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Source: *The Journal of Finance*, Vol. 44, No. 1 (Mar., 1989), pp. 183-194
Published by: Blackwell Publishing for the American Finance Association
Stable URL: http://www.jstor.org/stable/2328282
Accessed: 26/11/2009 07:24

THE JOURNAL OF FINANCE • VOL. XLIV, NO. 1 • MARCH 1989

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ABSTRACT

An asset-pricing model with an unobservable time-varying risk premium is used to price forward foreign exchange contracts. Specifically, the term spectrum of forward foreign exchange contracts is examined in order to focus on country-specific and maturity-specific information. The testable restrictions imposed by the model are consistent with both cross-country and cross-maturity forward contracts except at the short end of the maturity spectrum for cross-country forward exchange rates. This indicates that the intertemporal model is relatively robust in valuing forward contracts of different maturities and for different exchange rates but that it may fail when there are significant short-term country-specific shocks.

THEORETICAL AND EMPIRICAL studies have demonstrated that the forward exchange rate incorporates a time-varying risk premium in addition to an expected future spot exchange rate.[1] Recent works by Fama [9] and Hodrick and Srivastava [21] reveal the importance of the risk premium by showing that the variability in risk premiums exceeds the variability in expected change in spot exchange rates. These results are very general in that they are based on the definition that equates the forward exchange rate to the expected future spot exchange rate and a risk premium term. A logical but more demanding next step is to model and conduct specification tests of the risk premium itself.

The purpose of this paper is to use an intertemporal latent-variable asset-pricing model (ILVM) to examine the risk-return characteristics of the term structure of forward foreign exchange contracts. More specifically, the ILVM is used to price various bilateral dollar forward exchange contracts for each maturity as well as the different maturities for each bilateral forward exchange contract.

The ILVM was first utilized by Hansen and Hodrick [16] and Hodrick and Srivastava [20] to examine forward exchange risk premiums. The two studies are, however, confined to just one-month forward exchange contracts. In addition, the studies yield conflicting results. The application of the ILVM to other securities also yields mixed results. Gibbons and Ferson [12] fail to reject a single latent-variable model for U.S. common stocks. Campbell and Clarida [4] also

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[1] Theoretically, the presence of a forward exchange risk premium is established by a large number of international equilibrium asset-pricing models. These models include those of Grauer, Litzenberger, and Stehle [14], Fama and Farber [10], Frankel [11], Hodrick [19], and Stulz [28]. For empirical studies, see, for example, the works of Bilson [1], Hsieh [22], Hansen and Hodrick [16], and Hodrick and Srivastava [20].

find the behavior of the Eurocurrency interest rates to be consistent with the ILVM. On the other hand, Campbell [3] rejects the ILVM in a study of various U.S. securities. Since the assets examined in the extant literature generally lack common attributes, it is difficult to discern the reasons for the contradictory results. By examining the co-movements of the risk premiums in various bilateral dollar forward exchange rates and in the maturity structure of forward exchange contracts, the tests in this paper analyze the roles of currency-specific and maturity-specific information in the ILVM. Therefore, the evidence presented here provides pertinent information in the assessment of the ILVM in general.

I. Model

The equilibrium relationship used in the ILVM is induced by the optimal consumption-investment decisions of a representative investor who maximizes intertemporal expected utility functions defined over lifetime consumption expenditures subject to income and wealth constraints. The model, therefore, has the same roots as the class of equilibrium asset-pricing models constructed by Rubinstein [27], Lucas [23], Breeden [2], and Cox, Ingersoll, and Ross [5], among others.

