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PESO PROBLEMS, BUBBLES, AND RISK IN THE EMPIRICAL ASSESSMENT OF EXCHANGE-RATE BEHAVIOR

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Peso Problems, Bubbles, and Risk in the Empirical Assessment of Exchange-Rate Behavior

ABSTRACT

One of the most puzzling aspects of the post-1973 floating exchange rate system has been the apparently inefficient predictive performance of forward exchange rates. This paper explores some aspects of each of three leading explanations of forward-rate behavior. The paper first develops a simple rational-expectations model of the "peso problem" that generates some key empirical regularities of the foreign exchange market: seemingly predictable and conditionally heteroskedastic forward forecast errors, along with possible directional misprediction by the forward premium. The implications of bubbles for tests of forward-rate predictive efficiency are discussed next. It is argued that the existence of bubbles is extremely difficult (if not impossible) to establish empirically. Even though some types of bubble could distort standard tests on the relation between spot and forward exchange rates, it seems unlikely that these bubbles have been an important factor. Finally, the paper examines foreign-exchange asset pricing under risk aversion and suggests that a convincing account of forward-rate behavior should also help explain the results found in testing other asset-pricing theories, such as the expectations theory of the interest-rate term structure.

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One of the most puzzling aspects of the post-1973 floating exchange rate system has been the apparently inefficient predictive performance of forward exchange rates. Richard Meese's survey of research on international asset pricing ably documents both the evidence on forward rate determination and the attempts of researchers to pin down the empirical relationship between forward and spot exchange rates.

Table 1, which reproduces in part results I reported in a 1985 paper, is representative of the empirical record. Shown in the table are regressions linking the one-month percentage change in the spot dollar price of a foreign currency to the corresponding one-month nominal interest rate differential (or forward premium) from the previous month. (The sample period is February 1975-January 1985.) For the four foreign currencies tested, the forward premium tended on average to <u>mispredict</u> the direction in which the dollar exchange rate moved during the subsequent month. Forward premiums have therefore been badly biased predictors of subsequent exchange-rate movements. As Meese notes, a further indication of the poor forecasting performance of forward rates is the large body of evidence pointing to the ex ante predictability of forward forecast errors by past information.[1]

Three broad approaches to explaining the preceding facts can be

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[1] The table also shows that forward premiums have virtually no predictive value with respect to future depreciation. This fact is not necessarily evidence of some market failure; it is consistent with a world in which exchange rate movements are dominated by rational market reactions to unpredictable events. It is noteworthy that the early regressions of spot exchange rate <u>levels</u> on lagged forward rate levels seemed to indicate a greater predictive value of the forward rate than is evident in Table 1. Meese's argument that the forward- and spot-rate processes are cointegrated suggests that the latter findings are the spurious result of a common stochastic trend.

Table 1: One-Month Eurocurrency Interest-Rate Differentials and Subsequent Dollar Depreciation, February 1975-January 1985

Depreciation rate of dollar against:	Constant	Lagged Eurocurrency interest differential	D.-W.	$\bar{R}^2$
Yen	0.113 (0.056)	-2.188 (0.985)	1.97	0.03
French Franc	-0.092 (0.039)	-0.483 (0.773)	2.27	-0.01
Deutschemark	0.040 (0.071)	-1.778 (1.573)	2.23	0.00
Pound Sterling	-0.102 (0.037)	-1.477 (0.875)	1.93	0.02

Note: The dependent variable is the percentage change in the dollar price of the foreign currency, expressed on an annualized basis. The independent variables are a constant and the difference between the previous month's one-month Eurodollar deposit rate and one-month foreign Eurocurrency deposit rate. Exchange rate data are end-of-month data from OECD *Main Economic Indicators*. End-of-month interest rates come from Morgan Guaranty Trust Company of New York, *World Financial Markets*. Standard errors are given in parentheses below the coefficient estimates.

discerned in Meese's review. The first of these approaches holds that the empirical findings may result from distributional peculiarities such as the "peso problem" or from other factors that cause nonstationarities in the data. The second approach posits an important role for speculative bubbles and/or market irrationality in exchange-rate determination. The third approach observes that private risk aversion can imply time-varying discrepancies between forward premiums and expected future spot depreciation. Risk aversion thus offers another potential explanation of measured deviations from uncovered interest parity.

Economists have not reached a consensus interpretation of the data. Yet, the answer may have important implications for policymakers. A determination that the foreign-exchange market is driven by irrational trading, for example, would stengthen both the case for fixed exchange rates and the case for direct controls on international capital movements.

