Journal of International Money and Finance (1986), **5**, 5–21

The Covariation of Risk Premiums and Expected Future Spot Exchange Rates

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Eugene Fama (1984) analyzed the variability and the covariation of risk premiums and expected rates of depreciation. We employ three statistical techniques that do not suffer from a potential bias in Fama's analysis, but we nevertheless confirm his findings. In contrast to his interpretation the results are not necessarily at variance with the predictions of a theoretical model of the risk premium. Increases in expected rates of depreciation of the dollar relative to five foreign currencies are positively correlated with increases in the expected profitability of purchasing these currencies in the forward market, and risk premiums have larger variances than expected rates of depreciation.

Although the theoretical and empirical literature on the efficiency of the forward foreign exchange market is now quite broad, there is anything but consensus on the issues. It is now well understood that rejection of the unbiasedness hypothesis is not a rejection of market efficiency, since a risk premium can separate the forward rate from the expected future spot rate, but only recently has effort been directed to development of explicit tests of formal models of the risk premium. The empirical success of these models can only be judged as limited.[1]

In a recent paper, however, Eugene Fama (1984) has noted that by viewing the forward premium as the expected rate of depreciation of the domestic currency relative to the foreign currency *plus* a risk premium, it is possible to estimate empirically the degree of variation of the risk premium over time and

* We thank the conference participants and especially the organizers, Don Schlagenhauf and Michael Melvin, for their comments. We thank the Center for the Study of Futures Prices at Columbia University for its support of this project. We also thank Lauren Feinstone, Robert Flood, Jacob Frenkel, Richard Meese, Maurice Obstfeld, Kenneth Singleton, and participants in seminars at the NBER 1984 Summer Institute, Northwestern University, and Washington University for their comments.

also to learn about the covariation of the risk premium with the expected rate of depreciation. In particular, Fama found that the covariance of these two variables is negative and sufficiently large to imply that the variance of the risk premium exceeds that of the expected rate of depreciation. Fama also found the results somewhat troublesome. He states (p. 327), 'A good story for negative covariation between P_t (the risk premium) and $E(S_{t+1} - S_t)$ (the expected rate of depreciation) is difficult to tell.'

In Section I, we investigate the plausibility of the finding of negative covariation. We first argue that it is intuitively plausible before investigating whether the intuitive reasoning is supported by the general equilibrium model developed by Lucas (1982). Fama (1984) noted that the Lucas model might be used for this purpose, but he did not pursue the analysis. We show that a sufficient condition for negative covariation is that the covariance between the ratio of expected nominal marginal rates of substitution of the two currencies and the expected ratio of the nominal marginal rates of substitution be negative. Since it is difficult to determine the sign or magnitude of this expression in general, we examine a second-order approximation to the expression. Based on this approximation, it appears that the covariance of the risk premium and the expected rate of depreciation can be positive only if the covariance between the intertemporal nominal marginal rates of substitution of the two currencies is large and positive. Since the analysis based on the approximation is not entirely satisfactory, we provide an example based on Cobb–Douglas utility functions and conditionally lognormal exogenous processes in which the covariance is indeed negative. This example economy is exactly the type of environment in which Fama claims to find difficulty justifying negative covariation. It has complete markets and flexible goods prices that satisfy purchasing power parity.

Fama's other empirical result, that the variance of the risk premium is greater than the variance of the expected rate of depreciation, also seems to be an important result. It is significant because it implies that a time-varying risk premium plays a fundamental role in the determination of spot and forward exchange rates. In contrast, most theoretical rational expectations models of exchange rate determination have focused almost exclusively on the expected rate of depreciation. This finding suggests that future work be devoted to understanding the importance of risk in foreign exchange markets.

Because of the potential importance of these findings, it is essential to investigate whether the statistical procedures employed in Fama's analysis are correct. Under the hypothesis that there is a time-varying risk premium, there is reason to suspect that the standard errors reported by Fama are somewhat biased. In Section II, we investigate the source of the bias which arises from the presence of weak serial correlation, and we argue that Fama's estimates, while biased, are probably not severely biased. We then provide alternative techniques for testing the hypothesis of negative covariation that are robust to the presence of serial correlation, and we find that our statistical analysis supports Fama's results.

Concluding remarks and some qualifications of our analysis are contained in Section III.

I. Consistency with Theory

The empirical results contained in Fama (1984) demonstrate that the covariation of the expected rate of change of the exchange rate and the risk premium on a forward contract is negative. Let S_t be the spot exchange rate of US dollars per unit of foreign currency, and let F_t be the one period forward rate. The risk premium, P_t, is defined to be the expected profit from buying US dollars on the forward market, $P_t \equiv F_t - E_t(S_{t+1})$ where $E_t(\cdot)$ is the conditional expectation operator. Consequently, the empirical results tell us that when there is an increase in the expected rate of depreciation of the US dollar relative to the foreign currency, *i.e.* $E_t[(S_{t+1} - S_t)/S_t]$ increases, there is a decrease in the expected profit from buying US dollars in the forward market, *i.e.* P_t falls.

At an intuitive level this covariation may seem puzzling, since one might argue that the higher the expected depreciation of the dollar, the higher might be the required expected nominal return on a dollar denominated security. However, the risk premium in the empirical work is the expected profit from purchasing dollars forward. Hence, the expected profit is *not* dollar denominated; it is denominated in foreign currency. The appropriate dollar denominated profit is $(S_{t+1} - F_t)$. It is obtained by selling dollars in the forward market for foreign currency and using that foreign currency to buy dollars at the future spot rate. The expected profit from selling dollars forward is $(-P_t)$. Hence, the covariation of $(-P_t)$ with the expected rate of depreciation of the dollar relative to foreign currencies is positive. Consequently, at an intuitive level there is no apparent inconsistency in finding negative covariation between P_t and $E_t[(S_{t+1} - S_t)/S_t]$.

The intuitive reasoning is also supported by an examination of the risk premium within the context of a formal general equilibrium model developed by Lucas (1982). In Hodrick and Srivastava (1984), we demonstrate how the Lucas framework can be extended to price forward contracts. A central idea in such intertemporal asset pricing frameworks is that nominal risk free bills in the domestic and foreign currencies are priced such that the inverse of the known nominal return is equal to the conditional expectation of the intertemporal marginal rate of substitution of the currency. When this is true, the forgone marginal utility of holding a one period bill is equal to the expected discounted marginal utility of the return on the bill.

