of common stocks and treasury bills.[8] In the international context, the desire of investors to smooth consumption over time causes a risk premium to emerge in forward foreign exchange. To see this, let r_{t+1}^j be the real return to a domestic resident from taking an uncovered, one-period investment in a foreign currency denominated, nominal bond. Thus, eq. (2) becomes

$$1 = \beta \mathrm{E}\left[\left.\frac{U'(C_{t+1})\left(1 + i_{j,t+1}^*\right)}{U'(C_t)} \frac{P_t}{P_{t+1}} \frac{S_{j,t+1}}{S_{j,t}}\right| I_t\right], \tag{3}$$

$$j = 1, 2, \ldots, M < N,$$

where P_t is the domestic currency price of the consumption good in t, $i_{j,t+1}^*$ is the one-period, nominal return on a bond maturing at $t+1$ denominated in currency j, and $S_{j,t}$ is the domestic currency price of one unit of currency j. Now, consider taking a covered position in the investment of the foreign bond. In this case, the equilibrium pricing relation (2) becomes

$$1 = \beta \mathrm{E}\left[\left.\frac{U'(C_{t+1})\left(1 + i_{j,t+1}^*\right)}{U'(C_t)} \frac{P_t}{P_{t+1}} \frac{F_{j,t}}{S_{j,t}}\right| I_t\right], \qquad j = 1, 2, \ldots, M, \tag{4}$$

[8]See, for example, Grossman and Shiller (1981), Hansen and Singleton (1982, 1983), Dunn and Singleton (1983), and Rotemberg (1984).

where $F_{j,t}$ is the one-period forward exchange rate on currency j. By the covered interest arbitrage condition, this return is identical to the real return from investing in a one-period domestic nominal bond of similar default risk.

Taking the difference between (3) and (4), and dividing this result by the discount factor, β, and the nominal return on the foreign bond, $(1+i^*_{j,t+1})$, yields the fundamental restrictions imposed by the model[9]

$$\mathrm{E}\left[\frac{U'(C_{t+1})}{P_{t+1}}\frac{P_t}{U'(C_t)}\frac{(S_{j,t+1}-F_{j,t})}{S_{j,t}}\middle| I_t\right]=0, \qquad j=1,2,\ldots,M. \tag{5}$$

That is, the model requires the product of $(U'(C_{t+1})/P_{t+1})/(U'(C_t)/P_t)$ and $(S_{j,t+1}-F_{j,t})/S_{j,t}$ to be orthogonal to all elements in I_t. Since $(1/P)$ is the purchasing power of one unit of the home currency, $U'(C)/P$ is the utility denominated value of a unit of the home currency. Thus the first term in the expectation in (5), being the ratio of the utility values of the domestic money at $t+1$ and t, is interpreted as the intertemporal marginal rate of substitution of money. The second term would be, at $t+1$, the ex-post profit from speculating forward on currency j, $(S_{j,t+1}-F_{j,t})$, normalized by $S_{j,t}$. Hence, eq. (5) states that the conditional first cross-moment between the intertemporal marginal rate of substitution of domestic money, and the normalized profits from foreign exchange speculation, be zero.

Exploiting (5) in the decomposition of the conditional covariance, the risk premium, defined as the expected profit on forward foreign exchange speculation, can be expressed as

$$\mathrm{E}\left[\frac{S_{j,t+1}-F_{j,t}}{S_{j,t}}\middle| I_t\right]=\frac{\operatorname{cov}\left[\dfrac{U'(C_{t+1})}{P_{t+1}}\dfrac{P_t}{U'(C_t)};\dfrac{S_{j,t+1}-F_{j,t}}{S_{j,t}}\middle| I_t\right]}{-\mathrm{E}\left[\dfrac{U'(C_{t+1})}{P_{t+1}}\dfrac{P_t}{U'(C_t)}\middle| I_t\right]}, \tag{6}$$

where $\operatorname{cov}(\cdot;\cdot|I_t)$ is the conditional covariance. Since the intertemporal marginal rate of substitution of money is always expected to be positive, the risk premium is seen to be proportional to the conditional covariance of the intertemporal marginal rate of substitution of money and the profit from foreign exchange speculation. The factor of proportionality and the conditional covariance can depend on sample information and need not be time invariant, hence, the risk premium can fluctuate over time.