The equilibrium condition equates the marginal utility of consumption today to today's expectation of discounted future returns weighted by the marginal utility of future consumption. It can be used to price the forward exchange risk premium by considering the covered and uncovered dollar returns from investing in a nominally riskless bond denominated in foreign currency. The covered return is priced so that

$$E_t(m_{t+k} r^j_{t+k} F^j_{t,k}) = S^j_t, \tag{1}$$

where expectations are assumed to be rational so that $E_t(\cdot) = E(\cdot \mid I_t)$ is the mathematical expectation conditional on I_t, the information available at date t; m_{t+k} is the discounted marginal utility ratio of consumption at time $t + k$ to time t; r^j_{t+k} is one plus the nominal rate of return on a nominally riskless bond denominated in foreign currency j from time t to $t + k$; $F^j_{t,k}$ is the forward exchange rate of U.S. dollars per unit of foreign currency j observed at time t to be exercised at time $t + k$ and S^j_t is the spot exchange rate of U.S. dollars per unit of foreign currency j observed at time t.

The uncovered return from investing in a foreign bond is priced so that

$$E_t(m_{t+k} r^j_{t+k} S^j_{t+k}) = S^j_t. \tag{2}$$

Equating (2) and (3) and noting that the return on the foreign bond is observed at time t leads to

$$E_t(P^j_{t+k,k}) = \frac{-\text{cov}_t(m_{t+k}, P^j_{t+k,k})}{E_t(m_{t+k})}, \tag{3}$$

where $P^j_{t+k,k} = (S^j_{t+k} - F^j_{t,k})/S^j_t$ is the usual definition of an ex post foreign exchange risk premium and $\text{cov}_t(\cdot)$ is the conditional covariance operator. The risk premium is standardized by S^j_t in order to induce stationarity in the risk premium series for estimation purpose. According to (3), since the conditionally anticipated marginal rate of substitution is always positive, the expected profits from forward exchange speculation will be nonzero when the marginal rate of substitution is conditionally correlated with the rewards from speculation.

The relationship (3) can alternatively be stated so as to avoid the use of marginal utility of consumption by substituting it with the return on a benchmark portfolio. First, note that the return on a risk-free asset is defined by

$$R^F_{t+k} = [E_t(m_{t+k})]^{-1}. \tag{4}$$

Next, observe that an ex post return on the benchmark portfolio that satisfies the equilibrium condition is

$$R^b_{t+k} = m_{t+k}[E_t(m^2_{t+k})]^{-1}. \tag{5}$$

This return is not unique as any portfolio that lies on the conditional mean-variance frontier can be used as the reference portfolio. When markets are complete, (5) corresponds to the minimum conditional second moment return (Hansen, Richard, and Singleton [17]). Substituting (4) and (5) into (3) results in

$$E_t(P^j_{t+k,k}) = \beta^j_{t,k}E_t(R^b_{t+k} - R^F_{t+k}), \tag{6}$$

where $\beta^j_{t,k} = \mathrm{cov}_t(R^b_{t+k}, P^j_{t+k,k})/\mathrm{var}_t(R^b_{t+k})$ and $\mathrm{var}_t(\cdot)$ is the conditional variance operator. Specification (6) provides the basis for the ILVM studied in this paper. The ILVM is similar to the Sharpe-Lintner single-period capital asset pricing model (CAPM) in that it measures the risk of an asset relative to a portfolio that lies on the mean-variance frontier as a single constant beta. It is, however, different from the CAPM in that it explicitly allows for shifts in the investment opportunity sets over time. It also acknowledges Roll's [26] critique of asset-pricing models concerning the identification of the benchmark portfolio by treating the reference portfolio as an unobservable variable.

II. Methodological Issues

Specifications (3) and (6) are both derived from the same equilibrium condition. Both are highly general with little empirical content, and additional restrictions are needed to test the models. It is, therefore, important to emphasize that the tests are joint tests of (3) or (6) with the auxiliary constraints. Specifically, in order to test (3), it is necessary to assume a specific preference function of consumption. Mark [24] uses a utility function that displays constant relative risk aversion and obtains results that are inconsistent with the joint behavior of the consumption series and the forward exchange risk premiums. While this approach has the advantage of leaving unspecified the nature of the forcing variables in the economy, it is contingent on the choice of a valid utility function and the availability of reliable consumption data. Therefore, his rejection need not indicate a violation of the Euler condition but may merely indicate the failure of the additional assumptions.