Below I explore some aspects of each of the three leading explanations of forward-rate behavior. Section I develops a simple rational-expectations model of the peso problem that generates some key empirical regularities of the foreign exchange market: seemingly predictable and conditionally heteroskedastic forward forecast errors, along with possible directional misprediction by the forward premium. Section II discusses the implications of bubbles for tests of forward-rate predictive efficiency. Section III expands on Meese's discussion of foreign-exchange asset pricing under risk aversion. I suggest there that a convincing account of forward-rate behavior should also help explain the results found in testing other asset-pricing theories, such as the expectations theory of the interest-rate term structure.

I. Peso Problems

A peso problem arises when market forecasts reflect the possibility of major events that occur relatively infrequently in the data set available to the econometrician. Even though market expectations may be entirely reasonable ex ante, market forecasts appear biased and forecast errors appear serially correlated in the ex post sample. To the best of my knowledge, the effects of the peso problem on efficiency tests were first noticed by Kennth Rogoff in connection with the behavior of Mexican peso futures prior to the August 1976 devaluation of that currency. (An elegant formal model of the episode has been developed by José Saúl Lizondo.) Stephen Salant and Dale Henderson presented the first theoretical analysis of a peso-type problem in studying the behavior of gold prices under the threat of official gold-market intervention.

The peso problem is essentially one of small-sample inference. An announcement by the United States Treasury Secretary that the dollar is "too strong" may lead the market to repeatedly underestimate the dollar's future strength if monetary policies are not promptly adjusted to push the dollar down. Market expectations will be reassessed if it becomes clear that official pronouncements are not being backed by concrete policy shifts. But an econometrician looking at a limited data set would erroneously interpret hypothesis tests based on asymptotic distribution theory as rejections of market inefficiency. Reliable inference would require a data set covering many similar episodes, including some in which official pronouncements did lead to the implied policy actions.

An example of a floating-rate peso problem shows how some aspects of the seemingly anomalous behavior of exchange rates can be modelled without appeal to irrationality or risk aversion. Such a model requires the description of a general-equilibrium framework within which exchange rates are determined. To make my main points as simply as possible, I adopt the monetary approach to exchange-rate determination, as described by Michael Mussa. Karen Lewis has used this model to construct a complementary example in which a peso problem arises because agents are uncertain about the monetary regime currently in effect.

The monetary approach is based on an exchange rate equation of the form

$$(1)\quad s(t) = \lambda\{E_t[s(t+1) - s(t)]\} + \alpha'\phi(t),$$

where $s(t)$ is the natural logarithm of $S(t)$, the domestic-currency price of foreign exchange, λ is a constant that measures the spot rate's response to expected future depreciation, $E_t(.)$ is a rational expectation conditional on time-t information, α is a coefficient vector, and $\phi(t)$ is a vector of exogenous variables. The equation is derived from an uncovered interest-rate parity condition together with a theory of nominal interest rates and price levels in which the key variables are $s(t)$ and $\phi(t)$. The vector $\phi(t)$ could in principle include an exogenously-varying risk premium, but such an exogeneity assumption is implausible. The model embodied in (1) is therefore best viewed as describing a hypothetical world economy with risk-neutral investors.

The "bubble-free" solution to (1) is

$$(2)\quad s(t) = [1/(1+\lambda)] \sum_{i=0}^{\infty} [\lambda/(1+\lambda)]^i E_t[\alpha'\phi(t+i)].$$

This solution is found by iterating (1) forward and imposing a stationarity condition that is described in detail in the next section. For my example, it is sufficient to assume that $\phi(t)$ consists of a single variable (the log of the money supply) and that $\alpha = 1$.

To model a peso problem, assume that on any date t, the log money supply is the sum of two random variables. The first is a random-walk process, $m(t)$, the sum of $m(t-1)$ and a completely unpredictable mean-zero disturbance, $\omega(t)$. The second random variable contributing to the money-supply process is itself the product of two random variables, denoted $d(t)$ and $\mu(t)$. $d(t)$ has the distribution:

$$d(t) = \begin{cases} 1 \text{ (with probability } \pi) \\ 0 \text{ (with probability } 1 - \pi) \end{cases} \quad \text{conditional on } d(t-1) = 0;$$

$$d(t) = 1 \text{ (with probability 1) conditional on } d(t-1) = 1.$$

Conditional on $d(t-1) = 0$, $\mu(t)$ is generated by the first-order autoregressive process $\mu(t) = \rho\mu(t-1) + \varepsilon(t)$, where $0 \leq \rho < 1$ and $E_{t-1}[\varepsilon(t)] = 0$; but $\mu(t+j)$ $(j > 0)$ is constant at $\mu(t)$ conditional on $d(t) = 1$. $\phi(t)$ therefore follows the process

