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The Monetary Approach to Exchange Rate in an Efficient Foreign Exchange Market: Tests Based on Volatility

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ABSTRACT

The variance bounds on exchange rate movements implied by the monetary approach to exchange rate in an efficient foreign exchange market is shown to be violated by sample data. The paper also presents evidence showing that the forecast errors implied by the monetary model can be forecasted using historical data. The results are interpreted to suggest either the incompatibility of the monetary approach with sample data, or an inefficient foreign exchange market or both.

MODELS OF EXCHANGE RATE determination based on the monetary approach to the balance of payments imply, when coupled with an interest rate parity relation and the assumption of an efficient foreign exchange market, that exchange rates are related to a long average of money stocks differentials.[1, 2] The long average is analogous to the difference across countries of the long average of expected future money stocks which was supposed to characterize the price level of a single country (see, e.g., Sargent and Wallace [20]). Since long averages are known to have certain smoothing properties, it appears that the observed volatility of exchange rates may be inconsistent with these models. Thus, alternative tests of the models may be produced which rely on those properties. The aim of this paper is to determine if the observed volatility is consistent with those properties. The paper is the first to use tests based on the volatility of the exchange rate in an efficient market model of the monetary approach. The first set of tests is based on the property of implied variance bounds for exchange rate movements, and the second set consists of the more familiar efficient market tests of forecastability. The empirical results violate these two sets of tests of the model, thus casting doubt on the model.

Even though the claim that exchange rates are "too" volatile has been commonly made, few direct attempts have been made in the past to determine the validity of the claim. Some studies have presented estimates of variances of spot and forward rates but, lacking an objective criterion to determine what constitutes

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[1] The definition of market efficiency adopted here is one given by Fama [8]. This definition of market efficiency is the same as the definition of rational expectations given by Muth [18] when the latter is applied to the case of homogeneous expectations, and the terms market efficiency and rational expectations will be used interchangeably.

[2] Surveys of the literature have been done by Magee [16], Bilson [3, 4], Isard [13], Dornbusch [7], and Healy [12].

excess volatility, these studies have left the problem unresolved. To investigate this claim of excessive volatility, a benchmark which denotes "normal" volatility is necessary. The paper takes the volatility implied by market efficiency to be the norm. In adopting this criterion, we explicitly acknowledge that the claim of speculative excesses should be judged on the basis of ex ante anticipations rather than ex post occurrences. If the volatility of the exchange rates cannot be accounted for by changes in rational forecasts of future exchange rates, then the point can be made that the observed volatility in part reflect "speculative excesses."[3]

In the following section, I briefly describe the equilibrium condition studied in this paper. Section II presents the methodology and the empirical results. Tests of implied variance bounds for the exchange rates, and tests of equality restrictions on the cross covariance functions are presented in this section. The empirical analysis is confined to the period beginning March 1973, a date roughly coinciding with the beginning of the generally floating exchange rate regime, and ending March 1979. Three sets of monthly data were examined: the dollar price of *D*-mark, the dollar price of the pound, and the pound price of the *D*-mark.[4, 5] The choice of the countries was dictated by the presence in these countries of well-developed and integrated financial markets, and in particular the foreign exchange markets. The choice was also determined by the fact that these pairs of countries are the usual subjects of inquiry for past studies of monetary approach to exchange rate determination with rational expectations. Section III concludes the paper.

I. The Model

The monetary approach to exchange rate with rational expectations can be stated as[6]

$$s_t = x_t + \alpha[E_t(s_{t+1}) - s_t], \tag{1}$$

where $x_t = (m_t - m_t^*) - \beta(y_t - y_t^*)$, m_t, y_t, and s_t the logarithms of nominal money stock, real income, and spot exchange rates quoted in terms of home currency per unit of foreign currency respectively.[7] β is the elasticity of demand

[3] In most studies concerned with the volatility of exchange rate, the distinction between stabilizing and destabilizing speculation is often discussed. As has recently been emphasized by Bilson [4] and Kohlhagen [14], before it is possible to determine empirically if speculation has been stabilizing, the assumption of an underlying model that determines the exchange rate is necessary. Since stability is defined with respect to an equilibrium condition, by definition, an efficient market is one where speculation is stabilizing with respect to an equilibrium exchange rate. Thus, tests conducted for the present study can be used to determine if speculation has been stabilizing with respect to the monetary approach to exchange rate determination with rational expectations.