On the other hand, the ILVM proceeds by imposing two different assumptions. First, the beta coefficient is assumed to be a constant. Since the conditional covariances and conditional variances are, in general, functions of the variables available at time t, the ILVM may fail when the sample data are incompatible with this assumption. Second, the excess return on the benchmark portfolio is acknowledged to be unobservable and is treated as a latent variable. Specifically, since expectations are assumed to be rational, by the law of iterated projections, the expectation of the benchmark portfolio's excess return can be conditioned on a subset of I_t. The best linear projection of the benchmark portfolio's excess return is then

$$R^b_{t+k} - R^F_{t+k} = \alpha' Z_t + \varepsilon_{t+k,k},$$

where $\varepsilon_{t+k,k}$ is the forecast error that is orthogonal to Z_t; $\alpha' = (\alpha_1, \alpha_2, \cdots, \alpha_q)$ is a q-vector of parameters; and Z_t is a $(q \times 1)$ column vector of elements in I_t with typical element z_{it} that is used to predict the excess returns of the hedge portfolio.

Under the two additional assumptions, the risk-return relationship (6) becomes

$$P^j_{t+k,k} = \phi^j_k Z_t + e^j_{t+k,k}, \tag{7}$$

where $\phi^j_k = \beta^j_k \alpha'$ (with beta no longer a function of date-t information);

$$e^j_{t+k,k} = u^j_{t+k,k} + \beta^j_k \varepsilon_{t+k,k};$$

and

$$u^j_{t+k,k} = P^j_{t+k,k} - E_t(P^j_{t+k,k})$$

is the rational expectations error. The distribution of the disturbance term $e^j_{t+k,k}$ has the following properties:

(i) $E(e^j_{t+k,k} e^j_{t+k-i,k}) = 0$ for $i \geq k$ when the projection error $\varepsilon_{t+k,k} = 0$;
(ii) $E_t(e^j_{t+k,k} e^j_{t+k-i,k}) = f(I_t)$, where $f(\cdot)$ denotes a function; and
(iii) $E(e^j_{t+k,k} z_{it}) = 0$.

The first property (i) specifies that, when nonoverlapping data are used and there are no projection errors, there is no serial correlation in the residuals. Since projection errors can be serially correlated with unknown order, this condition is generally violated. In addition, when maturities longer than one month are analyzed, only a small number of observations are available for the current floating exchange rate period. This is important because Z_t are predetermined variables and are not strictly exogenous, and, therefore, the estimations must be based on asymptotic distribution theory. Thus, weekly data are used in the analysis. This results in a moving average (MA) error process of order $n - 1$ for an n-week maturity contract. For the cross-maturity estimates, since data from different maturities are employed, the MA error process corresponds to that for the longest maturity used.[2]

The second property (ii) states that, for all i, the conditional covariance of the residuals is a function of the information available at time t. Thus, general time-

[2] Nothing in the estimation need change to account for possible serial correlation in the projection errors. This is because, in practice, the size of the rational expectations error is likely to dominate since unanticipated exchange rate changes are quite large compared to plausible values of the expected movements in the return on the benchmark portfolio. Furthermore, the Newey-West [25] adjustment used below allows for potentially infinite order autocorrelation as the size of the sample increases.

varying conditional second moments are allowed and it is unnecessary to specify their exact form. This is important because studies by Cumby and Obstfeld [6], Domowitz and Hakkio [7], Giovannini and Jorion [13], and others have documented the presence of conditional heteroscedasticity in forward foreign exchange risk premiums.

The third property (iii) is the orthogonality condition exploited in the estimation.

Consider the case of estimating (7) simultaneously for $j = 1, 2, \cdots, p$ while holding the maturity k fixed. This results in a p equation system:

$$P_{t+k,k} = \phi_k Z_t + e_{t+k,k}, \tag{8}$$

where $P_{t+k,k}$ is a $(p \times 1)$ vector of ex post risk premiums, ϕ_k is a $(p \times q)$ matrix of parameters with typical element $\beta_k^j \alpha_i$, and $e_{t+k,k}$ is a $(p \times 1)$ vector of residuals. An analogous procedure is used to test the single-latent-variable specification for risk premiums of various maturities for a specific exchange rate.