$$(3)\quad \phi(t) = m(t) + d(t)\mu(t) = m(t-1) + \omega(t) + d(t)\mu(t).$$

The interpretation of equation (3) is this. One component of the money supply process is a random walk, which in itself would lead to no expected future change in money. It is the second component, the product $d(t)\mu(t)$, which causes a peso problem. Each period, there is a probability π that the money supply will jump permanently by the amount $\mu(t)$, given that no such jump has yet taken place. The expected value of this possible jump is $\rho\mu(t-1)$ in the previous period. The $\mu(t)$ process is assumed to damp out monotonically in mean to capture the intuitive notion that the market expects to put progressively less weight on the significance of a possible policy change the longer the interval over which the change has not occurred.[2]

The simple exchange-rate model set out above can now be used to analyze the difficulties associated with peso problems. For this purpose, it is assumed that $d(t) = 0$ for all observations t available to the econometrician. In other words, even though the market viewed a major policy switch as a possibility, the available data are conditioned on the non-occurrence of that event. The peso problem modelled here has three major implications:

1. Apparent conditional bias in market exchange-rate forecasts. Assume for simplicity that the econometrician can directly observe the market's expectation, $E_t s(t+1)$, of next period's exchange rate, but not the true, ex ante money-supply process itself. In forecasting the exchange rate, the market took into account the possibility that d might equal 1 on some future date, even though the event $d = 1$ turned out not to occur ex post. The possibility that d might equal 1 is reflected in the sequence of expected future money supplies, $E_t\phi(t+i)$, $i > 0$.

That sequence is computed as follows. If $d(t) = 0$, the event $d = 1$ will have occurred by date t+i with probability $[1 - (1-\pi)^i] = \pi + \pi(1-\pi) + \pi(1-\pi)^2 + \ldots + \pi(1 - \pi)^{i-1}$. On date t, the expected value of $\phi(t+i)$ is therefore the sum of $m(t)$ and $\{\pi + \pi(1-\pi)\rho + \pi[(1-\pi)\rho]^2 + \ldots + \pi[(1-\pi)\rho]^{i-1}\}\rho\mu(t)$, where the latter term is just the expected date-(t+i) value of the second component of the money-supply process. Thus, conditional on $d(t) = 0$,

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[2] Alternatively, one could assume a time-varying π to capture this type of effect.

$$E_t\phi(t+i) = m(t) + \frac{1 - [(1-\pi)\rho]^i}{1 - (1-\pi)\rho}\pi\rho\mu(t)$$

for all $i > 0$.

Combined with equation (2) (for the case $\alpha = 1$), the above formula implies

$$(5) \quad s(t) = m(t) + \frac{\pi\lambda\rho}{[1 + \lambda - (1-\pi)\lambda\rho]}\mu(t).$$

A similar calculation shows that the rational one-period-ahead exchange-rate forecast is

$$(6) \quad E_t[s(t+1)] = m(t) + \frac{\pi(1 + \lambda)\rho}{[1 + \lambda - (1-\pi)\lambda\rho]}\mu(t).$$

Together, (5) and (6) yield a date-(t+1) forecast error of

$$(7) \quad s(t+1) - E_t[s(t+1)] = \omega(t+1) + \frac{\pi\lambda\rho\varepsilon(t+1) - [\rho + (1-\rho)\lambda\rho]\pi\mu(t)}{[1 + \lambda - (1-\pi)\lambda\rho]},$$

given that $d(t+1) = 0$.

Note that the forecast error conditional on $d(t+1) = 0$ depends on the lagged variable $\mu(t)$. If $\mu(t) > 0$ over some interval, for example, the market appears systematically to overestimate the home currency's future depreciation. This pattern of ex post forecast errors does not imply a conditionally biased forecast: the market also took into account the likelihood of the event $d(t+1) = 1$, which would have occasioned a sharp increase in the money supply.

More formally, if $d(t+1) = 1$ had occurred, the equilibrium exchange rate, $s(t+1) = m(t+1) + \mu(t+1)$, would have resulted in a forecast error of

$$(8)\quad s(t+1) - E_t[s(t+1)] = \omega(t+1) + \varepsilon(t+1) + \frac{[\rho + (1-\rho)\lambda\rho](1-\pi)\mu(t)}{[1 + \lambda - (1-\pi)\lambda\rho]}.$$

The conditional expectation of the forecast error with respect to time-t information only is

$$\pi E_t[\text{error}(t+1)|d(t+1) = 1] + (1-\pi)E_t[\text{error}(t+1)|d(t+1) = 0] =$$

$$E_t\{\omega(t+1) + \frac{\pi(1 + \lambda)\varepsilon(t+1)}{[1 + \lambda - (1-\pi)\lambda\rho]}\} = 0.$$

Expectations are therefore conditionally unbiased.