Hansen and Hodrick (1983) base their econometric analysis on eq. (6), and attempt to explain the behavior of the risk premium through time variation of

[9] The division is legitimate since β is a constant and $i^*_{j,t+1}$ is an element of I_t.

the factor of proportionality alone. The joint hypothesis which they test involves a distributional assumption concerning the constancy of a conditional covariance as well as the implications of the underlying economic model. As Hansen and Hodrick (1983) and Hodrick and Srivastava (1984b) note, the ultimate rejection of the variants of eq. (6) which they investigate cannot be viewed as a conclusive rejection of the underlying model due to the joint nature of the test. In other words, it is possible that violations of the time invariance assumption of the conditional covariance led to the rejections. Thus, one way to proceed might be to conduct an explicit general equilibrium analysis and specify the state valuation function in full detail. This would permit explicit modeling of the dependence of the risk premium on the state variables of the system, and the development of empirically testable hypotheses. The aims of this paper, however, are more modest and focus on the restrictions, embodied in eq. (5), of the underlying economic model generating the risk premium. In what follows, a direct test of these restrictions is undertaken, employing a methodology which does not require that the complete stochastic environment be specified.

3. Estimation

In order to proceed with estimation, economic agents' utility function must be parameterized. In this paper, the constant relative risk aversion form is used,

$$U(C) = \delta C^{1-\gamma}/(1-\gamma), \qquad \gamma \geq 0, \tag{7}$$

where γ is the parameter of relative risk aversion and δ is an arbitrary constant. Risk lovers ($\gamma < 0$) are ruled out on a-priori grounds. Using this form of the utility function in (5), the theoretical restrictions arising from optimal investor behavior are restated as

$$\mathrm{E}\left(c_{t+1}^{\gamma} p_{t+1} \Pi_{j,t+1} \mid I_t\right) = 0, \qquad j = 1,2,\ldots,M, \tag{8}$$

where for notational ease, we define $c_{t+1} = C_t/C_{t+1}$, $p_{t+1} = P_t/P_{t+1}$, and $\Pi_{j,t+1} = (S_{j,t+1} - F_{j,t})/S_{j,t}$. In words, the product of the intertemporal marginal rate of substitution of money between time t and $t+1$ and nominal speculative profits from trading currency j one period forward at time t is unforecastable, based on the information, I_t, which economic agents have an incentive to acquire and which is available to them at t. That $c_{t+1}^{\gamma} p_{t+1} \Pi_{j,t+1}$ ($j = 1,2,\ldots,M$) is predicted by the theory to be uncorrelated with any and all elements in I_t, forms an empirically testable hypothesis which is examined in detail.

The model is estimated using the distribution free, generalized instrumental variables procedure outlined in Hansen and Singleton (1982). The statistical motivation and large sample properties upon which inference is based is developed in Hansen (1982). I assume throughout that all of his regularity conditions hold. The methodology, as it pertains to the current problem, is outlined briefly below.

Assume that the data are jointly stationary and ergodic, and let $X'_{t+1} = (c_{t+1}, p_{t+1}, \Pi_{1,t+1}, \ldots, \Pi_{m,t+1})$, $m \leq N$. Define the $(m \times 1)$ vector function h by

$$h(X_{t+1}, \gamma)' = (c^{\gamma}_{t+1} p_{t+1} \Pi_{1,t+1}, \ldots, c^{\gamma}_{t+1} p_{t+1} \Pi_{m,t+1}). \tag{9}$$

Let Z_t be an observable, q-dimensional vector of variables in agents' information set at t and define the $(mq \times 1)$ vector functions f and g by

$$f(X_{t+1}, Z_t, \gamma) = h(X_{t+1}, \gamma) \otimes Z_t,$$

$$g(\gamma) = \mathrm{E}[f(X_{t+1}, Z_t, \gamma)],$$

where $\otimes$ is the Kronecker product and $\mathrm{E}(\cdot)$ is the unconditional expectations operator.

