NBER WORKING PAPER SERIES

TESTS OF RATIONAL EXPECTATIONS AND NO RISK PREMIUM IN FORWARD EXCHANGE MARKETS

David A. Hsieh

Working Paper No. 843

NATIONAL BUREAU OF ECONOMIC RESEARCH
1050 Massachusetts Avenue
Cambridge MA 02138

January 1982

I would like to thank Rudiger Dornbusch, Stanley Fischer, Lars Hansen and Robert Litterman for helpful comments. The research reported here is part of the NBER's research program in International Studies. Any opinions expressed are those of the author and not those of the National Bureau of Economic Research.

NBER Working Paper #843
January 1982

TESTS OF RATIONAL EXPECTATIONS AND NO RISK PREMIUM IN FORWARD EXCHANGE MARKETS

Abstract

This paper tests the hypothesis that traders have rational expectations and charge no risk premium in the forward exchange market. It uses a statistical procedure which is consistent under a large class of heteroscedasticity, and a set of data which takes into account the institutional features of the forward exchange market. The results show that inferences using this procedure are very different from those using the standard assumption of homoscedasticity.

David A. Hsieh
Graduate School of Business
University of Chicago
1101 East 58th Street
Chicago, IL 60637

312-962-7418

1. Introduction

This paper tests the "simple efficiency" hypothesis in forward exchange markets. "Simple efficiency," as defined in Hansen and Hodrick (1980), means that traders have rational expectations and that there is no risk premium in the forward exchange market. A testable implication of this hypothesis is that the error committed by the forward rate in forecasting the spot rate has a zero mean and is uncorrelated with any known information.

The present work represents an improvement over the existing literature in two respects. One, all researchers have assumed that the aforementioned forecast error is stationary. Their test statistics are not necessarily consistent under nonstationarity. This paper uses a statistical procedure which is consistent for a large class of heteroscedasticity. Inference results using this procedure are quite different from those using the standard assumption of homoscedasticity. Two, a new set of data is used, which takes into account the institutional features of the forward market in matching the appropriate spot rate to the forward rate.

Using the new statistical procedure and the new data, I found that the forecast error is correlated with variables which are assumed to be in traders' information set, such as past values of spot and forward rates. These results provide the strongest rejection of the simple efficiency hypothesis thus far in the published literature, and are consistent with the findings of the other authors, such as Geweke and Feige (1979) and Hansen and Hodrick (1980). This does not necessarily imply that the forward market is "inefficient." Traders may charge a risk premium, which is correlated with the spot and forward rates.

The paper is in five sections. The next section discusses the "simple efficiency" hypothesis. The test procedure is introduced in section three.

The construction of the data is presented in section four. Empirical results and some conclusions are contained in the final two sections.

2. Tests of the Simple Efficiency Hypothesis

The notation of this paper is as follows. $s(t)$ denotes the natural logarithm of the spot exchange rate at date t, and $f(t, n)$, the natural logarithm of the forward exchange rate contracted at date t for delivery date $t + n$.[1] $E[s(t + n)|I(t)]$ is the expectation of $s(t + n)$ conditioned on the information set $I(t)$, which is assumed to contain all present and past values of spot and forward rates, and the stochastic processes describing these rates.

$M[s(t + n)|t]$ is the "market's expectation" of $s(t + n)$ at time t. The hypothesis of rational expectations is that the market's expectation is the true expectation:

$$\text{H1:} \qquad M[s(t + n)|t] = E[s(t + n)|I(t)] \; .$$

The concept of a single expectation for the entire market may be uncomforting. But all traders are assumed to have the same information set. Under rational expectations, they will have the same expectation.

A second hypothesis is required to relate the forward exchange rate to expectations. The market is assumed to set the forward rate equal to the expected spot rate on delivery:

$$\text{H2:} \qquad f(t, n) = M[s(t + n)|t] \; .$$

---

[1] I use logarithms here for two reasons. One, most researchers have used logarithms. To make my results comparable, I shall follow the same convention. Two, if the stochastic processes are log-normal, then using logarithms yield an approximation of the return. This is discussed in Hansen and Hodrick (1981). I note here that the results do not change if I use rates of change in discrete time.

Sufficient conditions yielding this result are (a) that all traders are risk neutral, and (b) that markets are competitive. However, if traders are risk averse, then it is possible to obtain an equilibrium forward rate which is different from the market expectation, because of the presence of a risk premium. Thus, H2 is called the hypothesis of "no risk premium" in the forward rate.