An econometrician viewing a sample conditioned on $d(t) = 0$ would fail to reach this last conclusion using standard methods of inference. If the forecast error (7) is regressed on the lagged depreciation forecast, for example, a negative regression coefficient will be found with high probability (as implied by results in the next paragraph). But such a finding is not evidence of an inefficient market in the present context. The resulting regression equation provides an estimate of $E_t[\text{error}(t+1)|d(t+1)=0]$, which is generally nonzero. Since $E_t[\text{error}(t+1)] = 0$, however, forecast errors are indeed uncorrelated with past information given the true distribution of the money supply (and hence of forecast errors).

2. Exchange-rate misprediction. This particular model implies also that rational market forecasts will on average mispredict the exchange rate's direction of change as long as $d = 1$ does not occur. By (5) and (6), the market forecast of currency depreciation, given that $d(t) = 0$, is

$$E_t[s(t+1)] - s(t) = \frac{\pi\rho}{[1 + \lambda - (1-\pi)\lambda\rho]}\mu(t).$$

If $\mu(t) > 0$, for example, the market expects the currency to depreciate, and the home interest rate consequently exceeds the foreign rate under uncovered interest parity. The actual change in the exchange rate, however, is

$$s(t+1) - s(t) = \omega(t+1) + \frac{\pi\lambda\rho\varepsilon(t+1) - (1-\rho)\pi\lambda\rho\mu(t)}{[1 + \lambda - (1-\pi)\lambda\rho]}.$$

If $\omega(t+1) = \varepsilon(t+1) = 0$, the currency <u>appreciates</u> (still assuming $\mu(t) > 0$) as the market revises downward by the factor $1-\rho$ the expected future path of the money supply's second component. In a sample where $d(t) = 0$ for all t, realized and expected depreciation will therefore be negatively correlated, even though interest parity holds and expectations are rational. (Table 1 showed that recent data are characterized by a negative correlation between forward premiums and subsequent depreciation rates.)

3. <u>Conditional heteroskedasticity of forecast errors</u>. Peso problems can result in time-dependence in the conditional covariances of forecast errors. Robert Cumby and I first presented evidence of such conditional heteroskedasticity in forward-rate forecast errors, and our finding has been confirmed by Robert Hodrick and Sanjay Srivastava, Alberto Giovannini and Philippe Jorion, and many others. As Meese notes, conditional heteroskedasticity may pose problems for statistical inference. In a world of risk-averse investors, the phenomenon also has a crucial implication about the nature of foreign-exchange risk premiums, as detailed below in section III.

Assume that the random variables $\omega(t)$ and $\varepsilon(t)$ themselves have constant conditional variances, so that $E_t[\omega(t+1)^2] = \sigma_\omega^2$ and $E_t[\varepsilon(t+1)^2] = \sigma_\varepsilon^2$. In spite of these assumptions, the conditional variance of the exchange-rate forecast error depends on lagged information in the preceding model. If d = 1 has occurred, then the exchange-rate forecast error is just $\omega(t)$. Thus, $E_t[\text{error}(t+1)^2] = \sigma_\omega^2$, where the conditioning set contains d(t). If d = 1 has not occurred, however, the conditional variance of the forecast error is different. Assume $\rho = 0$ for simplicity. Under that assumption, the population conditional variance of the forecast error, $E_t[\text{error}(t+1)^2]$, is $\pi(\sigma_\omega^2 + 2\sigma_{\omega\varepsilon} + \sigma_\varepsilon^2) + (1-\pi)(\sigma_\omega^2) = \sigma_\omega^2 + \pi(2\sigma_{\omega\varepsilon} + \sigma_\varepsilon^2)$, where the conditioning set again contains d(t).

An econometrician may well find evidence of conditional heteroskedasticity in a sample conditioned on d(t) = 0. For example, regressions of the squared forecast error (7) on variables including the squared forward premium, $\pi^2\rho^2\mu(t)^2/[1 + \lambda - (1-\pi)\lambda\rho]^2$, are likely to be significant. Cumby and I used this type of regression specification to test for conditional heteroskedasticity.

