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Excess Volatility and the Asset-Pricing Exchange Rate Model with Unobservable Fundamentals

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Abstract

This paper presents a method to test the volatility predictions of the textbook asset-pricing exchange rate model, which imposes minimal structure on the data and does not commit to a choice of exchange rate "fundamentals." Our method builds on existing tests of excess volatility in asset prices, combining them with a procedure that extracts unobservable fundamentals from survey-based exchange rate expectations. We apply our method to data for the three major exchange rates since 1984 and find broad evidence of excess exchange rate volatility with respect to the predictions of the canonical asset-pricing model in an efficient market.

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I. Introduction

Whether exchange rates are "too volatile" with respect to the behavior of their underlying determinants is one of the most persistent, and yet unsettled, issues in the international finance debate. The matter is clearly critical for policy purposes: if exchange rates are too volatile with respect to a reasonable benchmark in an efficient market, then there may be grounds for throwing "sand in the gears" of currency markets, so as to slow down changes in exchange rates and keep these rates in line with their "fundamentals" (see, for instance, Eichengreen, Tobin, and Wyplosz, 1995). Conversely, if exchange rate volatility is largely consistent with that predicted by conventional models, then one may simply have to swallow the rates' abrupt swings as the potentially efficient response to underlying shocks.

Applying the methods suggested by Shiller (1981) in the context of stock price volatility, a number of studies in the 1980s, including Huang (1981), Vander Kraats and Booth (1983), and Wadhwani (1987), first tried to settle this issue. These studies benchmarked the volatility of exchange rates against the predictions of the monetary model, and popularized the perception that exchange rates over the post-Bretton Woods period had been too volatile for their behavior to be consistent with their fundamentals in an efficient foreign exchange market.

Later research, however, cast doubt on these results. Diba (1987), for instance, pointed to the miscalibration of Huang's (1981) and Vander Kraats and Booth's (1983) tests. When correctly calibrated, these tests failed to provide evidence of excess volatility. More generally, Kleidon (1986), Marsh and Merton (1986), and others showed that statistical biases in early volatility tests of asset price models—including those of exchange rate models—caused those models to be rejected too often in finite samples. Some of these problems have been later addressed (in the international field, for instance, see West, 1987, Gros, 1989, and Bartolini and Bodnar, 1996), while a parallel line of research has tried to establish to what extent fluctuations in exchange rates could be ascribed to changes in fundamentals, or to the realizations of speculative bubbles or "sunspots" (see Frankel and Rose, 1995, for a survey of this effort, and Jeanne and Masson, 1998, for a more recent contribution). Like previous work, however, these studies relied heavily on specific structural models (e.g., the monetary model) as benchmarks for exchange rates' *normal* volatility.

The problem with reliance on specific structural models is that evidence of excess volatility may simply indicate that the assumed structural relationships fail to track exchange rates accurately—not a surprising finding, given the poor empirical record of most structural exchange rate models in recent decades. Measurement issues are also crucial for the finding of excess volatility: the econometrician must capture theoretical fundamentals by choosing between many possible empirical counterparts and must usually abstract from political, psychological, and other unobservable factors explaining investors' demand for currencies. Ultimately, the more inadequate is the model's empirical implementation (e.g., the narrower the choice of fundamentals), the more one is likely to find unexplainable—and hence excessive—movements in exchange rates, clearly turning the incentives for proper hypothesis testing upside down. Overall, while the view that exchange rates are too volatile to reflect efficient currency trading is widespread (see, for instance, Boertje and Garretsen, 1996, Masson, 1998, and Jeanne and Rose, 1999, among many others), its empirical support remains weak and anchored to controversial model- and data-specific evidence.

To partially address this gap, this paper presents a new approach to volatility tests of the standard asset-pricing view of the exchange rate which requires minimal commitment to an exchange rate model and no commitment at all to a choice of relevant fundamentals. Our approach is developed as the combination of (i) existing tests of excess volatility of asset prices (see Mankiw, Romer, and Shapiro, 1991, and Bollerslev and Hodrick, 1995); (ii) a technique suggested by Flood, Mathieson, and Rose (1991) and Flood and Rose (1996) to retrieve unobservable fundamentals from data on exchange and interest rates, assuming uncovered interest parity; and (iii) the insight that one can dispose of the uncomfortable interest parity condition by using survey-based data to measure exchange rate expectations directly.

As a result of this effort, our testing approach features several desirable properties. First, it relies only on the assumption that exchange rates satisfy a rather general asset-pricing relationship—namely, the canonical log-linear pricing model. Second, it allows for the widest possible definition of "fundamentals," requiring no commitment to any set of measurable or nonmeasurable variables. Third, it inherits the properties of the methodology of Mankiw, Romer, and Shapiro (1991), such as its good small-sample performance and its robustness to speculative bubbles. Aside from its application to volatility tests, our approach also suggests an easy way to construct unobservable fundamentals for potentially wide use in empirical work on exchange rates.