Denote the 'true' value of the risk aversion parameter by γ_0. Theoretically, $c^{\gamma_0}_{t+1} p_{t+1} \Pi_{j,t+1}$ is uncorrelated with elements in I_t. Since Z_t is a subset of I_t, it can be seen that $g(\gamma) = 0$ when evaluated at $\gamma = \gamma_0$. This generates a family of mq orthogonality conditions which are to be used in estimation.

Let T be the sample size. The method of moments estimator of $g(\gamma)$ is the mq-dimensional vector

$$g_T(\gamma) = \frac{1}{T} \sum_{t=1}^{T} f(X_{t+1}, Z_t, \gamma).$$

If the model is correctly specified, $g_T(\gamma)$ should be close to zero when evaluated at $\gamma = \gamma_0$. There is only one parameter to estimate, however, and there are mq orthogonality conditions, so that estimation of the model cannot proceed by minimizing $g_T(\gamma)$ when $mq > 1$. Thus Hansen (1982) proposes that γ be chosen to minimize the 'distance' function

$$J_T(\gamma) = g_T(\gamma)' W_T g_T(\gamma), \tag{10}$$

where W_T is an $(mq \times mq)$ symmetric weighting matrix. The value of γ that minimizes the distance function (10) is the generalized method of moments estimator of γ_0 in that it makes $g_T(\gamma)$ close to zero according to the metric defined by W_T. For an appropriate choice of W_T, Hansen (1982) shows that the minimizer of (10) is consistent with the following large sample properties.

Let

$$W_0 = \left[\mathrm{E} f(X_{t+1}, Z_t, \gamma_0) f(X_{t+1}, Z_t, \gamma_0)'\right]^{-1},$$

$$d_0 = \mathrm{E}\left[\frac{\partial f(X_{t+1}, Z_t, \gamma_0)}{\partial \gamma}\right],$$

and let γ_T be the minimizer of (10). Then $\sqrt{T}(\gamma_T - \gamma_0)$ is distributed as $\mathrm{N}(0, (d_0' W_0 d_0)^{-1})$ in large samples. In order to conduct inference, a two-step estimation procedure is indicated. Let

$$W_T(\gamma) = \left[\frac{1}{T}\sum_{t=1}^{T} f(X_{t+1}, Z_t, \gamma) f(X_{t+1}, Z_t, \gamma)'\right]^{-1},$$

$$d_T(\gamma) = \frac{1}{T}\sum_{t=1}^{T} \frac{\partial f(X_{t+1}, Z_t, \gamma)}{\partial \gamma}.$$

First, an initial guess value, γ^1, is chosen to form $W_T(\gamma^1)$. Minimization of (10) using this weighting matrix yields a first-step estimator, γ_T^1. This estimator is consistent, but it does not have the desired asymptotic properties. To obtain a consistent estimator of the asymptotic variance, W_T must be constructed using a consistent estimator of γ_0. Minimizing (10) using $W_T(\gamma_T^1)$ yields a second-step estimator, γ_T^2, and $[d_T(\gamma_T^2)' W_T(\gamma_T^2) d_T(\gamma_T^2)]^{-1}$ forms a consistent estimator of $(d_0' W_0 d_0)^{-1}$.

Economic theory provides little guidance as to what value γ should take. However, the empirically testable hypothesis of principle interest here concerns the overall specification of the model generating the risk premium. The first-order condition from the minimization of (10) is

$$\left[\frac{\partial g_T(\gamma)}{\partial \gamma}\right]' W_T(\gamma) g_T(\gamma) = 0. \tag{11}$$

The model is exactly identified when $mq = 1$. To estimate γ_0, a linear combination of the mq orthogonality conditions is set to zero. This leaves $(mq - 1)$ unconstrained linearly independent orthogonality conditions which should be close to zero, provided that the model is correctly specified. Hansen (1982) has shown that T times the second-step minimized value of (10) has a chi-square distribution with $(mq - 1)$ degrees of freedom in large samples and can be used to test the overidentifying restrictions imposed by the model.