Other statistical difficulties may distort standard efficiency tests even when peso problems are absent. One of these, discussed extensively by Meese, is the unit-root problem, which invalidates much of the asymptotic distribution theory on which many of the tests in the literature depend. Another instance of nonstationarity, also mentioned by Meese, is probably very important in practice. Because the processes generating exchange-market data are subject to structural change, there is no guarantee that the unconditional covariances of the relevant economic variables remain constant through time. The possibility of such structural changes has implications for econometric attempts to identify another anomalous exchange-rate phenomenon, the speculative bubble.

II. Speculative Bubbles

Speculative bubbles can be modelled within the monetary model of the previous section. Equation (2), which gives the equilibrium exchange rate just analyzed, is only one possible solution to the stochastic difference equation (1). A general solution takes the form

$$(9) \quad s(t) = [1/(1+\lambda)] \sum_{i=0}^{\infty} [\lambda/(1+\lambda)]^i E_t[\alpha'\phi(t+i)] + x(t)[(1+\lambda)/\lambda]^t,$$

where $\{x(t)\}$ is any stochastic process with the martingale property,

$$E_t[x(t+1)] = x(t).$$

Solution (2) corresponds to the choice $x(t) = 0$, for all t, but because $x(t)$ is multiplied in (9) by the exploding term $[(1+\lambda)/\lambda]^t$, choice of a nonzero x-process will impart explosive behavior to the exchange rate. Such explosive paths capture the idea of speculative bubbles driven by self-fulfilling anticipations.

It is sometimes asserted that bubbles give rise to serially-correlated market forecast errors or to biased forecasts, but this is not the case when the bubble's path is described by an equation like (9). Define $v^{t+i}(t+1) = E_{t+1}[\alpha'\phi(t+i)] - E_t[\alpha'\phi(t+i)]$ and $\eta(t+1) = x(t+1) - E_t[x(t+1)] = \Delta x(t+1)$. When expectations are rational and the model (9) is used to forecast future exchange rates, the one-step ahead forecast error is

$$s(t+1) - E_t[s(t+1)] = [1/(1+\lambda)] \sum_{i=0}^{\infty} [\lambda/(1+\lambda)]^i v^{t+i+1}(t+1) + \eta(t+1)[(1+\lambda)/\lambda]^{t+1}.$$

This error is, however, uncorrelated with time-t information (including lagged forecast errors), and the market forecast is therefore conditionally unbiased.

While market forecasts remain efficient, the presence of bubbles will generally invalidate the usual efficiency tests. Consider, for example, a test for bias in exchange-rate forecasts. Assume, for simplicity, that $\alpha'\phi(t)$ itself follows a martingale process, so that its conditional expectation equals its own lagged value. Then

$$(10)\quad s(t+1) - E_t[s(t+1)] = \nu^{t+1}(t+1) + \eta(t+1)[(1+\lambda)/\lambda]^{t+1}.$$

A significance test for the sample mean

$$M(T) = (1/T)\sum_{t=1}^{T}\{s(t+1) - E_t[s(t+1)]\}$$

is a test for unconditional bias in exchange-rate forecasts. But the variance of M(T) can be calculated as

$$\mathrm{Var}[M(T)] = (1/T^2)\sum_{t=1}^{T}\{\sigma_\nu^2 + \sigma_\eta^2[(1+\lambda)/\lambda]^{2(t+1)}\}$$

when ν and η are stationary and uncorrelated. This variance does not go to zero as $T \to \infty$. It follows that the statistic M(T) will not necessarily give a consistent estimate of the mean forecast error in the presence of a bubble. Even though agents are making optimal forecasts, the ususal tests might give strong indications of inefficiency.

As Meese notes, however, bubbles are an unlikely explanation of the puzzle of forward exchange rate behavior. Under the type of bubble described by (9), the variance of one-step-ahead forecast errors explodes over time at rate $[(1+\lambda)/\lambda]^2$, as (10) shows. Taking the logarithm of the forward rate as an approximation to the expected logarithm of the future spot rate gives a rough empirical measure of this forecast error under risk neutrality. The stylized facts recounted by Meese imply that the variance of forward forecast errors does not explode over time. Nor is such behavior observed over subperiods. In particular, a bubble as in (9) would be inconsistent with the approximate random walk followed by spot and forward rates.[3]

There are two other reasons for dismissing divergent speculative bubbles as a characterization of exchange-rate behavior. First, there is a strong theoretical case for ruling them out (as discussed in my paper with Rogoff). Second, as first pointed out by Robert Flood and Peter Garber, bubbles are observationally equivalent to possible changes in the $\phi(t)$ process that the public may expect with good reason.