We apply our method to data from the three major currency markets and find evidence of excess volatility of exchange rates, conditional on a significantly narrower set of assumptions than previous research. This evidence clearly contradicts the joint assumptions that the markets we study are efficient and that the canonical asset-pricing model of the exchange rate adequately captures the behavior of our sample currencies. Because of the limited structure we impose on the data, this evidence should be regarded as stronger than that available from previous research.

The paper is structured as follows. Section 2 presents our methodology. Section 3 calibrates the tests, applies them to data for the three major exchange rates, and discusses variants of our methodology and the robustness of our findings. Section 4 concludes.

II. Volatility Tests of the Canonical Asset-Pricing Exchange Rate Model

Our starting point is the familiar forward-looking, log-linear model of the exchange rate:

$$s_t = f_t + \alpha\left(\mathrm{E}_t\left[s_{t+1}\right] - s_t\right), \tag{1}$$

where s_t denotes the (log) exchange rate, defined as the domestic price of a unit of foreign currency, f_t denotes variables fundamental to the determination of the exchange rate, and $\mathrm{E}_t[\,.\,]$ denotes the usual rational expectation operator, based on information available at time t.

Equation (1) expresses the exchange rate as the sum of current fundamentals and a linear function of its own expected change, and summarizes the "asset market view" of the exchange rate that has become standard in the literature since Mussa (1976). Model (1) can be viewed, inter alia, as the reduced form of a monetary model linking the exchange rate to money supplies and incomes, with money-demand equations possibly derived from money-in-utility (see Stockman, 1980) or cash-in-advance assumptions (see Lucas, 1982); it can also be interpreted as a more general equilibrium model, such as a model with currency-substitution (see Calvo and Rodriguez, 1977). Flexible interpretation of f_t may include announcements (or signals) of current and future monetary policies, political factors, and a variety of measurable or hard-to-measure variables that affect investors' demand for currencies. The analogy of model (1) with models of stock valuation is also apparent—f_t plays the role of dividends and α that of the discount factor applied to expected future capital gains—and partly motivates the methodology that follows.

Solving equation (1) forward for s_t, up to time $t+h$, yields:

$$s_t = \frac{1}{1+\alpha}\left(\sum_{i=0}^{h-1}\left(\frac{\alpha}{1+\alpha}\right)^i \mathrm{E}_t\left[f_{t+i}\right]\right) + \left(\frac{\alpha}{1+\alpha}\right)^h \mathrm{E}_t\left[s_{t+h}\right], \tag{2}$$

where h is the holding period for the position opened by purchasing foreign currency at time t.

It is common in the literature to assume the absence of speculative bubbles, namely, that $\lim_{i\to\infty}\left(\frac{\alpha}{1+\alpha}\right)^i \mathrm{E}_t\left[s_{t+i}\right]=0$, an assumption that allows solving equation (2) forward solely as a function of fundamentals. However, the tests considered here do not require a no-bubble assumption and can be performed directly on (2).

Following the method used by Mankiw, Romer, and Shapiro (1991) in the context of stock price models, let's now define the perfect-foresight (or "fundamental") rate s_t^* as the value that the exchange rate would take if investors could predict with certainty future fundamentals f_{t+i}, $i=1,\ldots,h-1$, and the future exchange rate s_{t+h}. s_t^* is obtained by dropping the expectation operator from (2):

$$s_t^* = \frac{1}{1+\alpha}\left(\sum_{i=0}^{h-1}\left(\frac{\alpha}{1+\alpha}\right)^i f_{t+i}\right) + \left(\frac{\alpha}{1+\alpha}\right)^h s_{t+h}, \tag{3}$$

so that, by definition, $s_t=\mathrm{E}_t\left[s_t^*\right]$.

Next, we construct a "benchmark" exchange rate, denoted by s_t^o, from which the volatility of both market and fundamental rates can be measured. The main reason behind the need of a benchmark rate is that exchange rates are typically nonstationary, and therefore volatility must be measured on deviations of exchange rates from a specific (stochastic) trend. Accordingly, the first requirement for the choice of s_t^o is that the differences $(s_t - s_t^o)$ and $(s_t^* - s_t^o)$ be stationary, to assure their mean square errors to be well defined. The second requirement is that s_t^o be known at time t to investors. This condition assures the orthogonality of s_t^o with rational forecast errors based on information available at t, a property whose usefulness for our tests will soon become clear.

Many different definitions of s_t^o would satisfy these two requirements, however. We begin by following Mankiw, Romer, and Shapiro (1991) who, studying stock price volatility, suggest defining s_t^o as some "naive" price forecast, for instance as the price that would prevail in the market if investors expected dividends to evolve as a random walk. We follow a similar approach and define s_t^o as the rate that would prevail if investors expected fundamentals to follow a random walk, i.e., $f_t = \mathrm{E}_t[f_{t+i}], i \geq 0$. Under this assumption, equation (1) yields $s_t^o = f_t$. Thus, if fundamentals were to follow a random walk, under the assumptions of model (1) so would the benchmark rate s_t^o. Note that s_t^o need be neither a *rational* nor an empirically accurate forecast. (That said, a random-walk forecast may have claim to realism, as it is usually difficult to distinguish exchange rates from a random walk at short horizons.) The choice of s_t^o affects the outcome of the tests, however, by determining the information on which the tests are conditioned, as discussed below. We document this dependence in Section 4 by considering alternative definitions of s_t^o.