[4] It can be legitimately argued that it would be more interesting to look at daily price variance since market participants are more concerned with daily time horizons. However, given the availability of data, monthly sampling interval will have to suffice.

[5] All the data are pulled from the International Financial Statistics tapes. Money supply figures are from line number 34..b, and real production indices are from line number 66..c.

[6] See the articles referenced by the surveys mentioned in Footnote 2.

[7] The use of logarithms for the spot rate makes the results insensitive to the choice of the numeraire and circumvents what is known as Siegel's paradox (see Siegel [24] and Fleming, Turnovsky, and Kemp [9]).

for money with respect to income and α is the semielasticity of the demand for money with respect to nominal interest rates. The asterisk denotes the foreign country. The variable x_t can be interpreted as the nominal money balances relative to income where income is adjusted for income elasticity demand for money. E_t is the conditional expectations operator defined by $E_t(s_{t+1}) \equiv E(s_{t+1} \mid I_t)$ where E is the expectations operator and I_t is the information set used by the market participants at time t so that $I_t \supset I_{t-1} \supset I_{t-2} \cdots$.

Equation (1) is a monetary approach in that x_t describes the money market parameters which determine both interest rates and prices, and which in turn determine the exchange rate.

Equation (1), however, could not be directly tested without an additional assumption, since $E_t(s_{t+1})$ is an unobservable variable. But the assumption of an efficient market means that expectations are formed rationally, so that

$$E_t(s_{t+\iota}) = \frac{1}{1+\alpha}\{E_t(x_{t+\iota}) + \alpha E_t[E_{t+\iota}(s_{t+1+\iota})]\} \qquad i \geq 1.$$

But, by the law of interated projections, we can write

$$E_t[E_{t+\iota}(s_{t+\iota+1})] = E_t(s_{t+\iota+1}),$$

so that by repeating the procedure iteratively, we derive

$$s_t = (1-\gamma)\sum_{\iota=0}^{\infty} \gamma^{\iota} E_t(x_{t+\iota}), \tag{2}$$

where $\gamma = \dfrac{\alpha}{1+\alpha}$, and the convergence assumption that $\lim_{\iota\to\infty}\left(\dfrac{\alpha}{1+\alpha}\right)^{\iota} E_t(s_{t+\iota}) = 0$ is made here. Upon rearranging, Equation (2) can be written as:

$$(s_t - x_t) = (1-\gamma)\sum_{\iota=1}^{\infty}\sum_{j=\iota}^{\infty} \gamma^{j} E_t(Dx_{t+\iota}) = E_t(A_t), \tag{3}$$

where $A_t = \sum_{\iota=1}^{\infty} \gamma^{\iota}(Dx_{t+\iota})$, and D is the difference operator, defined by the operation $Dx_t \equiv x_t - x_{t-1}$. Therefore, A_t is the value of $(s_t - x_t)$ when the actual subsequent values of Dx_t's are known with certainty at time $t + \infty$. As Mussa [17], Dornbusch [7], and others have noted, Equation (3) states that exchange rates are determined by expectations and the movements can be attributed to changes in the current and expected future values of x_t's which move in response to new information shocks.

II. Methodology and Empirical Results

From a technical standpoint, the tests used in this paper are along the lines which Shiller used to study the volatility of long-term bonds [22], and the volatility of stock prices [23]. Papers by LeRoy and Porter [15] and Singleton [25] have also addressed the volatility issue (for stock prices and bond prices respectively).

A. *Tests of Implied Variance Bounds I*

Equation (3) can be stated as:

$$(s_t - x_t) = A_t - u_t, \tag{4}$$

where $u_t = A_t - E_t(A_t)$ is the realized forecast error that is uncorrelated with currently available information, information which does not include A_t. Therefore,

$$\mathrm{Var}(A_t) = \mathrm{Var}(s_t - x_t) + \mathrm{Var}(u_t) \geq \mathrm{Var}(s_t - x_t). \tag{5}$$