In the Lucas model, agents in two countries have identical preferences but different stochastic endowments of two goods. In period t, agents in country 0 get $2X_t$ of good X and nothing of good Y while agents of country 1 get $2Y_t$ of Y and nothing of X. Preferences are given by

$$\langle 1 \rangle \qquad E_0\left\{\sum_{t=0}^{\infty} \beta^t U(X_{it}, Y_{it})\right\}, \qquad 0 < \beta < 1.$$

where X_{it} and Y_{it} are the consumptions of X and Y of the representative agent in country i, $i=0, 1$. The real state is $\psi_t = (X_t, Y_t)$ which is a realization of a known Markov process. Since agents are assumed to be able to trade in markets that are complete in the sense of Arrow (1964) and Debreu (1959), they will share risk perfectly and always consume half of each good. Consequently, the relative price of Y in terms of X is

$$\langle 2 \rangle \qquad p_y(\psi_t) = U_y[X_t, Y_t]/U_x[X_t, Y_t]$$

where U_y and U_x are marginal utilities at period t.

Agents are required to purchase the good of a country with the money of that country, and agents know ψ_t when they make trades in money and securities.[2] Therefore, the dollar price of X and the pound price of Y are $P_x(\psi_t, M_t) = M_t/2X_t$ and $P_y(\psi_t, N_t) = N_t/2Y_t$ where M_t and N_t are 'dollars' and 'pounds', the monies of country 0 and country 1. Monies also follow known Markov processes.

The exchange rate is found by arbitrage:

$$\langle 3\rangle \qquad S_t(\psi_t, M_t, N_t) = [P_x(\psi_t, M_t)/P_y(\psi_t, N_t)]\, p_y(\psi_t) = \frac{M_t Y_t}{N_t X_t} p_y(\psi_t).^2$$

Define $\pi_t^M = [1/P_x(\psi_t, M_t)]$ and $\pi_t^N = [1/P_y(\psi_t, N_t)]$ to be the purchasing powers of the two monies over X and Y respectively. The expected value of one dollar in one period in terms of marginal utility is $E_t[\beta U_x(\psi_{t+1})\pi_{t+1}^M]$. If b_{xt} is the dollar price at t of one dollar delivered at $t+1$, then

$$\langle 4\rangle \qquad b_{xt} = E_t[\beta U_x(\psi_{t+1})\pi_{t+1}^M / U_x(\psi_t)\pi_t^M].$$

A similar argument gives the period t pound price of one pound delivered at $t+1$ as

$$\langle 5\rangle \qquad b_{yt} = E_t[\beta U_y(\psi_{t+1})\pi_{t+1}^N / U_y(\psi_t)\pi_t^N].$$

The discount bill prices are the conditional expectations of the intertemporal marginal rates of substitution of the two currencies:

$$\langle 6\rangle \qquad \mathcal{Q}_{t+1}^M = \beta U_x(\psi_{t+1})\pi_{t+1}^M / U_x(\psi_t)\pi_t^M$$

$$\langle 7\rangle \qquad \mathcal{Q}_{t+1}^N = \beta U_y(\psi_{t+1})\pi_{t+1}^N / U_y(\psi_t)\pi_t^N.$$

The intertemporal marginal rate of substitution of money is an index of the change in the purchasing power of money weighted by the intertemporal marginal rate of substitution of goods.

By covered interest arbitrage, a one period forward exchange rate must satisfy

$$\langle 8\rangle \qquad b_{yt} = (F_t/S_t)b_{xt}.$$

We use equations ⟨2⟩ through ⟨8⟩ to develop an analysis of the covariation of the risk premium and the expected rate of depreciation. In order to conserve notation, let $s_{t+1} \equiv (S_{t+1} - S_t)/S_t$ and let $p_t \equiv [F_t - E_t(S_{t+1})]/S_t$. Then it can be demonstrated that the following must hold:

$$\langle 9\rangle \qquad p_t = [E_t(\mathcal{Q}_{t+1}^N)/E_t(\mathcal{Q}_{t+1}^M)] - E_t(\mathcal{Q}_{t+1}^N/\mathcal{Q}_{t+1}^M)$$

$$\langle 10\rangle \qquad E_t(s_{t+1}) = E_t(\mathcal{Q}_{t+1}^N/\mathcal{Q}_{t+1}^M) - 1.$$

Consequently, the covariance of the risk premium in ⟨9⟩ and the expected rate of depreciation in ⟨10⟩ is

$$\langle 11\rangle \qquad C[p_t, E_t(s_{t+1})] = C\left[\frac{E_t(\mathcal{Q}_{t+1}^N)}{E_t(\mathcal{Q}_{t+1}^M)}, E_t\left(\frac{\mathcal{Q}_{t+1}^N}{\mathcal{Q}_{t+1}^M}\right)\right] - V\left[E_t\left(\frac{\mathcal{Q}_{t+1}^N}{\mathcal{Q}_{t+1}^M}\right)\right]$$

where $C(\cdot, \cdot)$ denotes the unconditional covariance operator and $V(\cdot)$ is the unconditional variance. Thus, a sufficient condition for negative covariation is

that the covariance term on the right hand side of ⟨11⟩ be negative. However, we have been unable to determine the sign or magnitude of this term in general. To get some insight into conditions implying a negative covariance between p_t and the expected rate of change of the spot rate, $E_t(s_{t+1})$, we consider a second order approximation to the expressions in ⟨9⟩ and ⟨10⟩. Second order approximation to ⟨10⟩ yields

$$\langle 12\rangle \qquad E_t(s_{t+1}) = \frac{E_t(Q^N_{t+1})}{E_t(Q^M_{t+1})} + \frac{V_t(Q^M_{t+1})}{[E_t(Q^M_{t+1})]^2} \times \left\{ \frac{E_t(Q^N_{t+1})}{E_t(Q^M_{t+1})} - \frac{C_t(Q^N_{t+1}, Q^M_{t+1})}{V_t(Q^M_{t+1})} \right\} - 1$$

which is obtained by expanding (Q^N_{t+1}/Q^M_{t+1}) around $E_t(Q^N_{t+1})$ and $E_t(Q^M_{t+1})$ and taking the conditional expectation. The same approximation in ⟨9⟩ yields

$$\langle 13\rangle \qquad p_t = -\frac{V_t(Q^M_{t+1})}{[E_t(Q^M_{t+1})]^2} \left\{ \frac{E_t(Q^N_{t+1})}{E_t(Q^M_{t+1})} - \frac{C_t(Q^N_{t+1}, Q^M_{t+1})}{V_t(Q^M_{t+1})} \right\}.$$