The observational equivalence of bubbles and anticipated future changes is illustrated by an example related to the one analyzed in the previous section.[4] Suppose again that α is a scalar equal to 1 and that $\phi(t)$ is the money supply. Until a known date T in the future, $\phi(t)$ is expected to equal the random variable $m(t)$ with probability 1, where $m(t) = m(t-1) + \omega(t)$, as before. There is a possibility, however, that the $\phi(t)$ process will change permanently on date T. Specifically, the public knows that the money supply for $t' \geq T$ will be given by the random variable $\phi(t') = m(t') + \mu(T)$, where $\mu(t) = [1 - d(t)][\mu(t-1) + \varepsilon(t)]$ (for $t \leq T$) and $d(t)$ and $\varepsilon(t)$ are as defined earlier. In words,

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[3] There are some models in which the bubble term involves the power of a root that is smaller than unity in absolute value. These convergent bubbles are difficult to rule out by inspecting the data, but on the other hand they do not necessarily destroy the validity of the standard efficiency tests. It is therefore unlikely that convergent bubbles are responsible for the apparent failure of risk-neutral foreign exchange pricing.

[4] A similar example illustrating this equivalence is give by James Hamilton.

the public expects a sharp jump in the money supply at time T if the event d(t) = 1 does not occur in the interim. The expected size of the possible jump, $\mu(t)$, evolves as a random walk.

Consider next the "bubble-free" exchange-rate solution (2), which is now equal to

$$s(t) = m(t) + (1-\pi)^{T-t}[1/(1+\lambda)] \sum_{i=T-t}^{\infty} [\lambda/(1+\lambda)]^i \mu(t).$$

Rewrite this equation as

$$s(t) = m(t) + K[\mu(t)/(1-\pi)^t][(1+\lambda)/\lambda]^t$$

(where $K = [(1-\pi)\lambda/(1+\lambda)]^T$); since $E_t[\mu(t+1)/(1-\pi)^{t+1}] = (1-\pi)\mu(t)/(1-\pi)^{t+1} = \mu(t)/(1-\pi)^t$ (conditional on d(t) = 0), the result is a special case of (9), corresponding to $E_t[\alpha'\phi(t+i)] = m(t)$ for <u>all</u> i. The equation therefore shows that an econometrician who is unaware that the public expected a possible change in the money-supply process will be unable to distinguish the exchange rate's bubble-free path from that of a "crashing" bubble such as the one analyzed by Olivier Blanchard. The strong theoretical case against divergent bubbles suggests that econometric results purporting to show their existence are more likely to be the result of misspecifying agents' information sets.

It is noteworthy that the example just given and (by implication) some bubbles may lead to peso problems. This coincidence does not imply that peso problems and bubbles are the same. The model of the last section, for example, shows that peso problems need not give rise to explosive exchange-rate behavior.

The bubble model embodied in equation (9) assumes rational one-step-ahead forecasting by market participants, but it imposes no terminal convergence condition to tie the exchange rate to economic fundamentals. As Meese notes, some researchers have suggested that market forecasts are not rational, and that exchange rates may be prone to "over-reaction" or "excessive" volatility. Under risk neutrality, the recently-developed volatility tests applied to stock and bond markets offer a "portmanteau" methodology that potentially detects bubbles as well as other deviations from efficient asset pricing based on fundamentals. How one should interpret these tests in practice is currently a matter of some controversy. (Joe Mattey and Meese provide a survey of the econometric issues in the debate, together with the results of Monte Carlo experiments.) In paticular, volatility tests, like tests for bubbles, require strong identifying assumptions about market expectations.

Even if one accepts the methodology of volatility testing in principle, it is not clear how it can be applied to the foreign exchange market. Economists have developed widely accepted theories of both stock pricing and the term structure of interest rates under risk neutrality and rational expectations; it is these theories upon which the volatility tests build. No correspondingly clear-cut theory of risk-neutral nominal or real exchange-rate determination exists. A consensus theory of exchange rates would require agreement on the determinants of relative demands for fiat monies, as well as on the dividing line between money and near-money. Such a theory would also require a consensus model of demand and supply in national output markets.

III. Risk

Risk aversion on the part of market participants provides a third potential explanation of apparent deviations from uncovered interest parity. Although much of the empirical literature on the risk premium reports negative results, more recent research lends limited support to the empirical relevance of risk-averse asset pricing.