To put this specification test in perspective, if γ_0 were known, an analogous procedure would be to regress $(c_{t+1}^{\gamma_0} p_{t+1} \Pi_{j,t+1})$ on sample information available to agents at t. The null hypothesis is this case would be that all regression coefficients be insignificantly different from zero. What is done here is similar in spirit, except that γ_0 is replaced with a consistent estimator, γ_T^2.

4. Data and instruments

The sampling interval of the data is monthly. The exchange rates used are the Canadian dollar, German mark (DM), Netherlands guilder, and U.K. pound, all relative to the U.S. dollar. All exchange rate data are Friday closing quotations from the Harris Bank *Foreign Exchange Weekly Review* and are sampled as follows. To construct profits from forward foreign exchange speculation, spot rates are sampled on Fridays which fall nearest to the end of the calendar month, and forward rates (and the normalizing spot rates[10]) are sampled four Fridays prior to the Friday falling nearest to the end of the month. To construct the forward premium, both spot and forward rates are sampled four Fridays prior to the Friday falling nearest to the end of the month.[11] The seasonally adjusted U.S. consumption data is from the Commerce Department's *National Income and Products Accounts*. Aggregate consumption is deflated by monthly population estimates, obtained from the Census Bureau's *Current Population Reports*, to arrive at per capita terms. Commodity prices are measured by the implicit price deflators for consumption. The estimation period extends from 3/73 to 7/83.[12]

Two measures of consumption were used. Expenditures on non-durables plus services, and expenditures on non-durables only (denoted by NDS and ND, respectively). Use of the latter measure is motivated by the fact that expenditures on items which would appear to display a very small degree of intertemporal substitution in the short run, such as spending on housing and household operation, are recorded in the service account.

The model is estimated jointly across the four currencies. Accordingly, this specifies $m = 4$, and $h(X_{t+1}, \gamma)$ in eq. (9) is a four-dimensional vector.

[10] This is the spot exchange rate appearing in the denominator of the expectation in eq. (5).

[11] This was done to match the observations as close as possible, given this data set; however, some misalignment remains. First, there is a need to match the exchange rate data to calendar time since the consumption data is reported monthly. Fridays occurring nearest to the end of the month, however, are separated by either four or five weeks, so the sampling interval is not constant. Second, since the Harris data is comprised of weekly observations, a slight mismatch of spot and forward rates is present when calculating forward and risk premia, e.g., a one-month forward contract sold on Friday, November 4 is due on Monday, December 4. This second source of mismatching, while recognized in the literature, does not appear to be very serious [e.g., see Hodrick and Srivastava (1984b)].

[12] Hansen and Hodrick (1983) argue that there may have been initial uncertainty with regard to the operation of flexible exchange rates in that agents may have been expecting a return to some sort of fixed exchange rate arrangement. They further argue that this source of uncertainty was eliminated when a formal amendment to the IMF articles of agreement in which the flexible exchange rate system was formally ratified in January 1976. Estimation over the shorter period, 2/76–7/83, yielded very similar results and are not reported. These results are available from the author upon request.

As is usually the case with instrumental variables estimation, little guidance is available for actual selection of the instruments. Consequently, elements forming the instrument vector, Z_t, are restricted to those variables which might reasonably be of importance to agents in solving their forecasting problem. Since agents are concerned about predicting the future return in terms of the consumption good from investing in assets denominated in a foreign currency, the current consumption ratio, c_t, is likely to provide important information. Moreover, current and past movements in the exchange rates of the other currencies as well as movements in the exchange rate of the individual currency in question are likely to provide important information due to the contemporaneous correlation of exchange rate changes across currencies. In particular, two specifications of the instrument vector are considered. The first employs the current consumption ratio along with current and past values of realized profits from foreign exchange speculation in each of the four currencies in the sample (lags 0 through 2 in the instrument vector). For example, with k lags, the instrument vector is written as

$$Z_t' = (c_t, \Pi_{1,t}, \ldots, \Pi_{4,t}, \ldots, \Pi_{1,t-k}, \ldots, \Pi_{4,t-k}).$$