Intuition about the sign of the covariance between $E_t(s_{t+1})$ and p_t can now be determined by examining their comovements as the various terms in expressions ⟨12⟩ and ⟨13⟩ change. It is straightforward to verify that $E_t(s_{t+1})$ and p_t move in opposite directions with changes in $E_t(Q^N_{t+1})$, $V_t(Q^M_{t+1})$ and $C_t(Q^N_{t+1}, Q^M_{t+1})$. With any of these changes, therefore, $C[p_t, E_t(s_{t+1})] < 0$. The only potentially ambiguous change comes in the case of a change in $E_t(Q^M_{t+1})$. If $E_t(Q^M_{t+1})$ increases, for example, then $C[p_t, E_t(s_{t+1})] < 0$ if $C_t[Q^N_{t+1}, Q^M_{t+1}] < 0$. If this last term is positive but small relative to $V_t(Q^M_{t+1})$, then $C[p_t, E_t(s_{t+1})]$ can be positive only if the conditional covariance between the intertemporal marginal rates of substitution of the two currencies is large and positive.

A sufficient condition for $C[p_t, E_t(s_{t+1})] < 0$ when $E_t(Q^M_{t+1})$ changes is $C_t(Q^M_{t+1}, Q^N_{t+1}) < 0$. Only if this last term is large relative to $V_t(Q^M_{t+1})$ is $C[p_t, E_t(s_{t+1})]$ potentially greater than zero.

Since this comparative statistics exercise may not be satisfactory due to the approximation and the types of changes being considered, we now present an example economy in which $C[p_t, E_t(s_{t+1})]$ is negative.

Suppose the period t utility function is Cobb–Douglas, $U = AX_t^{\alpha} Y_t^{(1-\alpha)}$, which is evaluated at the equilibrium consumption levels. Then, the marginal utilities with respect to X and Y at time t are $U_x = \alpha A X_t^{(\alpha-1)} Y_t^{(1-\alpha)}$ and $U_y = (1-\alpha) A X_t^{\alpha} Y_t^{-\alpha}$. Assume X_{t+1}, Y_{t+1}, M_{t+1}, N_{t+1} are conditionally log normal, that there is no contemporaneous correlation, and let lower case letters of these variables denote natural logarithms of their upper case counterparts.[3] Then, $X_t = \exp(x_t)$, $Y_t = \exp(y_t)$, $M_t = \exp(m_t)$, $N_t = \exp(n_t)$, and the conditional distributions of the lower case variables are normal:

$$\langle 14a\rangle \qquad x_t \sim N(E_{t-1}x_t, \sigma^2_{xt}),$$

$$\langle 14b\rangle \qquad y_t \sim N(E_{t-1}y_t, \sigma^2_{yt}),$$

$$\langle 14c\rangle \qquad m_t \sim N(E_{t-1}m_t, \sigma^2_{mt}),$$

$$\langle 14d\rangle \qquad n_t \sim N(E_{t-1}n_t, \sigma^2_{nt}).$$

Then,

$$\langle 15a\rangle \qquad Q^N_{t+1} = \frac{\beta X^{\alpha}_{t+1} Y^{(1-\alpha)}_{t+1} N_t}{X^{\alpha}_t Y^{(1-\alpha)}_t N_{t+1}},$$

$$\langle 15b\rangle \qquad Q^M_{t+1} = \frac{\beta X^{\alpha}_{t+1} Y^{(1-\alpha)}_{t+1} M_t}{X^{\alpha}_t Y^{(1-\alpha)}_t M_{t+1}}, \text{ and}$$

$$\langle 15c\rangle \qquad \frac{Q^N_{t+1}}{Q^M_{t+1}} = \frac{M_{t+1} N_t}{N_{t+1} M_t}.$$

From ⟨15c⟩ and the distributional assumptions in ⟨14⟩, we find

$$\langle 16\rangle \qquad E_t\left(\frac{Q^N_{t+1}}{Q^M_{t+1}}\right) = \exp[E_t m_{t+1} - E_t n_{t+1} + n_t - m_t + \tfrac{1}{2}(\sigma^2_{mt+1} + \sigma^2_{nt+1})].$$

From ⟨15a⟩ and ⟨15b⟩ and the distributional assumption in ⟨14⟩ we find

$$\langle 17\rangle \qquad \frac{E_t(Q^N_{t+1}}{E_t(Q^M_{t+1})} = \exp\{n_t - m_t - E_t n_{t+1} + E_t m_{t+1} + \tfrac{1}{2}(\sigma^2_{nt+1} - \sigma^2_{mt+1})\}.$$

The results in ⟨16⟩ and ⟨17⟩ may now be used to demonstrate that

$$\langle 18\rangle \qquad p_t = -\exp\{n_t - m_t - E_t n_{t+1} + E_t m_{t+1} + \tfrac{1}{2}(\sigma^2_{nt+1} + \sigma^2_{mt+1})\} \times [1 - \exp(-\sigma^2_{mt+1})]$$

and

$$\langle 19\rangle \qquad E_t(s_{t+1}) = \exp\{n_t - m_t - E_t n_{t+1} + E_t m_{t+1} + \tfrac{1}{2}(\sigma^2_{nt+1} + \sigma^2_{mt+1})\} - 1.$$

Since p_t and $E_t(s_{t+1})$ in ⟨18⟩ and ⟨19⟩ are determined by the same six variables $z_t = \{n_t, {}^{\mathrm{T}} m_t, E_t m_{t+1}, E_t n_{t+1}, \sigma^2_{mt+1}, \sigma^2_{nt+1}\}$, and because $\partial p_t / \partial z_t$ is opposite in sign to $\partial E_t s_{t+1} / \partial z_t$ on an element by element comparison, the covariance of p_t and $E_t(s_{t+1})$ must be negative. Hence, contrary to Fama's analysis of this issue, we find negative covariation between the risk premium and the expected rate of depreciation to be quite consistent with economic theory.