Note that H1 and H2 are totally independent hypotheses. Traders may have rational expectations, but still require a risk premium for forward contracts. If they are risk averse, they may (rationally) expect a loss, in order not to have to carry any exchange risk themselves.

Also, H1 and H2 are not separately testable, because $M[s(t + n)|t]$ cannot be observed. However, they jointly imply:

$$f(t, n) = E[s(t + n)|I(t)] \tag{2.1}$$

which is called the "simple efficiency" hypothesis in the literature.

A testable implication of (2.1) is the following. Define the n-period forecast error $u(t, n)$ of the forward rate:

$$u(t, n) = s(t) - f(t - n, n) \ . \tag{2.2}$$

Then the simple efficiency hypothesis implies that $u(t, n)$ has zero mean and is uncorrelated with any information in $I(t - n)$, which is assumed to contain $\{s(t - n - j), f(t - n - j, n): j = 0, 1, 2, \ldots\}$.

Thus, the simple efficiency hypothesis implies:

T1: $u(t, n)$ has zero mean and is uncorrelated with $u(t - n, n)$

T2: $u(t, n)$ has zero mean and is uncorrelated with the n-period holding yield $r(t - n, n)$ and the n-period forward discount $d(t - n, n)$, which are defined by:

(2.3) $$r(t, n) = s(t) - s(t - n)$$

(2.4) $$d(t, n) = s(t) - f(t, n)$$

T1 was first tested by Geweke and Feige (1979). Using quarterly data on the three month forward rate, i.e., $n = 1$, they regressed $u(t, 1)$ on itself lagged once and a constant term:

(2.5) $$u(t, 1) = a + b\, u(t - 1, 1) + e(t, 1) .$$

The null hypothesis is that $a = 0 = b$ and $e(t, 1)$ is serially uncorrelated. Assuming that $u(t, 1)$, and hence $e(t, 1)$, is covariance stationary, they tested $a = 0 = b$ using the standard F-statistic.

Hansen and Hodrick (1980) tested a more general version of (2.5). Using weekly data for the three month forward rate, i.e., $n = 13$, they regressed $u(t, 13)$ on a constant term and lags of $u(t - 13, 13)$:

(2.6) $$u(t, 13) = a + B(L)\, u(t - 13, 13) + e(t, 13) ,$$

where $B(L)$ is a polynomial in the lag operator L. The null hypothesis is $a = 0 = B(L)$, and $e(t, 13)$ is a moving average of order 12, assuming stationarity. Hansen and Hodrick noted that ordinary least squares (OLS) yields consistent estimates of the coefficients in (2.6). The test of the null hypothesis $a = 0 = B(L)$ was conducted with a covariance matrix which accounted for the fact that the error term was a moving average.

This paper differs from the above work in two respects. First, weekly data on the seven-day forward exchange rate is used. (Note that $n = 1$.) This means I can employ as many observations as Hansen and Hodrick, while retaining serial uncorrelation of the error term, as in Geweke and Feige. Second, I use a heteroscedastic-consistent covariance estimate of the OLS coefficients, relaxing the assumption of stationarity made in the previous work. This estimator is discussed in the next section.

3. A Heteroscedastic-Consistent Covariance Estimator of OLS

Hsieh (1981) proved that a heteroscedastic-consistent covariance estimator for OLS can be obtained for time series regressions. The estimator is stated in the general regression model:

(3.1) $$y(t) = x(t)'\beta + \varepsilon(t), \quad t = 1, \ldots, T$$

where $x(t) = [x_1(t) \ldots x_k(t)]'$.

It is assumed that:

(i) $E[\varepsilon(t) | x(t), x(t-1), \ldots, \varepsilon(t-1), \ldots] = 0$, with probability 1,

(ii) $E[x(t)\, x(t)'] = V_t$

(iii) $\sum_t V_t/T \to V$, as $T \to \infty$, where V is a positive definite matrix.

(iv) $\sum_t^T E[\varepsilon(t)^2 x(t)\, x(t)']/T \to M$, as $T \to \infty$, where M is a positive definite matrix.