To see how risk aversion gives rise to departures from interest parity, consider the relation between the nominal interest rates on one-period bonds denominated in home and foreign currency. An international investor following an optimal consumption/investment plan must be indifferent among the three alternatives of using an additional domestic-currency unit to augment his domestic-currency assets, allocating the money to foreign-currency assets, or spending the money on consumption goods today. Indifference occurs when each course of action yields the same expected marginal return. Let the investor's period utility function be $u[C(t)]$, let $\beta < 1$ denote the subjective time-preference factor, let R (R^*) denote the domestic (foreign) one-period nominal interest rate, and let P (P^*) be the local money price of the typical home (foreign) consumption bundle. For a domestic investor, the optimality condition described above is

$$\frac{u'[C(t)]}{P(t)} = \beta[1+R(t)]E_t\left\{\frac{u'[C(t+1)]}{P(t+1)}\right\} = \beta[1+R^*(t)]E_t\left\{\frac{u'[C(t+1)]S(t+1)}{P(t+1)S(t)}\right\},$$

where $S(t)$, as before, is the level (rather than the log) of the exchange rate. Define $Q(t+1)$ to be the ex post marginal rate of substitution between domestic currency units available on dates t and $t+1$; thus, $Q(t+1) = \{\beta u'[C(t+1)]/P(t+1)\}\div\{u'[C(t)/P(t)\}$. The equation above can be rewritten in terms of this definition as

$$(11)\quad 1 = [1+R(t)]E_t[Q(t+1)] = [1+R^*(t)]E_t[Q(t+1)S(t+1)/S(t)],$$

which implies

$$(12)\quad [1+R(t)]/[1+R*(t)] = E_t[Q(t+1)S(t+1)/S(t)]/E_t[Q(t+1)].$$

Because the one-period forward exchange rate F(t) is equated by covered interest parity to S(t)[1+R(t)]/[1+R*(t)], equations (11) and (12) above lead to Meese's equation (2) for the forward rate.

While (11) and (12) were derived by considering the alternatives available to a typical domestic investor, consideration of a foreign investor would lead to the same expressions. As noted earlier, $E_t[Q(t+1)]$ is an expected marginal rate of substitution between present and future home-currency units. $E_t[Q(t+1)S(t+1)/S(t)]$ can likewise be interpreted as an expected marginal rate of substitution between present and future units of the foreign currency. In the equilibrium of an internationally integrated asset market, all investors should have the same expected intertemporal substitution rate for each currency they trade. Both the numerator and the denominator on the right-hand side of (12) are therefore independent, in equilibrium, of the particular investor considered, provided all investors have access to the same bond markets.[5]

The implications of (12) for uncovered interest parity are understood most easily if that equation is written in the form

$$(13)\quad \frac{1+R(t+1)}{1+R*(t+1)} - E_t\left[\frac{S(t+1)}{S(t)}\right] = \frac{\mathrm{Cov}_t[Q(t+1),S(t+1)/S(t)]}{E_t[Q(t+1)]}.$$

Equation (13) shows the deviation from uncovered interest parity that

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[5] In my 1987 paper I test the international equality of intertemporal marginal substitution rates between the U.S. and Germany and between the U.S. and Japan. For post-1972 data, the tests I conduct fail to reject those equalities.

arises in the present model; the right-hand side of the equation is usually referred to as the (one-period) foreign exchange risk premium on domestic currency. The risk premium is positive if the home currency's depreciation against foreign currency is unexpectedly high at times when the increase in the marginal utility from spending a domestic currency unit is also unexpectedly high. In other words, the home currency is riskier if it tends to lose foreign-exchange value in states of the world where it is most needed. A positive risk premium implies that the domestic-foreign interest ratio exceeds the expected depreciation of home currency, while a negative risk premium implies the opposite.

The right-hand side of equation (13) equals the one-period forward premium, $F(t)/S(t)$, less expected depreciation. Thus, the equation also reveals that the forward-premium forecast error, $[S(t+1) - F(t)]/S(t)$, consists of a true, random forecast error less a risk premium that generally varies over time. Since the risk premium is a function of time-t information, forward-premium forecast errors are also predictable on the basis of that information. Many researchers therefore ascribe the apparent partial predictability of forward-premium errors to time-varying risk premiums. This explanation is sensible only if risk premiums are sufficiently variable that their influence is not dominated by that of the pure expectational errors.