To conclude this section, we note that there are factors other than a risk premium which can separate the forward rate from expected future spot rates. For example, suppose that in the Lucas model, the representative consumer in each country is risk neutral. Then, the relationship between F_t and $E_t(S_{t+1})$ is given by

$$\langle 20\rangle \qquad F_t = E_t(S_{t+1}) + \left\{\frac{E_t[\pi^N_{t+1} p_y(\psi_t)]}{E_t(\pi^M_{t+1})} - E_t\left[\frac{\pi^N_{t+1} p_y(\psi_{t+1})}{\pi^M_{t+1}}\right]\right\}.$$

The last term on the right hand side can vary through time as has been noted by Stockman (1978), Frenkel and Razin (1980), and Engel (1984), even though it is not a risk premium. Risk aversion magnifies the above deviation of F_t from $E_t(S_{t+1})$. Given ⟨20⟩, our empirical findings as well as Fama's findings can have an interpretation even with risk neutrality. We prefer to call any deviation a risk premium because of the large differences that characterize other expected asset returns. This evidence has strongly conditioned our prior beliefs regarding agents' risk aversion.

II. The Empirical Analysis

In this section we analyze whether the empirical technique used by Fama (1984) in his analysis of this issue is appropriate. Since the empirical analysis is relevant to any market containing spot and forward rates, it is desirable that estimators with correct properties are used. We argue here that some serial correlation may be present in the residuals of Fama's ordinary least squares and seemingly unrelated regressions which would potentially bias the standard errors. Since it appears in this case that the degree of the bias is not particularly severe, we do not investigate it formally. Instead, we discuss several econometric approaches to this issue that do not suffer from this potential bias.

As in the theoretical analysis, let F_t and S_t denote the forward and spot exchange rates, and let P_t be the risk premium. Both exchange rates are measured in domestic currency (US dollars) per unit of foreign currency. The risk premium is defined to be the expected profit from selling the foreign currency forward which is equivalent to buying US dollars forward. Market efficiency in the presence of a risk premium therefore implies

$$F_t = E_t(S_{t+1}) + P_t \tag{21}$$

where $E_t(S_{t+1})$ denotes the conditional expectation of S_{t+1} given information at time t. In ⟨21⟩, the forward rate is decomposed into two parts which are not observable to the econometrician, but each part is known to agents in the market. Since S_t is known at time t, ⟨21⟩ can be rewritten as:

$$(F_t - S_t)/S_t = E_t[(S_{t+1} - S_t)/S_t] + p_t \tag{22}$$

which decomposes the forward premium into the conditional expectation of the rate of change of the spot rate plus a normalized risk premium, $p_t \equiv P_t/S_t$. We use ⟨22⟩ in the empirical analysis because it is more likely than ⟨21⟩ to satisfy assumptions of covariance stationarity.[4] In order to simplify notation in what follows, let $s_{t+1} = [(S_{t+1} - S_t)/S_t]$ and $f_t = [(F_t - S_t)/S_t]$.

Because the actual rate of change of the spot rate is equal to its conditional expected value plus a prediction error, we can write

$$s_{t+1} = E_t(s_{t+1}) + \eta_{t+1}, \tag{23}$$

where the prediction error, η_{t+1}, is orthogonal to all variables in the information set at time t. From ⟨22⟩, it follows that

$$C(f_t, s_{t+1}) = C[E_t(s_{t+1}), s_{t+1}] + C(p_t, s_{t+1}) \tag{24}$$

where $C(\cdot, \cdot)$ denotes the covariance operator. Using ⟨23⟩ and the orthogonality of η_{t+1} to all time t information, we find

$$C(f_t, s_{t+1}) = V[E_t(s_{t+1})] + C[p_t, E_t(s_{t+1})] \tag{25}$$

where $V(\cdot)$ denotes the variance operator.

The covariance between the forward premium and the actual rate of change of the spot rate can be measured as the left hand side of ⟨25⟩. Neither of the terms on the right hand side of ⟨25⟩, the variance of the expected rate of depreciation of the domestic currency relative to the foreign currency and the covariance between the risk premium and the expected rate of depreciation, is observable. Since $V[E_t(s_{t+1})]$ is positive, the observable covariance on the left-

hand side of ⟨24⟩ yields an estimated upper bound for $C[p_t, E_t(s_{t+1})]$. Consequently, if $C(f_t, s_{t+1})$ is statistically significantly negative, then $C[p_t, E_t(s_{t+1})]$ is negative and greater in absolute value than $V[E_t(s_{t+1})]$.

As Fama (1984) notes, negative covariation between the risk premium and the expected rate of change of the spot rate sufficient to make the measurable covariance in ⟨25⟩ negative implies that the variance of the risk premium is greater than the variance of the expected rate of depreciation. This is easily demonstrated by examining the variance of the forward premium:

$$V(f_t) = V[E_t(s_{t+1})] + V(p_t) + 2C[p_t, E_t(s_{t+1})]. \tag{26}$$

Fama (1984) uses equations ⟨22⟩ and ⟨23⟩ to note that an ordinary least squares (OLS) regression of s_{t+1} on f_t as in:

$$s_{t+1} = \alpha + \beta f_t + \varepsilon_{t+1} \tag{27}$$

produces an estimated $\hat{\beta}$, that has a probability limit given by:

$$\begin{aligned} \beta &= C(s_{t+1}, f_t)/V(f_t) \\ &= \{V[E_t(s_{t+1})] + C[p_t, E_t(s_{t+1})]\}/V(f_t), \end{aligned} \tag{28}$$

since OLS provides the projection of s_{t+1} onto a constant and f_t. From ⟨23⟩ we know that $\varepsilon_{t+1} = \eta_{t+1} + \mu_t$ where μ_t is the linear prediction error from projecting $E_t(s_{t+1})$ onto a constant and f_t, a subset of the time t information. With nonoverlapping data η_{t+1} is serially uncorrelated, but if μ_t is serially correlated, which is likely, the traditional computation of the OLS standard errors cannot be justified since ε_{t+1} will be serially correlated. Fama (1984) did not find much evidence against the null hypothesis of no serial correlation in the autocorrelations of the residuals of equations like ⟨27⟩ using standard statistical tests, but there is substantial evidence (see, for example, Hodrick and Srivastava, 1984), that the forward premiums of other currencies, which are serially correlated, have statistically significant explanatory power in predicting future spot rates in equations like ⟨27⟩. Thus, it seems that the residuals of ⟨27⟩ may be characterized by weak serial correlation which could bias Fama's standard errors.