(v) There exists $\delta > 0$, $B > 0$, such that

$$E[|\varepsilon(t)^2 x_i(t) x_j(t) x_k(t)|^{1+\delta}] \leqq B, \quad \text{for all } i, j, k, \text{ and } t$$

$$E[|x_i(t) x_j(t) x_k(t) x_m(t)|^{1+\delta}] \leqq B, \quad \text{for all } i, j, k, m \text{ and } t.$$

(vi) For all $\{\alpha_{ij} : i = 1, \ldots, k, \quad j = i, \ldots, k\}$,

$$\sum_{i=1}^{k}\sum_{j=i}^{k}\sum_{t}^{T} E\Big[\big|E[\varepsilon(t)^2 x_i(t)x_j(t)\alpha_{ij}] - E[\varepsilon(t)^2 x_i(t)x_j(t)\alpha_{ij} | x(t-1), \ldots, \varepsilon(t-1), \ldots]\big|\Big]$$

$$= 0\Big(\sum_{i=1}^{k}\sum_{j=1}^{k}\alpha_{ij}E[\varepsilon(t)^2 x_i(t)x_j(t)]\Big) \; .$$

Then, (a) $b_{OLS} = (X'X)^{-1}(X'y) \to \beta$ in probability.

(b) $T^{1/2}(b_{OLS} - \beta)$ has an asymptotic normal distribution, with mean 0 and covariance $V'M(V')^{-1}$.

(c) $V^{-1}MV^{-1'}$ can be consistently estimated by:

$$V_{HC} = \Big(\frac{X'X}{T}\Big)^{-1} \Big(\sum_{t} u(t)^2\; x(t)\; x(t)'/T\Big) \Big(\frac{X'X}{T}\Big)^{-1} \tag{3.2}$$

where $u(t) = y(t) - x(t)b_{OLS}$.

Note that if $\varepsilon(t)$ is conditionally (and unconditionally) homoscedastic, i.e., $E[\varepsilon(t)^2 | X(t), X(t-1), \ldots, \varepsilon(t-1), \ldots] = \sigma^2$ with probability 1, then $V^{-1}M(V^{-1})' = \sigma^2 V^{-1}$, which can be consistently estimated by:

$$V_{OC} = s^2\Big(\frac{X'X}{T}\Big)^{-1}, \tag{3.3}$$

where s^2 is the sum of squared residuals divided by the number of observations. (No adjustment is made for the degrees of freedom lost in estimating the coefficient b_{OLS} since this is only an asymptotic result.) Under homoscedasticity, V_{HC} and V_{OC} tend to be the same matrix asymptotically.

If $\varepsilon(t)$ is indeed heteroscedastic, then V_{OC} may not be consistent, but V_{HC} is always consistent, provided the assumptions (i) through (vi) hold. The difference between V_{HC} and V_{OC} is:

$$V_{HC} - V_{OC} = \left(\frac{X'X}{T}\right)^{-1}\left(\sum_{t}^{T} (\varepsilon(t)^2 - s^2)x(t)x(t)'/T\right)\left(\frac{X'X}{T}\right)^{-1} .$$

Asymptotically, V_{OC} and V_{HC} tend to the same matrix, if $\varepsilon(t)^2$ and $x(t)x(t)'$ are not correlated. V_{OC} tends to underestimate (overestimate) the true covariance if $\varepsilon(t)^2$ and and $x(t)x(t)'$ are positively (negatively) correlated. Since homoscedasticity is not implied by the simple efficiency hypothesis, I prefer to conduct T1 and T2 using V_{HC} rather than V_{OC}, although both sets of results are reported.

4. Construction of the Data

Most empirical studies of the forward exchange market ignore the timing of delivery of forward contracts. The most important feature is that forward contracts do not have constant lengths. For example, a one-month contract sold on July 18 is due on August 18, assuming that August 18 is a business day. (See Riehl and Rodriquez (1977) for a detailed discussion.) This institutional feature means that forward rates and spot exchange rates must be properly matched if tests of the simple efficiency hypothesis are to be conducted.

Another interesting feature of forward markets is that the forward contract is delivered two business days after the contract is due. In the above example, delivery will take place on August 20, assuming the 19th and the 20th are both business days. For a hedger, i.e., someone holding a covered position, this means that the one-month forward contract is actually longer than a month. But for a speculator, i.e., someone holding an open position, this is not true. He must purchase foreign exchange on the spot market to cover his position. Since spot transactions are also delivered two business days after the trades are made, a speculator will trade in the spot market at the time when the forward contract is due, i.e., on August 18, and not when the contract is delivered, i.e., August 20.