This choice is motivated by the idea that future speculative profits might be partly predictable from past profits. The second formulation replaces these realized speculative profits with current and past forward premia on these currencies. This choice is motivated by the notion that the forward premium can be expressed as the sum of the expected rate of change of the exchange rate and the risk premium and, therefore, contains information about expected speculative profits.[13]

Since the estimation and inference procedures assume that the data is covariance stationary, the first twelve sample autocorrelations of the consumption ratio, the commodity price ratio, (normalized) speculative profits and forward premia are reported in table 1. As is seen, the sample autocorrelations for each series drops off quickly so that the data do not appear to violate the stationarity assumption. Since subtracting one from the consumption and price ratios approximates the difference in log-levels of consumption and prices, these ratios display features of stationary time series even though their levels may be non-stationary.[14] The ex-post speculative profits appear to display little if any serial correlation, while the forward premia appear to follow low-ordered autoregressive processes.

[13]See Fama (1984) and Hodrick and Srivastava (1984a) who exploit this idea and present evidence that most of the variation in the forward premium is due to variation in the risk premium.

[14]See Nelson and Plosser (1982) who argue that differencing is the appropriate way to achieve stationarity of the data.

Table 1

Sample autocorrelations and standard deviations of the data March 1973 to July 1983.

Series	Consumption	Exchange rate	Autocorrelations $\hat{\rho}_1$	$\hat{\rho}_2$	$\hat{\rho}_3$	$\hat{\rho}_4$	$\hat{\rho}_5$	$\hat{\rho}_6$	$\hat{\rho}_7$	$\hat{\rho}_8$	$\hat{\rho}_9$	$\hat{\rho}_{10}$	$\hat{\rho}_{11}$	$\hat{\rho}_{12}$	Standard deviation
c_{t+1}[a]	NDS[e]		−0.25	0.12	0.05	−0.05	0.03	0.01	−0.10	0.11	−0.01	0.14	0.07	−0.14	0.005
	ND[f]		−0.30	0.09	0.07	−0.03	−0.03	−0.03	−0.02	0.00	−0.03	0.22	−0.05	−0.07	0.008
p_{t+1}[b]	NDS		0.19	0.19	0.18	0.28	0.18	0.31	0.13	0.08	0.17	0.21	0.16	0.12	0.003
	ND		0.33	0.33	0.29	0.29	0.18	0.31	0.22	0.10	0.32	0.33	0.14	0.12	0.005
Π_{t+1}[c]		\$/Can.\$	−0.01	−0.11	0.11	−0.01	0.08	−0.11	−0.11	0.10	−0.04	−0.01	0.13	−0.25	0.012
		\$/DM	0.16	−0.08	0.07	−0.07	−0.17	−0.07	0.12	−0.04	0.13	0.18	−0.05	0.00	0.030
		\$/guilder	0.14	0.03	0.08	0.01	−0.16	−0.07	0.06	−0.03	0.04	0.25	0.05	−0.02	0.029
		\$/pound	0.13	0.05	−0.05	−0.13	0.17	0.11	−0.04	0.09	−0.16	0.06	0.27	−0.02	0.026
$(F_t - S_t)/S_t$[d]		\$/Can.\$	0.49	0.22	0.14	0.01	0.02	−0.01	0.00	0.02	0.14	−0.01	0.05	0.00	0.002
		\$/DM	0.74	0.55	0.39	0.27	0.26	0.26	0.30	0.35	0.39	0.35	0.32	0.29	0.002
		\$/guilder	0.56	0.48	0.32	0.21	0.22	0.20	0.22	0.25	0.26	0.21	0.17	0.15	0.003
		\$/pound	0.77	0.66	0.55	0.54	0.44	0.33	0.30	0.25	0.27	0.26	0.26	0.23	0.003

[a] c_{t+1} is the ratio of per capita consumption in period t to period $t+1$.
[b] p_{t+1} is the ratio of period t to period $t+1$'s implicit price deflator for consumption expenditures.
[c] Π_{t+1} is the spot rate observed on the Friday nearest to the end of the month minus the forward rate observed four Fridays earlier, normalized by the spot rate observed at the time the forward rate is sampled.
[d] $(F_t - S_t)/S_t$ is the forward rate minus the spot rate (normalized by the spot rate) all observed four Fridays prior to the Friday nearest to the end of the month.
[e] NDS is the expenditures on non-durables plus services.
[f] ND is expenditures on non-durables only.