At this point there are two questions to be addressed. One is why do the residuals of ⟨27⟩ satisfy standard tests for serial correlation. A second related question is if there is statistically significant serial correlation but standard tests cannot detect it, is there significant bias in the estimation.

In addressing the first question consider the figures reported in Table 1 that are taken from Fama's (1984) Table 1. The standard statistical test for serial correlation compares autocorrelation coefficients to an asymptotic standard error which is $1/\sqrt{T}$ where T is the sample size. With a sample of ten years of monthly data, the standard error is approximately 0.09. Hence, an autocorrelation coefficient must exceed 0.18 in absolute value to be judged statistically significant at traditional levels. Notice that none of the first order autocorrelations in s_{t+1} is greater than 0.18. One interpretation of these statistics could be that s_{t+1} is not serially correlated. An alternative interpretation is that the large forecast errors in s_{t+1} make inference about the degree of serial correlation in the series quite difficult. Notice also that the autocorrelations of the f_t series are highly significantly different from zero.

TABLE 1.

Currency	ρ_1^s	$S(s_{t+1})$	$V(s_{t+1})$	ρ_1^f	$S(f_t)$	$V(f_t)$	$S(s_{t+1})/S(f_t)$
French franc	−0.04	3.01	9.06	0.65	0.44	0.19	6.84
Japanese yen	0.16	3.05	9.30	0.85	0.64	0.40	4.77
Swiss franc	0.01	3.76	14.14	0.86	0.37	0.14	10.16
UK pound	0.15	2.58	6.56	0.87	0.35	0.12	7.37
West German mark	0.01	3.08	9.49	0.78	0.24	0.056	12.83

Note: $S(\cdot)$ denotes the standard deviation, $V(\cdot)$ denotes variance, and ρ_1^s and ρ_1^f are the first order autocorrelation of s_{t+1} and f_t. Data are from Fama (1984).

Since these variables are statistically significant in regressions like ⟨27⟩, we know that there is a statistically meaningful expected rate of change in exchange rates. However, applying the $1/\sqrt{T}$ test to the residuals will not reject the null hypothesis of no serial correlation because the residuals also contain the unanticipated change in the exchange rate, η_{t+1}. Based on the regressions in Hodrick and Srivastava (1984) in which a set of forward premiums was statistically significant in explaining s_{t+1}, it is entirely plausible to conjecture that the residuals in equations like ⟨27⟩ contain approximately the same amount of residual serial correlation as the raw series used as the dependent variables. It is also plausible based on the statistics in Table 1 to consider that the variance of η_{t+1}, the unanticipated change of the exchange rate, is perhaps 50 to 100 times the variance of the projection error, μ_t. Thus, in this case it seems unlikely that the bias is very severe.

II.A. The GMM Alternative

Rather than investigate the severity of the bias directly, which is cumbersome and quite difficult in the case of stochastic regressors that are simply predetermined variables, we investigate three alternative strategies of analysis.

First, we consider estimation of the parameters in ⟨27⟩ as a problem in Hansen's (1982) Generalized Method of Moments (GMM). In this case we treat ⟨27⟩ as a model that supplies us with two orthogonality conditions. Let $z_t' = (1, f_t)$, and define the function $h(s_{t+1}, z_t, b_0) = \varepsilon_{t+1}$ where $b_0' = (\alpha, \beta)$ is the true parameter vector. Then the orthogonality conditions of the model are given by defining the function $f(s_{t+1}, z_t, b)$ as

$$f(s_{t+1}, z_t, b) = h(s_{t+1}, z_t, b) \otimes z_t, \tag{29}$$

where $\otimes$ denotes Kronecker product, and the model implies $E[f(s_{t+1}, z_t, b_0)] = 0$. Since b is unknown and must be estimated, a GMM estimator can be constructed by defining the function $g_0(b) = E[f(s_{t+1}, z_t, b]$ which has a zero at $b = b_0$. The method of moments estimator of the function g_0 for a sample of size T is:

$$g_T(b) = \frac{1}{T} \sum_{t=1}^{T} f(s_{t+1}, z_t, b), \tag{30}$$

and b can be chosen by minimizing the criterion function:

$$\langle 31 \rangle \qquad J_T(b) = g_T(b)' W_T g_T(b)$$

where W_T is an appropriately chosen weighting matrix. Since W_T is (2×2) and the model is linear, the choice of W_T does not affect the estimated values of the parameters. In this case the GMM estimates are the OLS estimates. The important point is that the choice of W_T does affect the standard errors of the parameters, and different auxiliary assumptions lead to different optimal choices of W_T where optimality implies the minimum asymptotic covariance for estimators that impose the same orthogonality conditions. The OLS standard errors are produced by the auxiliary assumptions of no serial correlation, $E(\varepsilon_{t+1} | z_t, \varepsilon_t, z_{t-1}, \varepsilon_{t-1}, \ldots) = 0$, and conditional homoscedasticity, $E(\varepsilon_{t+1}\varepsilon'_{t+1} | z_t, \varepsilon_t, z_{t-1}, \varepsilon_{t-1}, \ldots) = \sigma_\varepsilon^2$. These are the assumptions of case (i) of Hansen (1982), p. 1043). We argued previously that the ε_{t+1} series is arbitrarily serially correlated under the null hypothesis, which makes these assumptions inappropriate. More appropriate assumptions are case (v) of Hansen (p. 1045). We assume that $E(\varepsilon_{t+1})$ and $E(z_t)$ are zero and that the processes are linearly regular with fourth order cumulants that are zero. In this case the asymptotic covariance matrix of the GMM parameters for the optimal choice of W_T is $\Omega = (D_0' S_w^{-1} D_0)^{-1}$ where $w_t \equiv f(s_{t+1}, z_t, b_0)$, $S_w \equiv \sum_{j=-\infty}^{\infty} W_w(j)$, $R_w(j) \equiv E(w_t w'_{t-j})$ and $D_0 \equiv E(z_t z_t')$. In this case the optimal W_T is a consistent estimate of S_w^{-1} which can be obtained from estimation of the spectral density of the w_t process using the residuals from OLS estimation. The estimated covariance matrix also employs a consistent estimate of D_0 which is just the moment matrix of the regressors,