This paper matches up the forward rates and spot rates for a speculator using a seven-day forward contract. This has several advantages: (i) the length of the contracts are constant, (ii) many non-overlapping observations are available, and (iii) the nonoverlapping observations allow the use of the heteroscedastic-consistent covariance estimator for the OLS coefficients.

Covered Interest Arbitrage

Multinational banks frequently deal in one-, two-, and seven-day forward contracts. But data are not publically available. However, using the "covered interest arbitrage" formula, I can construct forward contracts from seven-day eurocurrency rates and the spot exchange rates.

Let me illustrate this for the case of the U.S. dollar/German mark rate. I can buy forward marks for dollars in two ways. I can buy a seven-day forward contract at $F(t, n)$ marks per dollar. Or, I can borrow from the eurodollar market at the rate $i(t, 1)$, sell the dollars in the spot exchange market for $S(t)$ marks, and deposit the marks in a eurobank at the rate $i^*(t, 1)$. The two methods should lead to the same number of forward marks for each dollar (aside from brokerage costs.) This equivalence is the "covered interest arbitrage" condition:

$$(4.1) \qquad F(t, 1) = S(t)\,(1 + i(t, 1))/(1 + i^*(t, 1)) \; .$$

Eaker (1980) showed that arbitrage opportunities using ninety-day forward and eurocurrency rates essentially do not exist.

This reader may object to the construction of the forward rate in this manner, because there are brokerage costs. Conversations with a foreign exchange trader revealed that brokerage costs are quite small--about 12 U.S. dollars per million U.S. dollars of transactions.

Data Sources

Seven-day forward rates are constructed in this manner for seven other currencies--the British pound, the French franc, the Swiss franc, the Dutch guilder, the Italian lira, the Canadian dollar, and the Japanese yen--in addition to the German mark. The spot exchange rates and eurocurrency interest rates are Friday closing rates in London, published by the Financial Times. The data start on June 9, 1978 and end on April 24, 1981, totaling 151 observations per exchange rate. The choice of currencies and dates are limited by the availability of data.

On several occasions, seven-day eurocurrency interest rates were not available (mostly in the case of the eurosterling.) They were replaced by the one-month rates if available, and with the rates from the previous day if one-month rates wre also not available. In addition, the London exchange was closed on April 4, 1980. The rates on April 3 were used instead. These imperfections may affect the results, particularly if these dates are outliers. (This is not the case.)

5. The Econometric Results

The econometric results are presented in Tables 1 through 4. The first two tables contain the results of T1, which runs the regression:

$$u(t, 1) = a + B(L)\, u(t - 1, 1) + e(t, 1) \quad (5.1)$$

The null hypothesis is $a = 0 = B(L)$.

As discussed in Geweke and Feige (1979), two versions of this test are possible: the single market and the multi-market test. The difference lies in the amount of information available to the trader.

Let superscript j denote the j-th currency. The single market test assumes the trader in the j-th market has the information set $I^j(t) = \{r^j(t, 1), r^j(t-1, 1), \ldots, d^j(t, 1), d^j(t-1, 1), \ldots\}$. The multi-market test assumes that each trader has the information set $I(t) = \{r^j(t, 1), r^j(t-1, 1), \ldots, d^j(t, 1), d^j(t-1, 1), \ldots : j = 1, \ldots, 8\}$.

In the single market test, (5.1) is a system of 8 univariate regressions, of the form:

$$(5.2) \qquad u^j(t, 1) = a^j + \sum_k b^j_k u^j(t-k-1, 1) + e^j(t, 1), \quad j = 1, \ldots, 8 .$$

The result of testing $a^j = 0 = b^j{}_k$ using 10 lags, are in Table 1. The first column reports the test statistic computed under the standard procedure, which assumes homoskedasticity of $e^j(t, 1)$. This uses the OC covariance in (3.3). The second column reports the test statistic using the HC covariance in (3.2), which is consistent under a large class of heteroscedasticity of $e^j(t, 1)$. The latter procedure leads to a higher rate of rejection of the null hypothesis. In fact, only the Canadian dollar fails under OC at the five percent significance level, while three out of eight currencies fail under HC.

In the multi-market test, (5.1) is a vector autoregression:

$$(5.3) \qquad u^j(t, 1) = a^j + \sum_k \sum_{\ell=1}^{8} b^j_{k\ell}\, u^\ell(t-k-1, 1) + e^j(t, 1), \quad j = 1, \ldots, 8 .$$

The results of testing $a^j = 0 = b^j{}_{k\ell}$, using 2 lags, are in Table 2. The null hypothesis is rejected at the five percent level for five currencies under OC, and six currencies under HC.