An appropriate procedure for estimating (8) with predetermined variables in the presence of serially correlated and conditionally heteroscedastic errors is Hansen's [15] generalized method of moments (GMM). Define the $(pq \times 1)$ vector function $f(P_{t+k,k}, Z_t, \phi_k) = e_{t+k,k} \otimes Z_t$, where $\otimes$ denotes the Kronecker product. The parameters are then chosen to minimize the criterion function:

$$J_T(\phi) = g_T(\phi)'\, W_T g_T(\phi), \tag{9}$$

where $g_T(\phi)$ is the sample mean of the orthogonality conditions $E[f(P_{t+k,k}, Z_t, \phi_k)]$, $E[\cdot]$ is the unconditional expectations operator, and W_T is the $(pq \times pq)$ symmetric weighting matrix. Hansen [15] proves that the optimal W_T is

$$W_T = \{ \textstyle\sum_{i=-k+1}^{k-1} [T^{-1} \sum_{t=1}^{T} f(P_{t+k,k}, Z_t, \phi_k) f(P_{t+k-i,k}, Z_{t-i}, \phi_k)']\}^{-1}. \tag{10}$$

It is optimal in the sense that, among the class of estimators that employ the orthogonality conditions, it yields the smallest asymptotic covariance matrix of the parameters.[3] The asymptotic covariance matrix is $(d_T' W_T d_T)^{-1}$, where

$$d_T = \frac{1}{T} \sum_{t=1}^{T} \frac{\partial f(P_{t+k,k}, Z_t, \phi_k)}{\partial \phi_k}. \tag{11}$$

There are a total of pq elements in the ϕ_k matrix. However, only the coefficients in the first row and the first column of the matrix need to be estimated. This is

[3] A problem often encountered with using the W_T estimator in practice is that it is not constrained to be positive definite by construction. An alternative to (10) is to use frequency domain procedures to estimate W_T, as suggested by Hansen [15]. This alternative, however, does not exploit the presence of only a finite number of autocovariances present in the weighting matrix and may, therefore, result in less precise estimates. Instead, I obtain a positive-definite matrix by adjusting W_T with a modified Bartlett lag window,

$$w(\tau) = \left[\frac{k - |i|}{k}\right]^{\tau}, \qquad i = -k+1, -k+2, \cdots, k-2, k-1,$$

by giving less weight to the higher autocovariances of estimated residuals. In the estimation, a τ value of one was found to be sufficient to yield positive-definite W_T. Thus, a triangular lag window is used, and a proof of the consistency of the resulting W_T estimator is provided by Newey and West [25].

because β_k^1 can be standardized at one since the benchmark portfolio is unobservable and because of the extensive nonlinear cross-equation restrictions in the system. This gives a total of $(p - 1) + q$ coefficients to be estimated. The remaining $(p - 1)(q - 1)$ overidentifying constraints provide the testable restrictions imposed by the ILVM and can be used to test the validity of the model. A theorem by Hansen [15] is used for this purpose. It states that T times the minimized value of $J_T(\phi)$ is asymptotically distributed as chi-square with $(p - 1)(q - 1)$ degrees of freedom.

III. Empirical Results

The ILVM (7) stipulates that the forward foreign exchange risk premiums move in proportion to the same single latent variable. The valuation formula is estimated for the nine bilateral U.S. dollar exchange rates over the period February 9, 1976 to May 9, 1986.[4]