If the relevant economic variables are jointly lognormally distributed (because their logs are generated by a linear Gaussian process), then (12) becomes

$$(14)\quad r(t) - r^*(t) - E_t[\Delta s(t+1)] = Var_t[\Delta s(t+1)]/2 + Cov_t[q(t+1),\Delta s(t+1)],$$

where $r(t) = \log[1+R(t)]$, $r^*(t) = \log[1+R^*(t)]$, and $q(t) = \log[Q(t)]$. Lars Peter Hansen and Hodrick (1983) first derived this widely-tested logarithmic model within a utility-maximizing framework. A major implication of this model is that the conditional variance and covariance on the right-hand side of (14) are both constant (given stationarity), implying a constant risk premium. Constancy follows from the fact that the innovations in q(t+1) and Δs(t+1), being normal, are independent of time-t variables, not just uncorrelated with them. Since the risk premium must vary over time to explain the predictability of forward forecast errors by past information, the logarithmic model is clearly inconsistent with the data. Another of the model's implications, the conditional homoskedasticity of forward-rate forecast errors, also contradicts the facts. Conditional heteroskedasticity may be casued by peso problems, as noted above, but there are obviously other possible causes.

As Meese documents, alternative models of the risk premium also seem to do poorly when confronted with the data. But I think it is too early to conclude, as does Jeffrey Frankel, that risk-averse behavior can explain only a negligible fraction of the observed serial correlation in forward forecast errors. In recent work, Cumby (1986a) presents direct estimates of foreign exchange risk premiums that provide some support for the risk-averse asset-pricing model described by equation (13).[6] My own preliminary estimates (1987) of a logarithmic, consumption-based model of real interest-rate differentials are also encouraging. The problems discussed in section I above are almost certainly important, however, and one goal of research should be to take these data problems explicitly into account. The role of risk aversion

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[6] Cumby's (1986b) results on stock-market prices in Germany, Japan, the United Kingdom, and the United States appear less supportive of the model.

cannot be assessed adequately until more refined procedures for purging the data have been developed.

Another promising direction for future research is to integrate studies of foreign exchange pricing with studies of price determination in other asset markets. The (rational) expectations model of the term structure, like uncovered interest-rate parity, has been widely tested and rejected. As documented by Robert Shiller, John Campbell, and Kermit Schoenholtz, the interest-rate forecasts implicit in the term structure, like forward foreign exchange premiums, often mispredict the direction of future rate movements in post-1959 data. It is more than an appeal to Occam's razor to suggest that a single theory should explain these similar phenomena. In view of the increasingly tight links between onshore and offshore asset markets, it is difficult to conceive of a convincing account of the foreign-exchange risk premium that does not simultaneously throw light on the characteristics of term premiums and stock-market risk premiums.

Consider, for example, the tight connection between term premiums and foreign-exchange premiums implied by covered interest parity. Let $F^{(i)}$ denote the i-period forward exchange rate and $R^{(i)}$ ($R*^{(i)}$) the i-period interest rate on a domestic (foreign) bond. Covered interest parity ensures that for any i,

$$F^{(i)}(t)/S(t) = [1+R^{(i)}(t)]^{i}/[1+R*^{(i)}(t)]^{i}.$$

To be concrete, apply the above expression to compute the ratio of the two-period forward premium to the one-period forward premium. The result is

$$(15)\qquad \frac{F^{(2)}(t)/S(t)}{F^{(1)}(t)/S(t)} = \frac{[1+R^{(2)}(t)]^{2}/[1+R^{(1)}(t)]}{[1+R*^{(2)}(t)]^{2}/[1+R*^{(1)}(t)]}.$$

Let τ ($\tau*$) denote the domestic (foreign) one-period term premium and let $\theta^{(i)}$ (i = 1,2) denote the i-period foreign-exchange risk premium. Reasoning similar to that followed above shows that

$$\tau(t) = \mathrm{Cov}_t\{Q(t+1)Q(t+2),[1+R^{(1)}(t+1)]\}/E_t[Q(t+1)Q(t+2)].$$

The foreign term premium $\tau*(t)$ depends in the same way on the covariance between the foreign one-period interest rate and the ex post marginal rate of substitution between foreign currency units delivered on dates t and t+2. Equation (15), expressed in terms of these risk premiums, can be shown to be

$$\frac{E_t[S(t+2)/S(t)] + \theta^{(2)}(t)}{E_t[S(t+1)/S(t)] + \theta^{(1)}(t)} = \frac{E_t[1+R^{(1)}(t+1)] + \tau(t)}{E_t[1+R*^{(1)}(t+1)] + \tau*(t)}.$$

Expressions such as this one imply restrictions on the joint behavior of foreign-exchange risk premiums and bond-market term premiums.

Intriguing results have come out of research that examines several asset markets at the same time. Richard Clarida and Campbell, for example, find that variables with explanatory power for term premiums also help predict returns to forward foreign exchange speculation. Giovannini and Jorion report similarities in the conditional distributions of foreign-exchange and U.S. stock-market risk premiums. In time, further investigation into findings such as these may help explain the puzzling behavior of forward exchange rates.

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