$$D_T = \frac{1}{T} \sum_{t=1}^{T} z_t z_t'.$$

The data used in this paper consist of spot and one-month forward exchange rates between the US dollar and five currencies, the French franc, the Japanese yen, the Swiss franc, the British pound and the West German mark. The estimation used 119 nonoverlapping observations, the sample period being July 1973 to September 1983.[5]

The results of this estimation are presented in Table 2. Not surprisingly, the results of the test are supportive of Fama's findings. All the currencies have negative $\hat{\beta}$'s, and those of the Japanese yen, the Swiss franc, and the UK pound are statistically significantly negative at standard marginal levels of significance. Rather than extend the analysis in the direction of seemingly unrelated regressions as was done by Fama, we explore two alternative ways of conducting the investigation. These tests are described in the next subsection of the paper.

II.B. Two Additional Complementary Techniques

Our first additional test allows us to examine whether the covariance in ⟨24⟩ is negative without assuming lack of serial correlation. The desire to develop a multivariate analogue to ⟨27⟩ that does not impose the assumption of no serial correlation in the residuals provides the motivation for our second additional test. Unfortunately, the coefficient of the own forward premium does not

TABLE 2.

GMM Estimates of $s_{t+1}=\alpha+\beta f_t+\varepsilon_{t+1}$		
Currency	$\hat{\alpha}$ (Std. Err.)	$\hat{\beta}$ (Std. Err.)
French franc	−0.25	−0.44
	(0.44)	(0.97)
Japanese yen	0.28	−1.22
	(0.35)	(0.51)
Swiss franc	0.94	−2.06
	(0.41)	(0.67)
UK pound	−0.58	−2.13
	(0.39)	(0.93)
West German mark	0.20	−0.89
	(0.37)	(0.94)

Note: Parameter estimates are OLS. Standard errors are calculated as in case (v) of Hansen (1982). Data are described in note 5.

decompose as in ⟨28⟩ if more regressors are added to ⟨27⟩. It is, nevertheless, still possible to test the hypothesis of negative covariation in the multivariate case. The procedure consists of predicting s_{t+1} using a multivariate analogue of ⟨27⟩ to compute an estimate of $V[E_t(s_{t+1})]$ which can be compared with $V(f_t)$. From ⟨26⟩, if $V[E_t(s_{t+1})] > V(f_t)$, the covariance of p_t and $E_t(s_{t+1})$ must be negative.

We turn next to a formal description of our alternative statistical tests. Note that the first case examines whether a covariance of two arbitrarily serially correlated time series is negative while the second case tests whether the difference of two variances is negative. Hypothesis testing in these cases requires the distribution of sample variances and covariances, and these are derived and estimated by using some elementary spectral analysis. The procedures followed here are similar to those employed by Meese and Singleton (1980) and Singleton (1980).

In the two tests we assume that the rate of change of the spot rate, s_{t+1}, and the forward premium, f_t, are time series that are stationary up to their fourth moments. For the purposes of our first test, they can be arbitrarily serially correlated. Define the autocovariance function of the vector $v_t'=[s_{t+1}, f_t]$ to be the matrices $R_v(v) \equiv E\{[v_t - E(v_t)][v_{t-v} - E(v_{t-v})]'\}$, $v = \ldots -1, 0, 1, \ldots$. Then, it is well known (Fishman 1969, pp. 61–64) that the spectral density function of v_t, $S_v(\lambda)$, is the Fourier transform of $R_v(v)$,

$$S_v(\lambda) = (1/2\pi) \sum_{v=-\infty}^{\infty} R_v(v)\exp(-i\lambda v), \qquad |\lambda| \leqslant \pi, \tag{32}$$

where $i=\sqrt{-1}$. Let $S_f(\lambda)$ and $S_s(\lambda)$ denote the spectral densities of the f_t and s_{t+1} series, and let $S_{fs}(\lambda)$ be the cross-spectral density. Since $R_{fs}(v)=R_{sf}(-v)$, $S_{sf}(\lambda)=S_{fs}(-\lambda)$.

Additional Test 1

Our first additional test uses the sample covariance, $C_{fs}(0)$, to test the hypothesis that $R_{fs}(0) < 0$.[6] When the sample covariance is computed in the usual way from a sample of size T as

$$\langle 33 \rangle \qquad C_{fs}(0) = T^{-1} \sum_{t=1}^{T} (f_t - \bar{f})(s_{t+1} - \bar{s})$$

where $\bar{f} = T^{-1} \sum_{t=1}^{T} f_t$ and $\bar{s} = T^{-1} \sum_{t=1}^{T} s_{t+1}$ are the sample means, it is easily demonstrated by substituting for $\bar{f}$ and $\bar{s}$ in ⟨33⟩, that

$$\langle 34 \rangle \qquad E[C_{fs}(0) - R_{fs}(0)] = -T^{-1} \left[\sum_{v=-(T-1)}^{T-1} \left(1 - \frac{|v|}{T}\right) R_{fs}(v) \right].$$

Hence, $C_{fs}(0)$ is a biased estimator of $R_{fs}(0)$, but from ⟨32⟩, the bias is proportional to $-2\pi S_{fs}(0) T^{-1}$ and is unimportant in large samples. Therefore, $C_{fs}(0)$ is a consistent estimator of $R_{fs}(0)$.