A. Cross-Currency Single-Maturity Evidence

The nine bilateral exchange rates are jointly examined to determine whether the ILVM can fit the risk premiums for forward speculation in contracts with a specific maturity. The maturities covered are for one-month, three-month, six-month, and twelve-month contracts. It is also necessary to specify the instrumental variables for the latent variable. Hansen and Hodrick [16] use lagged risk premiums and Hodrick and Srivastava [20] use the forward premiums available at time t. Since Hodrick and Srivastava obtain more powerful results, and in order to utilize the currently available information variables, forward premiums are used here. Specifically, the instrumental variables are the forward premiums of the nine foreign currencies plus a constant. Thus, there are a total of ten alphas and eight betas from nine equations and ten instruments to estimate using ninety orthogonality conditions. The beta for the Canadian dollar is fixed at one.[5]

The results are presented in Table I. The estimations yield relatively precise estimates resulting in significant coefficients. The test of the overidentifying

[4] The beginning date corresponds to that used by Hansen and Hodrick [16] and avoids the "burning-in" period of the current floating exchange rate regime. The source of the data is the Weekly Review of the Harris Trust and Savings Bank. The data are as of the last day of the week, and one-month, three-month, six-month, and twelve-month forward contracts are computed from the forward premiums of nine major foreign currencies all quoted in terms of the U.S. dollar. The exchange rates are the Canadian dollar (CA), the British pound (UK), the Belgian franc (BE), the French franc (FR), the German deutschemark (GE), the Italian lira (IT), the Dutch guilder (NE), the Swiss franc (SW), and the Japanese yen (JA).

[5] In minimizing (9), W_T must be computed with consistent parameter estimates. A two-step procedure for minimizing (9) suggested by Hansen and Singleton [18] is therefore adopted. In the first step, the residuals from estimating the unconstrained version of (7) by OLS are used to construct the W_T matrix used in the first minimization of $J_T(\phi)$. The starting alpha values for the first minimization process are obtained from estimating the first equation of (7) unconstrained and the beta values are all set at one. The residuals from the first minimization process are then used to construct the consistent estimator of W_T, which is used in the second round minimization of $J_T(\phi)$. The initial parameter values for the second round are the parameter estimates from the first round. The differentiations required for the standard errors are calculated analytically and not numerically.

Table I

Intertemporal Latent-Variable Model: Cross-Currency Estimates[a]

	COEFF[b]	SE[c]	MSIG[d]	COEFF	SE	MSIG
	One-Month Maturity			Three-Month Maturity		
	(NOBS[e] = 532)			(NOBS = 523)		
β_{UK}	−1.671	0.196	0.000	−2.002	0.044	0.000
β_{BE}	1.914	0.218	0.000	−0.105	0.032	0.000
β_{FR}	0.165	0.176	0.174	0.067	0.030	0.013
β_{GE}	1.023	0.171	0.000	0.763	0.033	0.000
β_{IT}	−0.106	0.171	0.267	0.261	0.035	0.000
β_{NE}	1.282	0.201	0.000	0.469	0.032	0.000
β_{SW}	0.683	0.214	0.001	1.316	0.038	0.000
β_{JA}	0.987	0.198	0.000	−0.424	0.031	0.000
α_0	−0.003	0.001	0.001	−0.023	0.002	0.000
α_{CA}	−1.033	0.197	0.000	−2.898	0.140	0.000
α_{UK}	0.437	0.145	0.001	1.837	0.084	0.000
α_{BE}	−1.014	0.126	0.000	1.650	0.101	0.000
α_{FR}	0.006	0.045	0.450	−0.368	0.032	0.000
α_{GE}	1.679	0.287	0.000	1.258	0.203	0.000
α_{IT}	0.048	0.034	0.079	−0.240	0.042	0.000
α_{NE}	−1.026	0.152	0.000	0.064	0.082	0.219
α_{SW}	0.275	0.123	0.013	0.187	0.096	0.026
α_{JA}	−0.483	0.142	0.000	−0.826	0.067	0.000
χ^2	104.114		0.008	39.860		0.999
	Six-Month Maturity			Twelve-Month Maturity		
	(NOBS = 510)			(NOBS = 484)		
β_{UK}	0.378	0.027	0.000	3.166	0.018	0.000
β_{BE}	1.321	0.033	0.000	−1.239	0.015	0.000
β_{FR}	1.481	0.030	0.000	1.067	0.014	0.000
β_{GE}	0.988	0.032	0.000	0.326	0.020	0.000
β_{IT}	1.324	0.026	0.000	1.777	0.014	0.000
β_{NE}	2.260	0.040	0.000	0.363	0.020	0.000
β_{SW}	0.390	0.041	0.000	3.757	0.037	0.000
β_{JA}	−0.636	0.036	0.000	−1.165	0.011	0.000
α_0	−0.019	0.001	0.000	−0.094	0.001	0.000
α_{CA}	−0.353	0.046	0.000	0.365	0.016	0.000
α_{UK}	0.329	0.029	0.000	0.221	0.007	0.000
α_{BE}	−0.322	0.027	0.000	−0.235	0.012	0.000
α_{FR}	−0.632	0.018	0.000	0.002	0.000	0.000
α_{GE}	0.165	0.058	0.002	0.570	0.020	0.000
α_{IT}	0.594	0.016	0.000	−0.512	0.006	0.000
α_{NE}	−2.330	0.056	0.000	−0.619	0.009	0.000
α_{SW}	1.497	0.033	0.000	0.392	0.010	0.000
α_{JA}	0.905	0.029	0.000	0.636	0.009	0.000
χ^2	20.428		1.000	10.336		1.000