Having thus defined fundamental and benchmark rates, the methodology we follow to test model (1) can now be introduced. In an efficient market, where current exchange rates rationally reflect available information on future rates and fundamentals, the forecast errors $(s_t^* - s_t)$ should be uncorrelated with any variable known at time t, including $(s_t - s_t^o)$. That is,

$$\mathrm{E}_t\left[\left(s_t^* - s_t\right)\left(s_t - s_t^o\right)\right] = 0 . \tag{5}$$

Therefore, squaring both sides of the identity:

$$\left(s_t^* - s_t^o\right) \equiv \left(s_t^* - s_t\right) + \left(s_t - s_t^o\right), \tag{6}$$

taking expectations, and using (5), yields:

$$\mathrm{E}_t\left[\left(s_t^* - s_t^o\right)^2\right] = \mathrm{E}_t\left[\left(s_t^* - s_t\right)^2\right] + \mathrm{E}_t\left[\left(s_t - s_t^o\right)^2\right], \tag{7}$$

or,

$$q_t \equiv \mathrm{E}_t\left[\left(s_t^* - s_t^o\right)^2\right] - \mathrm{E}_t\left[\left(s_t^* - s_t\right)^2\right] - \mathrm{E}_t\left[\left(s_t - s_t^o\right)^2\right] = 0 . \tag{8}$$

Thus, model (1), combined with the assumption of market efficiency, implies the restriction $q_t = 0$, and hence the restriction $\mathrm{E}[q_t] = 0$. Hence, the sample mean of the q_t's, $\bar{q}$, should be close to zero if the market is efficient and model (1) correctly describes the exchange rate process.

One can then perform a test of model (1) as follows. First, one must calibrate the parameter α and define a suitable fundamental series f_t. This allows computing the series q_t and its sample average $\bar{q}$. One can then reject the hypothesis that $\mathrm{E}[q_t]=0$ if $\bar{q}$ is different from zero, by judging the statistical significance of this deviation, for instance, by a Generalized Method of Moments distribution for $\bar{q}$ (see Bollerslev and Hodrick, 1995). This is a normal distribution:

$$\bar{q} \sim \mathrm{N}\left(0, \frac{V}{T}\right), \tag{9}$$

whose variance, V/T, can be estimated using the method proposed by Andrews and Monahan (1992), to make the test robust to heteroskedasticity and serial correlation in the series q_t.[2]

Next, excess volatility tests can be constructed by noting that equation (7) also implies the inequalities:

$$\mathrm{E}\left[\left(s_t^* - s_t\right)^2\right] \le \mathrm{E}\left[\left(s_t^\circ - s_t^*\right)^2\right], \tag{10a}$$

$$\mathrm{E}\left[\left(s_t - s_t^\circ\right)^2\right] \le \mathrm{E}\left[\left(s_t^\circ - s_t^*\right)^2\right]. \tag{10b}$$

Should either one of these inequalities be violated, then $\mathrm{E}_t[q_t]$ would be negative (though the opposite need not be true). Equation (10a) states that the market exchange rate s_t should be less volatile around the fundamental rate s_t^* than the benchmark rate s_t°, in terms of the usual mean-square error criterion. Similarly, (10b) states that the market rate should be less volatile around the benchmark rate than the fundamental rate. Thus, (10a) and (10b) can be used as excess volatility tests of model (1), and involve a simple intuition. For (10b), for instance, the fundamental exchange rate deviation from the benchmark rate, $\left(s_t^\circ - s_t^*\right)$, equals the market rate deviation from the same benchmark, $\left(s_t - s_t^\circ\right)$, *plus* a forecast error $\left(s_t^* - s_t\right)$. If markets are efficient and model (1) is valid, this error should merely add noise to current information, i.e., it should not be systematically related to information available at time t, including $\left(s_t - s_t^\circ\right)$. The volatility of the fundamental rate around the benchmark should then exceed that of the market rate, and similarly for (10a).

A number of steps in the testing procedure just outlined require discussion. First, a choice must be made as to the value of α that parameterizes model (1) and the volatility tests. Although α could—in principle—be estimated (see, for instance, Flood, Rose and Mathieson,

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[2]This method involves filtering q_t by an AR(1) process and estimating the asymptotic variance Ω of the filtered residuals, $\hat{u}_t$, by $\hat{\Omega} = C(0) + 2\sum_{j=1}^{h-1} \kappa\left(\frac{j}{h}\right) C(j)$, where $C(j) = \sum_{t=j+1}^{T} \frac{\hat{u}_{t+h}\hat{u}_{t+h-j}}{T}$, $\kappa(j/h)$ is a quadratic spectral kernel, and the bandwidth parameter h is chosen according to an automatic data-based procedure. $\hat{V}$ is then recovered by multiplying $\hat{\Omega}$ by the square of the inverse of the AR filter.