In order to conduct inference, we need the asymptotic distribution of the sample covariance, $C_{fs}(0)$. Fishman (1969, p. 121) demonstrates that $\sqrt{T}(C_{fs}(0) - R_{fs}(0))$ is asymptotically normally distributed with zero mean and variance

$$\langle 35 \rangle \qquad V[C_{fs}(0)] = 2\pi \int_{-\pi}^{\pi} [S_f(\lambda) S_s(\lambda) + S_{fs}^2(\lambda)]\, d\lambda$$

when f_t and s_{t+1} are normal. Valid inference about the hypothesis that the true covariance is negative can be conducted as a one-sided test of the null hypothesis $R_{fs}(0) \geqslant 0$ using the normal distribution described above.

The series were first subjected to Fourier transformation, and the periodogram was computed at 120 equally spaced frequencies. The spectral densities were then obtained by smoothing the periodogram ordinates using a Daniell window of width 7. This smoothing procedure was deemed satisfactory in that the estimates of the covariances obtained by integrating (summing) the estimated spectral densities were quite close to those computed directly. The results of our one-sided test are presented in Table 3. As can be seen, the sample covariance between the forward premium and the actual rate of change of the spot rate is negative for all five currencies, significantly so for the Japanese yen, the Swiss franc, and the UK pound. While not an identical test to our GMM estimation, this first alternative produces basically the same statistical results.

Additional Test 2

As indicated previously, the second test relies on a comparison of the variance of the forward premium with the variance of the estimated rate of change of the spot rate, $E_t(s_{t+1})$. If

$$\langle 36 \rangle \qquad V(f_t) - V[E_t(s_{t+1})] < 0,$$

then the covariance of the risk premium and the expected rate of depreciation must be negative. The statistical analysis in this case follows Singleton (1980) quite closely. The first step in the analysis is to form an estimate,

TABLE 3.

Currency	$C_{fs}(0)$	$\{V[C_{fs}(0)]/T\}^{1/2}$	ζ-value
French franc	−20.07	23.81	−0.88
Japanese yen	−56.72	29.77	−1.99
Swiss franc	−39.96	23.38	−1.82
UK pound	−34.70	18.73	−1.94
West German mark	−10.68	15.58	−0.71

Note: $C_{fs}(0)$ is given in ⟨33⟩, $V[C_{fs}(0)]$ is given in ⟨35⟩, and $\zeta = C_{fs}(0)/\{V[C_{fs}(0)]/T\}^{1/2}$. For a one-sided test of $C_{fs}(0) < 0$, the null hypothesis $\zeta \geqslant 0$ is rejected at the 95 per cent confidence level if $\zeta \leqslant -1.64$.

$\hat{s}_{t+1} = E(s^i_{t+1} | w_t)$, of the expected rate of change of the exchange rate for each of the five currencies conditional on an information set w_t. We choose $w'_t = [s'_t, f'_t]$, where $s'_t = (s^1_t \ldots s^5_t)$ and $f^1_t \ldots f^5_t)$. Let θ denote the (10×5) matrix of estimated coefficients corresponding to the regressions of $[s^1_{t+1} \ldots s^5_{t+1}]$ onto w'_t. It is well known (Fishman, 1969, p. 71) that the spectral density function of the vector $\hat{s}'_{t+1} = (\hat{s}^1_{t+1} \ldots \hat{s}^5_{t+1})$ is

$$S_{\hat{s}}(\lambda) = \theta' S_w(\lambda)\theta \tag{37}$$

where $S_w(\lambda)$ is the spectral density function of w_t.[7] Following arguments in Hannan (1970), Anderson (1971), and Singleton (1980), it can be shown that the sample variances of $\hat{s}_{t+1}$ and f_t have asymptotically normal distributions with means given by $\sigma^2_{\hat{s}}$ and σ^2_f, the true variances, and covariance matrix of the sample variances given by:

$$V\begin{vmatrix} \hat{\sigma}^2_{\hat{s}} \\ \hat{\sigma}^2_f \end{vmatrix} = \frac{4\pi}{T}\begin{vmatrix} \delta^2_{\hat{s}} & \delta_{\hat{s}f} \\ \delta_{f\hat{s}} & \delta^2_f \end{vmatrix} \tag{38}$$

where

$$\delta^2_{\hat{s}} = \int_{-\pi}^{\pi} S^2_{\hat{s}}(\lambda)\, d\lambda,\ \delta^2_f = \int_{-\pi}^{\pi} S^2_f(\lambda)\, d\lambda \quad \text{and} \quad \delta_{f\hat{s}} = \int_{-\pi}^{\pi} |S_{f\hat{s}}(\lambda)|^2\, d\lambda.^8$$

Here $S_f(\lambda)$ is the spectral density function of f and $S_{f\hat{s}}(\lambda)$ is the cross-spectral density function of f and $\hat{s}$. From these distributions we find that $(\hat{\sigma}^2_f - \hat{\sigma}^2_{\hat{s}})$ is asymptotically normal with mean $\sigma^2_f - \sigma^2_{\hat{s}}$ and variance $\delta^2 = (\delta^2_f + \delta^2_{\hat{s}} - 2\delta_{f\hat{s}})4\pi/T$. Examination of whether the variance of the expected rate of depreciation is greater than the variance of the forward premium can be done as a one-sided test of the null hypothesis $\sigma^2_f - \sigma^2_{\hat{s}} \geqslant 0$ based on this asymptotic distribution.

In performing this test, we used the same data and spectral estimation method as in the first test. The results of the test are presented in Table 4. As can be seen, the tests here based on the large magnitude of the ζ statistics indicate very strongly that the covariance between the risk premium and the expected rate of change on the spot rate is negative for all five currencies in our sample.

III. Conclusions

The primary purpose of this paper has been to investigate the covariation of the risk premium in the forward foreign exchange market and the expected rate of

TABLE 4.