[a] The estimates are obtained using the generalized method of moments. The dependent variables are forward risk premiums and the instruments are forward premiums plus a constant. Both heteroscedasticity and serial correlation are allowed. The chi-square statistic tests the overidentifying constraints imposed by the latent-variable specification.

[b] Coefficient estimate.

[c] Standard error.

[d] Marginal significance level.

[e] Number of observations.

restrictions is provided by the chi-square statistic with seventy-two degrees of freedom. For the one-month forward contracts, the constrained system is rejected at the one percent significance level. This result is consistent with that found by Hodrick and Srivastava [20] and is contrary to that obtained by Hansen and Hodrick [16]. Whereas Hodrick and Srivastava employ twenty-one additional nonoverlapping observations compared with Hansen and Hodrick, the number of nonoverlapping observations used here is over two times bigger and covers a longer sample period. Additionally, as opposed to five exchange rates in the two previous studies, nine are estimated here.

In sharp contrast to the rejection obtained for the one-month contracts, the ILVM cannot be rejected for the three-month, six-month, and twelve-month contracts. In fact, for all three cases, a chi-square statistic at least as big will always occur under the null hypothesis.

B. Cross-Maturity Single-Currency Evidence

The ILVM is next examined to determine whether it can price the risk premiums of various maturities for each foreign currency. Since the instrumental vector is not confined to forward premiums that match the maturities of the risk premiums, for the present purpose, it includes the current forward premiums for the four different maturities plus a constant. Thus, there are four equations and five instruments, and this results in the estimation of eight coefficients from twenty orthogonality conditions. The beta for the one-month maturity is set to one. Since contracts of various maturities are involved, the data are all expressed as annualized percentage rates. The empirical evidence is reported in Table II. It shows that, for all nine foreign exchange rates, the data cannot reject the constraints imposed by the ILVM. With the exception of the French franc, the data also appear to be well-behaved, with standard errors that are small relative to the parameter estimates.

C. Discussion

The results indicate that foreign currency-specific factors operating at the short end of the forward maturity spectrum render the ILVM invalid. Since the rejection is for a single-latent-variable model, it is possible that the restrictions are consistent with a two-latent variable model. Gibbons and Ferson [12] state the constraints for the case of multiple risk premiums.[6] Specifically, the two-state-variable model constrains the typical element of the ϕ_k matrix in (8) to be $\sum_{s=1}^{2} \beta_{ks}^{j}\alpha_{si}$. Since the two reference portfolios are unobservable, the beta coefficients can be standardized by setting $\beta_{k1}^{1} = \beta_{k2}^{2} = 1$ and $\beta_{k2}^{1} = \beta_{k1}^{2} = 0$. For the nine-equation and ten-instrument system, this results in the estimation of thirty-four parameters and fifty-six overidentifying restrictions. The chi-square value for the test of the overidentifying restrictions is 84.03, which is significant at the one percent level. Thus, the two-latent-variable model is also strongly rejected.[7]

[6] However, in contrast to Gibbons and Ferson [12], the estimations here permit the use of data sampled more frequently than the forecast horizon as well as time-varying second moments.