1991, and Bartolini and Bodnar, 1992), previous efforts have yielded such a broad and imprecise array of estimates that an agnostic approach seems preferable. Here we select three values of α (α=0.1, α=1, and α=5), which previous research suggests to be representative of the low, typical, and high ranges of α, respectively. Then, we present results of tests calibrated with all these values. Results of tests calibrated with different values of α are available upon request.

Second, one must obtain a series for f_t. Common practice in the exchange rate literature—particularly in the context of tests of excess volatility—is to specify a structural exchange rate model and let this model guide selection and aggregation of fundamentals. For instance, the textbook monetary model sets f_t as a linear combination of domestic and foreign money supplies and incomes. More general models with sticky prices and sluggish money adjustments would include prices and lagged money stocks among fundamentals.[3]

Instead, here we proceed as follows. First, we follow Flood, Rose and Mathieson (1991), and Flood and Rose (1996), and use model (1) to generate a series of fundamentals as:

$$f_t \equiv (1+\alpha)s_t - \alpha \mathrm{E}_t\left[s_{t+1}\right]. \tag{11}$$

Next, differently from Flood, Rose, and Mathieson (1991) and Flood and Rose (1996), who estimate expected future exchange rates using uncovered interest parity, i.e., $\mathrm{E}_t\left[s_{t+1}\right] \equiv s_t + i_t - i_t^*$, we use survey-based exchange rate expectations to measure $\mathrm{E}_t\left[s_{t+1}\right]$ directly.[4] Our use of survey data (see below for a description) allows us to dispose of the assumption of uncovered interest parity whose empirical record is, at best, questionable. This completes the parameterization of our test.

To clarify the implications of our method of calculating implied fundamentals for our testing procedure, first substitute (11) into (3) for all i from t to $t+h$. Then denote the exchange rate "news" at t (i.e., the forecast error realized at t) by $\epsilon_t \equiv s_t - \mathrm{E}_{t-1}\left[s_t\right]$, and rearrange terms so as to express s_t^* as a function of s_t and future forecast errors as:

$$s_t^* = s_t + \sum_{i=1}^{h}\left(\frac{\alpha}{1+\alpha}\right)^i \epsilon_{t+i} \equiv s_t + \eta_t\,, \tag{12}$$

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[3]Gardeazabal, Regúlez, and Vázquez (1997) are an exception to this rule, as they also test model (1) allowing for unobservable fundamentals. Beyond this similarity, however, these authors' technique and ours differ substantially, and their tests—in contrast with ours—do not reject model (1).

[4]In addition to using the interest parity condition, Flood and Rose (1996) also build fundamentals directly from macro data, as suggested by the monetary model of the exchange rate.

where $\eta_t \equiv \sum_{i=1}^{h} \left(\frac{\alpha}{1+\alpha} \right)^i \epsilon_{t+i}$ is the discounted stream of news between $t+1$ and $t+h$.

Now compare expression (12) with expression (5)—which is the fundamental restriction underlying our tests. Since, by (12), $s_t^* - s_t = \eta_t$, then a finding that $\bar{q}$ is (significantly) different from zero provides evidence that the composite forecast error η_t is correlated with information available at time *t*. This result, in turn, indicates that either exchange rates do not reflect efficiently all the information available to the market, or that the aggregation of forecast errors ϵ_t suggested by model (1) is invalid, or both.

Next, note some of the advantages of our testing procedure with respect to previous tests of excess volatility of exchange rates. The most important of these advantages is our ability to restrict the set of joint theoretical assumptions being tested to the assumptions of market efficiency and of model (1). This approach contrasts sharply with that of previous studies, which relied heavily on functional forms for money demands, price-adjustment equations, and interest-parity and purchasing-power-parity conditions, among others. Our approach also leaves the fundamental series f_t open to the widest possible interpretation of what fundamental factors may contribute to the determination of exchange rates in model (1).

Other advantages of our methodology are inherited directly from our use of the econometric procedure of Mankiw, Romer, and Shapiro (1991). For instance, our tests are independent of the presence (or lack thereof) of speculative exchange rate bubbles. This feature reflects the definition of the fundamental rate s_t^* as a discounted stream of fundamentals only up to time *t+h* (see expression (3)). Based on this definition, a hypothetical bubble term would be incorporated in both s_t and s_t^* through s_{t+h}, and hence would not alter the forecast error $\left(s_t - s_t^*\right)$.[5]

III. Empirical Results

A. Data and Calibration

We applied our methodology to data for the three major exchange rates: the British pound, the Deutsche mark, and the Japanese yen against the U.S. dollar. As noted above, our tests require only data on spot and expected future rates. However, it is important that the spot data be sampled on the same day in which the survey of expectations is conducted, for

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[5]Mankiw, Romer, and Shapiro (1991) also show that the general volatility test applied here exhibits better statistical properties than traditional regression-based tests of the same model. In particular, the test's finite-sample distribution is well approximated by its asymptotic distribution, causing the under-lying model to be rejected the correct number of times, on average. Traditional regression-based tests of asset-price models such as (1), by contrast, are likely to be biased in small samples when prices and fundamentals are nonstationary (as is typically the case with exchange rates), rejecting the underlying model too often in finite samples.

our procedure to generate fundamentals that are synchronous with the exchange rate they are supposed to determine. Conveniently, the *Financial Times Currency Forecaster* archive (previously *Currency Forecaster's Digest*) provides synchronized spot and survey-based expectations data. From this archive we drew data for all the available forecast horizons, namely, the rates expected to prevail one, three, six, and twelve months after each survey date. We used the entire sample available at the time of writing, including monthly data from April 1984 to December 1998.