Table 2

Estimation results when the consumption ratio and speculative profits are used as instruments.

Consumption	Lag[a]	γ_T	s.e.(γ_T)[b]	$T\min(J_T)$[c]	d.f.[d]	Confidence[e]
NDS[f]	0	1.42	42.43	19.247	19	0.559
	1	39.98	29.08	41.270	35	0.785
	2	50.38	23.82	59.777	51	0.813
ND[g]	0	0.00	24.93	19.006	19	0.544
	1	2.82	19.40	41.196	35	0.782
	2	7.25	15.65	60.202	51	0.823

[a]Lag indicates the lag length of the instrument vector.
[b]s.e.(γ_T) is the standard error of γ_T.
[c]$T\min(J_T)$ is the minimized value of the distance function (10) at the second step, multiplied by the sample size, and is distributed as χ^2(d.f.) under the null hypothesis.
[d]d.f. is the degrees of freedom.
[e]Confidence $= \int_0^{T\min(J_T)} \chi^2_{(\mathrm{d.f.})}(t)\,\mathrm{d}t$.
[f]NDS = non-durables plus services.
[g]ND = non-durables only.

5. Empirical results

The results from estimating the model using the current consumption ratio and current and past speculative profits as instruments appear in table 2.[15] As can be seen, point estimates of γ are generally quite large and vary substantially with the lags of the instrument vector and with the measure of consumption. For example, when consumption is measured by non-durables plus services (NDS), with the current values plus two lags of speculative profits (lag = 2) included as instruments, γ is estimated to be 50.38. At lag = 0 and consumption is measured by non-durable spending (ND), the algorithm corners to 0.0.

Overall, what is observed from table 2 is that lower estimates of γ are obtained when consumption is measured by spending on non-durables only. Furthermore, both the estimates of γ and their corresponding standard errors are large relative to those obtained by others [such as Hansen and Singleton (1982, 1983), Dunn and Singleton (1983), and Rotemberg (1984)] who have fit models of this type using aggregate stock market data. Statistically, however, these differences are not significant since most of the point estimates in table 2 lie within a standard error of zero. One reason for these qualitative differences may be that the stock market studies employ returns on a *portfolio* of assets; whereas, here, in order to highlight how risk premiums in the foreign exchange market can be generated, only individual asset returns are used.

[15]A number of starting values were experimented with without substantial differences in the final estimates. For table 2, a starting value of $\gamma_0 = 1$ was used. The first step estimates γ_T^1 for lags 0–3 are 1.43, 26.98, 31.61 with consumption = NDS and 0.0, 2.17, 4.85 with consumption = ND.

Table 3

Estimation results when the consumption ratio and forward premia are used as instruments.[a]

Consumption	Lag	γ_T	s.e.(γ_T)	$T\min(J_T)$	d.f.	Confidence
NDS	0	43.51	33.13	32.301	19	0.972[b]
	1	44.95	26.84	43.194	35	0.839
	2	41.45	23.65	56.450	51	0.721
ND	0	17.51	26.25	33.227	19	0.977[b]
	1	13.79	18.97	44.377	35	0.867
	2	12.67	16.64	57.079	51	0.741

[a] See notes to table 2.
[b] Significant at the 5% level.

While the distance function (10) is flat in the region of its minimum, its value is not very large. Tests of the overidentifying restrictions cannot reject the model at standard significance levels. In particular, the specification which provides the most evidence against the model is that in which consumption is measured by non-durables spending and lag = 2 for the instrument vector. The value of the test statistic is 60.202 and has a marginal significance level of 0.177 under the null hypothesis.