In T2, the following regression is run:

$$(5.2) \qquad u(t, 1) = a + C(L)\, x(t-1, 1) + D(L)\, d(t-1, 1) + e(t, 1) .$$

As in T1, the test can be done in the single market and multi-market context. In the former case, the one-period forecast error of each currency is regressed on a constant term and a distributed lag of its holding yields and its forward discount. (Five lags are used.) The results, reported in Table 3, show that three of eight currencies are rejected at the five percent level under OC, and six of eight under HC.

In the latter case, each forecast error is regressed on a constant term and one lag of the holding yields and forward discounts of all eight currencies. The results, in Table 4, show that the null hypothesis is rejected in one of eight cases under OC, and all eight cases under HC, at the five percent level.

Table 1

Results of T1: single market

Regression: $u^j(t, 1) = a^j + \sum_{k=1}^{10} b_k^j u^j(t - k - 1, 1) + e^j(t, 1)$

Tests of $a^j = 0 = b_k^j$ using

j	OC[a]	HC[a]
France	15.03	18.28*
Germany	13.86	24.79***
United Kingdom	7.91	12.91
Switzerland	11.37	15.38
Netherlands	11.92	16.47
Canada	18.89*	26.13***
Italy	16.78	16.31
Japan	25.36***	31.58***

Period: September 29, 1978 to April 24, 1981 (139 observations)

* Significant at the 10 percent level
** Significant at the 5 percent level
*** Significant at the 1 percent level

[a] The statistic is distributed chi-square, with 11 degrees of freedom. Critical values are: 17.2750, 19.6751, and 24.7250 at the 10%, 5% and 1% levels, respectively.

Table 2

Results of T1: multi-market

Regression: $$u^j(t, 1) = a^j + \sum_{k=1}^{2} \sum_{\ell=1}^{8} b^j_{k\ell} u^e(t - k - 1, 1) + e^j(t, 1)$$

Tests of $a^j = 0 = b^j_{k\ell}$ using

j	OC[a/]	HC[a/]
France	31.94***	33.25**
Germany	31.93**	35.18**
United Kingdom	21.59	24.50
Switzerland	27.78**	31.40**
Netherlands	30.42**	26.88**
Canada	20.30	20.85
Italy	23.46	28.84**
Japan	36.23***	33.70***

Period: September 29, 1979 to April 24, 1981 (139 observations)

* Significant at the 10 percent level
** Significant at the 5 percent level
*** Significant at the 1 percent level

[a/]The statistic is distributed chi-square, with 17 degrees of freedom. Critical values are: 24.7690, 27.5871, and 33.4087 at the 10%, 5%, and 1% levels, respectively.

Table 3

Results of T2: single market

Regression: $u^j(t,1) = a^j + \sum_{k=1}^{5} b_k^j r^j(t-k-1,1) + \sum_{k=1}^{5} c_R^j d^j(t-k-1,1) + e^j(t,1)$

Tests of $a^j = 0 = b_k^j = c_k^j$

j	OC[a]	HC[a]
France	16.03	25.25***
Germany	10.75	61.91***
United Kingdom	7.43	17.54*
Switzerland	15.32	22.37**
Netherland	9.25	15.43
Canada	27.50***	50.09***
Italy	21.63**	19.01**
Japan	21.90**	21.18**

Period: September 29, 1978 to April 24, 1981 (139 observations)

* Significant at the 10 percent level
** Significant at the 5 percent level
*** Significant at the 1 percent level

[a] The statistic is distributed chi-square, with 11 degrees of freedom. Critical values are: 17.2750, 19.6751, and 24.7250 at the 10%, 5%, and 1% levels, respectively.

Table 4

Results of T2: multi-market

Regression: $u^j(t, 1) = a^j + \sum_{\ell=1}^{8} b^j_\ell r^e(t - 1, 1) + \sum_{\ell=1}^{8} c^j_e d^e(t - 1, 1) + e^j(t, 1)$

Tests of $a^j = 0 = b^j_k = c^j_\ell$

j	OC[a/]	HC[a/]
France	27.48*	41.20***
Germany	28.15*	40.00***
United Kingdom	17.68	35.58***
Switzerland	20.45	37.46***
Netherlands	27.54*	35.72***
Canada	24.44	31.91***
Italy	24.34	33.68***
Japan	24.86*	28.57**

Period: September 29, 1978 to April 24, 1981 (139 observations)

* Significant at the 10 percent level
** Significant at the 5 percent level
*** Significant at the 1 percent level

[a/] The statistic is distributed chi-square, with 17 degrees of freedom. Critical values are: 24.7690, 27.5871, and 33.4087 at the 10%, 5%, and 1% levels, respectively.