[7] Extensions to more than two latent variables were not attempted as estimations prove to be difficult because of the large number of parameters and orthogonality conditions. For example, the three-latent-variable case involves estimation of forty-eight restricted coefficients with ninety orthogonality conditions.

Table II

Intertemporal Latent-Variable Model: Cross-Maturity Estimates[a]

Foreign Currency		β_{3M}	β_{6M}	β_{12M}	α_0	α_{1M}	α_{3M}	α_{6M}	α_{12M}	χ^2
CA	COEFF[b]	0.341	0.161	0.037	−3.662	−4.865	2.825	−1.312	1.322	8.896
	SE[c]	0.065	0.054	0.046	0.835	0.872	2.287	3.046	2.562	
	MSIG[d]	0.000	0.002	0.211	0.000	0.000	0.109	0.333	0.303	0.712
UK	COEFF	1.126	1.144	0.763	−9.287	1.204	−2.748	1.176	−4.189	5.670
	SE	0.056	0.085	0.079	1.412	0.708	1.169	5.011	5.032	
	MSIG	0.000	0.000	0.000	0.000	0.045	0.010	0.407	0.203	0.932
BE	COEFF	0.689	0.487	0.352	−2.626	0.630	−7.643	2.028	3.955	2.562
	SE	0.076	0.092	0.061	2.066	0.487	1.737	3.835	3.013	
	MSIG	0.000	0.000	0.000	0.102	0.098	0.000	0.299	0.095	0.998
FR	COEFF	1.157	7.345	0.718	0.054	0.062	−0.130	0.056	−0.001	5.922
	SE	2.541	22.25	2.117	0.186	0.192	0.399	0.178	0.002	
	MSIG	0.325	0.371	0.367	0.386	0.373	0.373	0.376	0.374	0.920
GE	COEFF	0.372	0.435	0.422	3.224	−10.24	14.54	−6.738	1.211	7.474
	SE	0.073	0.082	0.100	2.470	1.941	3.073	5.341	3.903	
	MSIG	0.000	0.000	0.000	0.096	0.000	0.000	0.104	0.378	0.825
IT	COEFF	0.769	0.497	0.308	−16.671	−.868	−1.757	1.302	−0.453	6.729
	SE	0.045	0.052	0.039	4.499	0.345	0.837	1.649	1.457	
	MSIG	0.000	0.000	0.000	0.000	0.006	0.018	0.215	0.378	0.875
NE	COEFF	0.938	0.706	0.512	8.771	−1.302	7.424	−14.136	3.301	8.001
	SE	0.080	0.089	0.065	2.330	0.922	3.015	3.984	3.319	
	MSIG	0.000	0.000	0.000	0.000	0.079	0.007	0.000	0.160	0.785
SW	COEFF	0.781	0.738	0.789	24.53	−4.189	0.398	9.655	−10.43	4.977
	SE	0.086	0.117	0.174	5.580	1.137	1.771	3.637	4.062	
	MSIG	0.000	0.000	0.000	0.000	0.000	0.411	0.004	0.005	0.958
JA	COEFF	0.922	0.827	0.964	15.714	2.418	−3.659	−2.620	−0.215	6.046
	SE	0.062	0.078	0.125	3.548	0.718	1.926	3.297	1.748	
	MSIG	0.000	0.000	0.000	0.000	0.000	0.029	0.214	0.451	0.914

[a] The estimates are obtained using generalized method of moments. The dependent variables are forward risk premiums and the instruments are forward premiums plus a constant. Both heteroscedasticity and serial correlation are allowed. The chi-square statistic tests the overidentifying constraints imposed by the latent-variable specification. The number of observations is 484.