Currency	$V(f)-V(\hat{s})$	δ	z-value
French franc	−98.82	21.16	−4.67
Japanese yen	−155.28	36.54	−4.25
Swiss franc	−193.81	36.08	−5.37
UK pound	−191.10	40.59	−4.71
West German mark	−95.70	24.19	−3.96

Note: $V(\hat{s})$ and $V(f)$ are the computed sample variances of $\hat{s}$ and f, δ^2 is specified in the text following ⟨38⟩, $z=[V(\hat{s})-V(f)]/\delta$, $z<-1.64$ implies that z is negative with 95 per cent confidence.

depreciation of the US dollar relative to five other currencies. Using alternative statistic techniques, we confirmed the findings reported in Fama (1984). If one views the forward premium as the sum of the expected rate of depreciation of the currency plus a risk premium, then our evidence indicates that the risk premium is negatively correlated with the expected rate of depreciation. The risk premium, when defined this way, is the expected return to purchasing dollars in the forward market. The expected return to selling dollars or buying foreign currency forward therefore covaries positively with the expected rate of depreciation of the dollar relative to all five foreign currencies. Although Fama (1984) found such a covariation puzzling and potentially inconsistent with economic theory, we have demonstrated that it is intuitively plausible and consistent with the prediction of the Lucas (1982) model.

The magnitude of the covariance also indicated that the variance of the risk premium is greater than the variance of the expected rate of depreciation. Since rational expectations models of spot exchange rate determination have focused almost exclusively on the latter term, this quantitative finding suggests that more work ought to be devoted to determining how risk affects the determination of spot exchange rates.

Of course, this analysis as well as Fama's and all modern rational expectations time series analysis relies on the statistical assumptions of stationarity and ergodicity. Krasker (1980) has argued that these assumptions may be incorrect in such analyses. Agents may care about events that have not occurred in the sample, and the probability of these events may fluctuate.[9] Developing estimation methods to handle these problems may not be as critical as determining what the factors actually are. Fama's Section 5 offers some alternative interpretations of the data that certainly demand some consideration. First, it is, of course, possible that the market is inefficient, although this hypothesis receives virtually no support in studies of other financial markets. A second alternative is attributed to Richard Roll who apparently suggested that a government may obstinately force appreciation of a currency precisely in those periods during which market forces are predicting depreciation. If the unbiasedness hypothesis were true, such a finding could only be consistent with the data if governments could consistently fool the public. Hence, we also give little credence to this explanation in a rational world. Fama's third possible explanation was offered by Michael Mussa. He argued that these markets may be characterized by periods of brief skewness in

the distribution of future exchange rates due to uncertainty in the direction of government policy. Without sufficient draws from these changing distributions, it is possible that *ex post* changes in exchange rates are consistently below their *ex ante* means which creates negative small sample covariation between the forward premium under the unbiasedness hypothesis and the actual rate of depreciation.[10] Tackling these problems of heteroscedasticity and changing skewness in the distributions of returns requires larger samples and explicit incorporation of additional theory regarding the linkage between the determination of exchange rates and government policies. This is certainly a fertile area for additional theoretical and empirical analysis.

Notes

1. See Hodrick and Srivastava (1984) for tests of a particular model and for extensive references to the empirical and theoretical literature on the subject. See also Frankel (1982), Domowitz and Hakkio (1985), Korajczyk (1985), and Mark (1985) for tests of alternative specifications of risk premiums.
2. This specification of the Lucas model has been criticized because the determination of the exchange rate lacks a forward looking component. The problem arises because the timing of trades in securities and goods makes the demand for money insensitive to the nominal interest rate. Alternative ways of overcoming this problem are discussed in Lucas and Stokey (1983) and Svensson (1983).
3. Domowitz and Hakkio (1985) employ these distributional assumptions in their analysis.
4. Fama (1984) specified his analysis in natural logarithms. Since $\ln(S_{t+1})-\ln(S_t)$ and $\ln(F_t)-\ln(S_t)$ are very nearly equal to $[(S_{t+1}-S_t)/S_t]$ and $[(F_t-S_t)/S_t]$ and have almost perfect correlation, it is highly unlikely that statistically significant differences would separate the results.
5. The data were supplied by Data Resources, Inc. They are a nonoverlapping sample beginning in July 1973 and ending in September 1983 in which Tuesday forward rates predict Thursday spot rates 30 days in the future with the next observation being the following Friday forward rate predicting a corresponding Monday spot rate. Hakkio (1983) argues correctly that the future spot rates do not match the precise value day that is specified by a forward contract. Fama (1984) used the Harris Bank data treating a Friday forward rate as predicting a Friday spot rate four weeks in the future which is also incorrect. Thus, our data contain slight measurement errors. Riehl and Rodriguez (1977) discuss the rules that regulate the execution of a forward contract. Meese and Singleton (1980), Hsieh (1984), and Cumby and Obstfeld (1984) match the data more precisely than here, and they find very little difference in inference regarding evidence against the unbiasedness hypothesis.
6. Because the distribution theory for a sample covariance requires consideration of the entire autocorrelation function of the vector time series, we have introduced new notation. Clearly, $R_{fs}(0)=C(f_t, s_{t+1})$ is the covariance of interest.
7. Dhrymes (1974, p. 532) notes that $S_i(\lambda)$ is a consistent estimator of the true spectral density whenever a consistent estimator of θ is used in $\langle 37 \rangle$.
8. Anderson (1971, p. 593) establishes conditions under which the sample variances have limiting normal distributions despite the fact that θ is estimated. Our analysis is valid given these regularity conditions.
9. Tests of excess volatility of stock prices relative to dividends as in Shiller (1981) are subject to the same critique although the problem is much broader than just these financial studies. All rational expectations estimation techniques require that large sample moments correspond to the moments of the true distributions. If this is not the case because drastic events have not occurred with sufficient frequency, then the studies make an error.
10. It is interesting to note that evidence of negative covariation can be found in Bilson (1981) who broke the forward premium into small and larger values on the basis of whether they

were smaller or larger than 10 per cent at an annual rate. His seemingly unrelated regression

$$\hat{s}_{t+1} = \underset{(0.15)}{0.26} f_t^S - \underset{(0.12)}{0.28} f_t^L$$

indicates that the negative β coefficient in equations like ⟨27⟩ may be due to extreme values in the data which supports Mussa's conjecture since relatively few of these large values occurred during the sample.

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