6. Conclusion

The empirical results are interesting for two reasons. First, they show that there is a large difference in inference between using the OC covariance (which assumes homoscedasticity) and the HC covariance (which allows for a large class of heteroscedasticity). This is particularly evident in the multi-market tests performed in Table 4. These results suggest that the OC covariance often overestimate the standard errors of the OLS coefficients.

Second, the results provide the strongest rejection of the simple efficiency hypothesis thus far in the published literature. However, these results are not strictly comparable to those of Feige and Geweke (1979) and Hansen and Hodrick (1980), because of the differences in forecast horizons.

There are at least three possibilities to account for the rejection of simple efficiency. One, traders may not know the full structural model of exchange rate determination. They may not know the intervention rule of the central banks, or they may not know some parameters of the model. In this case, serial correlation in forecast errors may be observed.

Two, there may be a risk premium in the forward market. This arises from many circumstances, including differences in risk preferences. Theoretical models which demonstrate this point are Solnik (1974), Grauer, Litzenberger, and Stehle (1976), Stockman (1978), Fama and Farber (1979), Frankel (1979), Roll and Solnik (1979), and Stulz (1980).

Three, traders may simply act in an irrational manner. Further study is needed to find out which of these alternative hypotheses is responsible for the failure of simple efficiency in the forward exchange market.

References

Eaker, M. R. (1980) "Covered interest arbitrage: new measurements and empirical results," Journal of Economics and Business, 32: 249-253.

Fama, E. F. and A. Farber (1979) "Money, bonds, and foreign exchange," American Economic Review, 69: 639-649.

Frankel, J. (1979) "Diversifiability of exchange risk," Journal of International Economics, 9: 379-396.

Frenkel, J. A. and R. M. Levich (1979) "Covered interest arbitrage: unexploited profits? Reply," Journal of Political Economy, 87: 418-422.

Geweke, J. and E. L. Feige (1979) "Some joint tests of the efficiency of markets for forward foreign exchange," Review of Economics and Statistics, 61: 334-341.

Grauer, F. L. A., R. H. Litzenberger, and R. E. Stehle (1976) "Sharing rules and equilibrium in an international capital market and uncertainty," Journal of Financial Economics, 3: 233-256.

Hansen, L. P. (1979) "The asymptotic distribution of least squares estimator of endogenous regressors and dependent residuals," Carnegie-Mellon University Working paper.

Hansen, L. P. (1980) "Large sample properties of generalized method of moments estimators," Carnegie-Mellon University Working paper.

Hansen, L. P. and R. J. Hodrick (1980) "Forward exchange rate as optimal predictors of future spot rates: an econometric analysis," Journal of Political Economy, 88: 829-853.

Hansen, L. P. and R. J. Hodrick (1981) "Risk averse speculation in the forward exchange market: an econometric analysis," Carnegie-Mellon University Mimeo.

Hsieh, D. A. (1981) "A heteroscedastic-consistent covariance estimator for time series regressions," University of Chicago, mimeo.

Levich, R. M. (1978) "Further results on the efficiency of markets for foreign exchange," Federal Reserve Bank of Boston Conference Series, no. 20.

Meese, R. A. and K. J. Singleton (1980) "Rational expectation with risk premia, and the market for spot and forward exchange," Internatonal Finance Discussion Paper, no. 165.

Riehl, H. and R. M. Rodriguez (1977) Foreign Exchange Markets, New York: McGraw-Hill.

Roll, R. and B. Solnik (1977) "A pure foreign exchange asset pricing model," Journal of International Economics, 7: 161-179.

Solnik, B. (1974) "An equilibrium model of the international capital market," Journal of Economic Theory, 86: 303-309.

Stockman, A. C. (1978) "Risk, information, and forward exchange rates," in J. A. Frenkel and H. G. Johnson, ed., The Economics of Exchange Rates, Reading, MA: Addison-Wesley.

Stulz, R. M. (1980) "The forward exchange rate and macroeconomics," University of Rochester, mimeo.

Theil, H. (1971) Principles of Econometrics, New York: John Wiley.