Our tests are parameterized by α, a plausible range of which we identified drawing from previous empirical studies. In surveying available evidence, we found studies attempting to estimate α structurally (e.g., as the annualized semi-elasticity of money demand to interest rates) typically yielding estimates between α=1 and α=5 (see, for instance, Fair, 1987, Goldfeld and Sichel, 1990, and Laidler, 1993). By contrast, we found reduced-form estimates of α, obtained directly from exchange rate data, to be much smaller, typically ranging between α=0.1 and α=1 (see, for instance, Flood, Rose, and Mathieson, 1991, and Bartolini and Bodnar, 1992). Overall, this evidence suggests viewing the value α=1 as a reasonable baseline for α, and the values α=0.1 and α=5 as representative of the low and high ranges for the same parameter, respectively. (In calibration, these annualized values must be normalized by dividing by the forecast horizon, defined in units of year.) The next section presents results for tests calibrated with these values.[6]

B. Baseline Results

The top panels of Tables 1-6 report standardized (asymptotically) normal statistics for tests of the null hypothesis H_o: $E[q_t]=0$, along with a superscript indicating whether inequalities (10a) and (10b) are violated. A failure of inequality (10a) is indicated by a superscript "a," while a failure of inequality (10b) is indicated by a superscript "b." The bottom panels of the tables report Phillips-Perron Z_t tests of unit roots in the series $(s_t^* - s_t)$, $(s_t - s_t^o)$, and $(s_t^* - s_t^o)$, the stationarity of which is required for the validity of our volatility tests. Our theoretical prior is that—given a long enough sample—these series should be recognized as stationary: we do not expect "market," "fundamental," or "forecast" rates to drift infinitely apart as the sample's length grows to infinity. In particular, having defined the benchmark rate as $s_t^o = f_t \equiv (1+\alpha)s_t - \alpha E_t[s_{t+1}]$ allows for s_t^o to be cointegrated with s_t and with s_t^*, provided that survey-based expectations are cointegrated with s_t.[7] However, the length of our sample is limited by the availability of survey data, and we must check for potential problems of nonstationarity due to short sample. To highlight potential problems of this sort, we write the volatility test statistics within single parentheses and double

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[6]Complete details of our tests are available upon request, along with the data and the Gauss program we used to perform them.

[7]Evidence of cointegration between spot rates and survey expectations from the *Financial Times Currency Forecaster* is provided in Giorgianni (1996).

parentheses, respectively, when any of the underlying stationarity tests fails to reject the null hypothesis of a unit root at the five and ten percent significance levels (the critical values are -2.91, and -2.59, respectively). The stationarity tests that we present follow the data-based methodology proposed in Andrews and Monahan (1992) to determine the amount of correction for auto-correlation in the residuals of the auxiliary regressions used to perform the unit root tests. Alternative stationarity tests we employed yielded very similar results.

Table 1-4 present results for tests computed for different forecast horizons: one month for Table 1, three months for Table 2, six months for Table 3, and twelve months for Table 4. The holding period h is held constant at three months in these tables; it is allowed to vary from one month to twelve months in Table 5, discussed below.

Results from Table 1-4, the columns of which report tests performed for different values of α, show the volatility test statistics and the stationarity properties of $\left(s_t^* - s_t\right)$ and $\left(s_t^* - s_t^{\circ}\right)$ to be fairly insensitive to changes in α. By contrast, violations of inequalities (10a) and (10b) depend on α.[8] There is some uniformity of results across the three rates, likely reflecting the common sensitivity of the three exchange rates to dollar shocks.

Two main sets of results emerge from Tables 1-4, the first of which pertains to tests of the hypothesis $\mathrm{E}\left[q_t\right]=0$. The joint assumptions of market efficiency and of applicability of model (1) can be rejected with confidence in all cases (i.e., for all forecast horizons and all values of α) for the mark/dollar rate, and in almost all cases for the sterling/dollar rate (the only exceptions occur for relatively high values of α at the forecast horizon of twelve months). The results are only slightly weaker for the yen/dollar rate, for which the same assumptions fail to be confidently rejected only at the forecast horizon of one month and—for higher values of α—at the forecast horizon of three months. For this specification of the tests, the validity of these results is confirmed by the stationarity of the series $\left(s_t^* - s_t\right)$, $\left(s_t - s_t^{\circ}\right)$, and $\left(s_t^* - s_t^{\circ}\right)$, documented in the bottom panels of Tables 1-4: the hypothesis of a unit root in our tests series can be rejected with confidence at the 5 percent significance level for all three exchange rates and all forecast horizons, and with even greater confidence in most cases.