[b] Coefficient estimate.

[c] Standard error.

[d] Marginal significance level.

The rejections, however, need not indicate the failure of the underlying model (6) to explain one-month forward exchange risk premiums as auxiliary assumptions were imposed to generate the testable hypotheses. Hodrick and Srivastava [20] suggest that the rejection is due to the failure of the beta parameters to remain stable over time. What the present analysis suggests is that, if beta instability is the culprit, given the other results in Table I and those in Table II, the causes of the beta instability are likely to be linked to short-run components specific to each foreign currency. Possible candidates for such components are short-term foreign real interest rates. Of course, since rational expectations is assumed throughout the analysis, market irrationality is also a distinct possibility.

The results also suggest that the previously observed rejection of the ILVM for the one-month forward contracts by Hodrick and Srivastava [20] is an exception rather than the rule. In a recent study, Campbell and Clarida [4] also fail to reject the ILVM for the term structure of Euromarket interest rates. Given the nature of the integration between the Eurocurrency interest rates and foreign exchange rates, this similarity may be expected. Indeed, the tests were repeated with Eurocurrency interest rate differentials instead of the forward premiums as instrumental variables without any qualitative differences in results.

Of course, this is not to say that the ILVM is the true model in cases the data fail to reject. It merely indicates that the ILVM is a reasonable model for the sample period covered for these cases. It also suggests that the model may be inappropriate in situations where cross-country effects that persist for relatively short periods of time are prevalent. The good news is that it is suitable for cross-foreign currency risk premiums for forward contracts of longer maturities and cross-maturity risk premiums.

It is also important to emphasize that, since the tests are based on asymptotic distribution theory, the results are contingent on an adequate sample size being used. It may also be contested that the tests are not powerful enough to reject, but the tests do reject for cross-currency one-month contracts where it may be argued that they are the least likely to.[8]

IV. Conclusion

This paper investigates the intertemporal pricing of forward foreign exchange contracts using an intertemporal latent variable model (ILVM). The results show

[8] In order to test the robustness of the results, the tests were repeated for different specifications. First, the forward rates may not match precisely with the subsequently realized spot rate. Therefore, for the longer maturities, the sensitivity of the results is explored with small variations in the number of overlapping periods. Second, instead of standardizing the beta in cross-currency tests for the Canadian dollar at one, the German deutschemark is set to one. Third, the Italian lira and the Belgian franc were dropped to form a seven-equation system for cross-currency tests. Fourth, for cross-maturity tests, the twelve-month forward contracts were dropped. Fifth, as mentioned above, Eurocurrency interest rate differentials as opposed to forward premiums were used as instruments in both sets of tests. In all cases examined, the results lead to the same conclusions as those in Tables I and II. It may also be desirable to estimate all nine exchange rates and four maturities simultaneously. This would result in estimating some two hundred orthogonality conditions, and, given the experience encountered in estimating ninety orthogonality conditions, this would likely prove to be prohibitively expensive and highly impractical. More importantly, in the event of a rejection, it is desirable to identify the nature of the rejection, and this purpose is better served by focusing on cross-currency and cross-maturity estimates separately.

that the ILVM is relatively robust with respect to pricing forward exchange risk premiums. Specifically, it is capable of explaining the cross-foreign currency risk premiums for specific maturities with mispricing occurring only at the short end of the maturity structure. It is also capable of explaining cross-maturity risk premiums for specific exchange rates. The results are also of interest for assessing asset-pricing models in general and the ILVM in particular as it identifies situations under which the models may hold. The evidence presented here suggests that the model may fail in the presence of nontrivial short-term country-specific effects. This provides some guidance in understanding more fully the reasons for the failure of the ILVM as well as the identification of state variables that produce the behavior observed here in future work. Possible extensions of the analysis here also include investigations of models with explicit specifications for the time variation in the conditional variances and covariances. A possible direction for this is the use of ARCH and GARCH processes described in Engle and Bollerslev [8].

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