Table 3 reports the estimation results from employing current and lagged forward premia in place of ex-post speculative profits as instruments.[16] Essentially, the same estimates are obtained from employing the second set of instruments as with the first, although they are somewhat larger here. Estimates of γ continue to be large with correspondingly large standard errors. Again, the estimates vary substantially across consumption measures and, to a lesser extent, over the lag length. The major difference to arise with the change in instruments is that tests of the overidentifying restrictions now yield some evidence against the model. In particular, the model can be rejected at the 5% level at lag = 0 when either consumption measure is used (marginal significance levels are 0.028 and 0.023 for consumption measured by NDS and ND, respectively). However, additional lags of the instrument vector appear to be orthogonal to $c_{t+1}^{\gamma} p_{t+1} \Pi_{j,t+1}$, $j = 1, \ldots, 4$.

It can be seen from the standard errors reported in table 1 the ex-post speculative profits display much (roughly five times) more variability than either the consumption or price ratios. Thus, the time series properties of $c_{t+1}^{\gamma} p_{t+1} \Pi_{j,t+1}$ are dominated by $\Pi_{j,t+1}$. It can be shown that $\Pi_{j,t+1}$ can be decomposed into the risk premium and the unanticipated change in the spot rate. Since it is likely that movements in this serially uncorrelated forecast error swamp the movements in the risk premium, the time series properties of

[16] A starting value $\gamma_0 = 1$ was used. First step estimates γ_T^1 for lags 0–3 are 32.37, 29.81, 25.91 with consumption = NDS and 14.20, 9.50, 8.43 with consumption = ND.

$\Pi_{j,t+1}$ are in turn dominated by this forecast error. This explains why past values of speculative profits tend to be of limited use in predicting $c_{t+1}^{\gamma} p_{t+1} \Pi_{j,t+1}$, and, hence, the lack of evidence provided against the model when these instruments are used.

That the forward premiums appear to be correlated with $c_{t+1}^{\gamma} p_{t+1} \Pi_{j,t+1}$ is consistent with the notion that the forward premium contains the risk premium and the expected rate of change of the spot rate. Since movements in the forward premium appear to be dominated by movements in the risk premium [see Fama (1984) and Hodrick and Srivastava (1984a) who provide evidence on this point], this explains why $\Pi_{j,t+1}$ and, hence, $c_{t+1}^{\gamma} p_{t+1} \Pi_{j,t+1}$ is partly forecastable based on the current forward premium.

6. Conclusions

In this paper, a generic, one-good, representative-agent model of optimal investor behavior was applied in an international setting to explain time varying risk premia in forward foreign exchange. The specification of the model placed restrictions on the correlations among various time series variables. Statistical tests of these restrictions yielded rejections at the 5% level in some cases, but the evidence against the model is not overwhelming.

Several comments are in order. First, part of the maintained hypothesis involves the particular parameterization of agents' utility function which may be restrictive. For instance, identification of the model requires the assumption that the utility function is *not* subject to random shifts. As argued by Garber and King (1984), violations of this assumption typically will lead to 'mongrel' parameter estimates. A related matter concerns the specific form of the utility function and the measures of the consumption flow. Possible generalizations might proceed along the lines of Dunn and Singleton (1984) who relax the assumption that preferences are time separable and include the implicit service flow from durable goods purchases in measuring consumption. Additionally, it might be useful to incorporate other important items, such as leisure, which do not appear in published consumption data. Secondly, the risk premium was generated by considering the pricing of individual assets denominated in foreign currencies. An alternative strategy might look at a portfolio of currencies; however, the problem here is to avoid averaging out the systematic risk. Finally, inference conducted in this paper is based on the available large sample theory and should be interpreted with caution. Indeed, estimation of the model yielded very large estimates of the coefficient of risk aversion and correspondingly large standard errors which indicate a potential small sample problem. The extent to which any small sample bias is present and the nature of the bias were not explored. This, along with the other extensions mentioned above, are left for future research.

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