The second set of results documented in Tables 1-4 pertains to the diagnostics of the model's rejection, namely, to evidence of violation of inequalities (10a) and (10b). The interesting evidence, in this case, is that in almost all cases, excess volatility is a main contributing cause of the rejection: actual spot rates are too volatile around s_t^*, around s_t°, or around both, for model (1) to capture our currencies' behavior in an efficient market. In

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[8]For this specification of s_t°, $\left(s_t - s_t^{\circ}\right)$ is independent of α, by construction.

particular, inequality (10a) tends to be rejected for lower values of α, while the converse is true for inequality (10b); at least one violation occurs in almost all cases.[9]

C. Alternative Specifications of the Test

To examine how our tests respond to changes in their specification, Table 5 reports tests performed for holding periods, h, of one, six, and twelve months and Table 6 reports tests performed for different definitions of the benchmark rate s_t°. (In each case, the other parameters are held fixed at their baseline values, with $\alpha=1$ and the survey forecasting horizon set at three months.)

Tests performed with holding periods of one and six months yield results qualitatively similar to those discussed for the case h=3 months: the joint assumptions of market efficiency and of applicability of model (1) to these markets can be rejected with confidence, and excess volatility of market rates around either fundamental rates or the selected benchmark is an underlying cause of rejection. The volatility tests statistics decline as h increases from one to twelve months, however. This feature likely reflects the loss of observations as h rises (by construction, a number h of observations is lost at the end of each sample), as well as the higher standard errors estimated for higher h (consecutive observations overlap information on fundamentals over intervals of length h-1, thus causing greater serial correlation). Furthermore, our confidence in the stationarity of the test series also declines with h. With a holding period of twelve months, in particular, our tests are invalid in the sample we examine, as we are unable to reject the hypothesis of a unit root in the series $\left(s_t^* - s_t\right)$ at the ten percent significance level.

Finally, to illustrate the flexibility of our approach with respect to the choice of s_t°, we report in Table 6 tests performed using alternative definitions of the benchmark rate, which was hitherto set at the "naive" value $s_t^{\circ} \equiv f_t$. We consider two alternative definitions of s_t°. First, we define s_t° as the model's prediction at time $t-1$ of the exchange rate expected to prevail at time t. This rate can be obtained by lagging equation (1) once and solving for $\mathrm{E}_{t-1}\left[s_t\right]$, yielding:

$$s_t^{\circ} \equiv \mathrm{E}_{t-1}\left[s_t\right] = \left(\frac{1+\alpha}{\alpha}\right) s_{t-1} - \frac{f_{t-1}}{\alpha}\,. \tag{13}$$

Alternatively, we also define $s_t^{\circ} \equiv s_{t-1}$, the last observed value of the market exchange rate. By this definition of s_t°, the volatility (or mean square error) of the difference $\left(s_t - s_t^{\circ}\right)$ corresponds to the familiar notion of conditional exchange rate volatility. Both these definitions meet the requirement that s_t° should be known at time t. As documented in the bottom panel of Table 6, both definitions also meet the requirement of making the series

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[9]Our test results proved robust to the choice of the sample period, as revealed by plots of recursively-generated sample average q's. The plots are available from the authors upon requests.

$(s_t^* - s_t)$, $(s_t - s_t^o)$, and $(s_t^* - s_t^o)$ stationary, leaving no doubt regarding our ability to reject the hypothesis of a unit root.

We report in Table 6 results for our baseline parameterization (α=1, h = 3, and survey forecast horizon equal to three months) when the benchmark rate is set at either one of these definitions. When $s_t^o \equiv E_{t-1}[s_t]$, model (1) is strongly rejected for all three exchange rates, although this time the tests reveal no evidence of excess volatility. By contrast, when $s_t^o \equiv s_{t-1}$, there is neither evidence against the model as a whole, nor—by necessity—evidence of excess volatility. One should not be surprised by these results, particularly by their apparent contrast with those reported in Tables 1-4: it is normal for our tests to yield different results when they are conditioned on different time-t information sets (that is, on different definitions of $(s_t - s_t^o)$). This is because, while some of these information sets may be correlated with future forecast errors if the market is inefficient, other sets may not be so. In particular, lack of evidence of excessive *conditional* volatility, and our inability to detect a significant correlation between the current innovation $(s_t - s_{t-1})$ and future forecast errors, reflects the well known difficulty (independent of market efficiency) of distinguishing exchange rate innovations from pure noise at short horizons.

IV. Concluding Remarks

For reasons ranging from calibration errors, small-sample biases, and—most importantly—strong reliance on specific exchange rate models, previous research on exchange rates has failed to provide reliable evidence of the ability of popular exchange rate models to match exchange rate volatilities observed over the post-Bretton Woods period. This paper has tried to partially address this gap by presenting volatility tests of the textbook asset-pricing model of the exchange rate that improve on previous tests in a number of ways. Most importantly, they do so by requiring a minimal set of assumptions on the intertemporal behavior of exchange rates and by allowing for a wide definition of their fundamental determinants, the observability of which is not required for the implementation of the tests. We applied these tests to data from the three major exchange rates and found broad evidence of excessive volatility with respect to the predictions of the canonical asset-pricing model of the exchange rate in an efficient market.

What do we make of this evidence? Although our approach to testing for excess volatility still involves a joint test of market efficiency and of a specific, if rather general, exchange rate model, it represents a considerable step forward with respect to previous research, given its reliance on a considerably smaller set of assumptions than previous studies. Weeding our list of (two) suspects further is, at this stage, essentially a matter of personal judgement. (We admit leaning towards the view that the problem is not so much that exchange rates are "too volatile," but rather that the canonical asset-pricing model of the exchange rate is unable to generate enough volatility.) Yet, our results may be useful in clarifying the agenda for future research on exchange rates. In reviewing empirical research on exchange rates conducted in recent years, for instance, Frankel and Rose (1995) note the sorry record of simple exchange rate models based on observable macroeconomic determinants (see also Flood and Taylor, 1996, for a similar assessment), and also note two promising lines of research: one line aimed at incorporating speculative bubbles into exchange rate models; and another line aimed at providing a detailed analysis of currency markets' micro-structure. Our analysis points to shortcomings of the simple asset-pricing view of the exchange rate even when unobservable fundamentals and speculative bubbles are accounted for. Thus, it suggests that if either lifeboat pointed by Frankel and Rose can rescue structural exchange rates models, it is unlikely to be the first one.

Table 1. Test Results: Forecast horizon = 1 month, holding period = 3 months

		$\alpha = 0.1$	$\alpha = 1$	$\alpha = 5$
Volatility Tests	sterling / dollar	-3.01 [a]	-2.92	-2.88 [b]
	mark / dollar	-2.63 [a]	-2.36	-2.31 [b]
	yen / dollar	-1.71 [a]	-1.62	-1.63 [b]
Z_t tests for $s_t^* - s_t$	sterling / dollar	-6.91	-6.40	-6.40
	mark / dollar	-6.85	-6.73	-6.75
	yen / dollar	-5.37	-5.18	-5.21
Z_t tests for $s_t - s_t^o$	sterling / dollar	-8.79	-8.79	-8.79
	mark / dollar	-7.30	-7.30	-7.30
	yen / dollar	-7.59	-7.59	-7.59
Z_t tests for $s_t^* - s_t^o$	sterling / dollar	-7.33	-8.69	-8.98
	mark / dollar	-6.96	-7.30	-7.39
	yen / dollar	-5.37	-7.30	-7.75

Table 2. Test Results: Forecast horizon = 3 months, holding period = 3 months

		$\alpha = 0.1$	$\alpha = 1$	$\alpha = 5$
Volatility Tests	sterling / dollar	-2.94 [a,b]	-2.30 [b]	-2.30 [b]
	mark / dollar	-3.43 [a,b]	-2.45 [b]	-2.38 [b]
	yen / dollar	-2.68 [a]	-1.48 [b]	-1.38 [b]
Z_t tests for $s_t^* - s_t$	sterling / dollar	-6.23	-4.84	-4.84
	mark / dollar	-5.66	-4.83	-4.98
	yen / dollar	-4.77	-3.60	-3.68
Z_t tests for $s_t - s_t^o$	sterling / dollar	-3.31	-3.31	-3.31
	mark / dollar	-2.99	-2.99	-2.99
	yen / dollar	-3.22	-3.22	-3.22
Z_t tests for $s_t^* - s_t^o$	sterling / dollar	-8.81	-5.49	-3.64
	mark / dollar	-8.26	-4.42	-3.12
	yen / dollar	-6.40	-4.89	-3.68

Notes: The values reported in the upper panels are the (asymptotically) standard normal test statistics for the null hypothesis that $E[q_t]=0$, where q_t is defined by equation (8) in the text. The superscripts "a" and "b" indicate that inequalities (10a) and (10b) were violated, respectively. The test statistics are written within single and double parentheses when any of the stationarity tests of the series $(s_t^* - s_t)$, $(s_t - s_t^o)$, and $(s_t^* - s_t^o)$ fails to reject the hypothesis of a unit root at the five percent and ten percent significance levels, respectively. Phillips-Perron Z_t stationarity tests (with critical values -2.91 and -2.59, respectively) are reported in the lower panels. The sample includes (173-h) observations from April 1984 to December 1998.

Table 3. Test Results: Forecast horizon = 6 months, holding period = 3 months

		α = 0.1	α = 1	α = 5
Volatility Tests	sterling / dollar	-2.94 [a,b]	-2.20 [a,b]	-2.10 [b]
	mark / dollar	-3.02 [a,b]	-2.49 [a,b]	-2.30 [b]
	yen / dollar	-3.71 [a,b]	-2.51 [a,b]	-2.31 [b]
Z_t tests for $s_t^* - s_t$	sterling / dollar	-5.63	-4.40	-4.66
	mark / dollar	-4.73	-4.29	-4.58
	yen / dollar	-4.60	-3.74	-3.97
Z_t tests for $s_t - s_t^o$	sterling / dollar	-3.67	-3.67	-3.67
	mark / dollar	-3.20	-3.20	-3.20
	yen / dollar	-3.77	-3.77	-3.77
Z_t tests for $s_t^* - s_t^o$	sterling / dollar	-9.67	-6.73	-4.26
	mark / dollar	-9.11	-5.40	-3.46
	yen / dollar	-7.39	-5.47	-4.43

Table 4. Test Results: Forecast horizon = 12 months, holding period = 3 months

		α = 0.1	α = 1	α = 5
Volatility Tests	sterling / dollar	-2.43 [a,b]	-1.78 [a,b]	-1.54 [b]
	mark / dollar	-2.84 [a,b]	-2.27 [a,b]	-2.03 [b]
	yen / dollar	-3.78 [a,b]	-2.89 [a,b]	-2.51 [b]
Z_t tests for $s_t^* - s_t$	sterling / dollar	-5.69	-4.26	-4.37
	mark / dollar	-4.66	-3.88	-4.32
	yen / dollar	-4.65	-3.66	-3.84
Z_t tests for $s_t - s_t^o$	sterling / dollar	-3.55	-3.55	-3.55
	mark / dollar	-3.29	-3.29	-3.29
	yen / dollar	-3.69	-3.69	-3.69
Z_t tests for $s_t^* - s_t^o$	sterling / dollar	-10.26	-7.96	-4.97
	mark / dollar	-9.68	-7.30	-4.36
	yen / dollar	-8.16	-5.66	-4.66

Notes: The values reported in the upper panels are the (asymptotically) standard normal test statistics for the null hypothesis that $E[q_t]=0$, where q_t is defined by equation (8) in the text. The superscripts "a" and "b" indicate that inequalities (10a) and (10b) were violated, respectively. The test statistics are written within single and double parentheses when any of the stationarity tests of the series $(s_t^* - s_t)$, $(s_t - s_t^o)$, and $(s_t^* - s_t^o)$ fails to reject the hypothesis of a unit root at the five percent and ten percent significance levels, respectively. Phillips-Perron Z_t stationarity tests (with critical values -2.91 and -2.59, respectively) are reported in the lower panels. The sample includes (173-h) observations from April 1984 to December 1998.

Table 5. Test Results: $\alpha = 1$, forecast horizon = 3 months

		$h = 1$ month	$h = 6$ months	$h = 12$ months
Volatility Tests	sterling / dollar	-3.50 [b]	-2.13 [a,b]	((-2.07 [a,b]))
	mark / dollar	-4.12 [b]	-2.27 [a,b]	((-2.24 [a,b]))
	yen / dollar	-2.36 [b]	(-2.43 [a,b])	((-2.34 [a,b]))
Z_t tests for $s_t^* - s_t$	sterling / dollar	-6.66	-3.28	-2.58
	mark / dollar	-5.43	-3.18	-2.58
	yen / dollar	-5.30	-2.78	-2.18
Z_t tests for $s_t - s_t^o$	sterling / dollar	-3.62	-3.26	-3.07
	mark / dollar	-3.07	-3.05	-2.98
	yen / dollar	-3.54	-3.32	-3.20
Z_t tests for $s_t^* - s_t^o$	sterling / dollar	-4.26	-6.36	-6.60
	mark / dollar	-3.05	-5.14	-4.99
	yen / dollar	-4.66	-5.10	-5.17

Table 6. Test Results: $\alpha = 1$, forecast horizon = 3 months, holding period = 3 months

		$s_t^o = \mathrm{E}_t[s_t]$	$s_t^o = s_{t-1}$
Volatility Tests	sterling / dollar	3.01	-0.92
	mark / dollar	2.95	-0.89
	yen / dollar	2.17	0.40
Z_t tests for $s_t^* - s_t$	sterling / dollar	-6.19	-6.19
	mark / dollar	-5.61	-5.61
	yen / dollar	-4.63	-4.63
Z_t tests for $s_t - s_t^o$	sterling / dollar	-8.31	-10.90
	mark / dollar	-7.37	-10.29
	yen / dollar	-8.18	-10.19
Z_t tests for $s_t^* - s_t^o$	sterling / dollar	-6.12	-8.35
	mark / dollar	-5.55	-7.83
	yen / dollar	-5.43	-7.06

Notes: The values reported in the upper panels are the (asymptotically) standard normal test statistics for the null hypothesis that $\mathrm{E}[q_t] = 0$, where q_t is defined by equation (8) in the text. The superscripts "a" and "b" indicate that inequalities (10a) and (10b) were violated, respectively. The test statistics are written within single and double parentheses when any of the stationarity tests of the series $(s_t^* - s_t)$, $(s_t - s_t^o)$, and $(s_t^* - s_t^o)$ fails to reject the hypothesis of a unit root at the five percent and ten percent significance levels, respectively. Phillips-Perron Z_t stationarity tests (with critical values -2.91 and -2.59, respectively) are reported in the lower panels. The sample includes (173-*h*) observations from April 1984 to December 1998.

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