But since A_t is a moving average, it tends to smooth the individual Dx_{t+i}'s, and therefore Dx_t has a higher variance than A_t. More precisely, since A_t is determined by passing Dx_t through a linear filter, it is well known that if the coefficients in (3) sum to one, the squared gain of the linear filter is $g^2(\omega) = (1-\gamma)^2/[1+\gamma^2 - 2\gamma\,\mathrm{Cos}(\omega)]$ where ω is the frequency which lies between zero and π. The squared gain declines monotonically as frequency increases with a maximum at $\omega = 0$ of one. It would then follow immediately that the spectrum of A_t is more concentrated in the lower frequencies than is the spectrum of Dx_t, and the variance of A_t is less than that of Dx_t. Consequently, adjusting for the fact that the summation in (3) does not sum to one, since it starts from $i = 1$, we have:

$$\left(\frac{\gamma}{1-\gamma}\right)^2 \mathrm{Var}(Dx_t) \geq \mathrm{Var}(A_t) \geq \mathrm{Var}(s_t - x_t).$$

The results of the test are striking enough that only the results of the tests based on the observable quantities, $(s_t - x_t)$ and Dx_t are presented below.[8] Thus, our first inequality test is:

$$\mathrm{Var}(s_t - x_t) \leq \left(\frac{\gamma}{1-\gamma}\right)^2 \mathrm{Var}(Dx_t). \tag{I1}$$

Before reporting the results of the inequality test (I1), we note the assumption necessary in carrying out the test. The variance tests of this paper have a distinct advantage over previous tests of (3) (or the equivalent expression (2)), in that previous tests usually consist of regressions which require specification of movements of money stock and real income overtime, which as a practical matter, necessarily require ad hoc assumptions. On the other hand, tests of variance inequalities such as (I1) require no such specification since it is not information-specific in that no element of the information set needs be specified. The variance tests here require only the assumption of stationarity. No fail-proof test of stationarity exists, but following Box and Jenkins [5], we can determine by examining the first few sample autocorrelations of the time series, the minimal degree of differencing necessary to induce stationarity. It is found that the sample autocorrelations of Dx_t have the usual pattern of a stationary time series.[9] More precisely, we assume that $\{Dx_t\}$ and $\{I_t\}$, the information set, are joint covariance stationary. This means that $\{Dx_t\}$ and $\{I_t\}$ are each stationary, and the finite $\mathrm{Cov}(Dx_t, I_{t+i})$ is only a function of i.[10, 11]

[8] A_t can be approximated by observing that $A_t = \gamma A_{t+1} + \gamma Dx_{t+1}$, so that one can start by assuming a terminal value for A_{t+1}, and then recursively working backward. Since $0 < \gamma < 1$, the importance of the terminal value chosen declines as one moves backward in time.

[9] The sample autocorrelations are available from the author upon request.

[10] It is to be understood that expressions such as $\mathrm{Var}(Dx_t)$ even though written with a time subscript, are not functions of time. Note that the existence of the entire autocovariance function of Dx_t is implied by the existence of a finite $\mathrm{Var}(Dx_t)$ as can be seen by applying Cauchy-Schwartz's inequality. In addition, since $\mathrm{Var}(Dx_t) = E[\mathrm{Var}(Dx_t \mid I_t)] + \mathrm{Var}[E_t(Dx_t)]$, therefore $\mathrm{Var}[E_t(Dx_t)]$ must also be finite.

[11] The reason for not taking the first difference of (2) in order to induce stationarity, and then

The results of the sample variances of (I1) are shown in Table I. In computing the variances, we need to specify the values of $\gamma = \frac{\alpha}{1+\alpha}$ and β, the income elasticity. A casual examination of (I1) shows that the right-hand side of the inequality increases monotonically with the value of α. The monetary approach is usually assumed, based on past money demand studies, to be consistent with the view that α lies between 0 and 3 with 95 percent confidence limits.[12] Consequently, in order to make the strongest possible case against the model (3), the results of the inequalities for $\alpha = 3$ are reported. The choice of $\alpha = 3$ for the two other inequality tests derived below are made for the same reason. In fact the results in Table I for (I1) show that the value of α would have to be implausibly high for the inequality not to be violated. Conversely, since interest rate elasticity is equal to the product of α and the interest rate, for (I1) not to be violated, interest rates during the sample period would have to be unrealistically low, or the interest rate elasticities would have to be unrealistically high. As for the value of β, the literature on the monetary approach generally assumed that, based on past studies of money demand, it lies between 0.5 and 1.5 with 95 percent confidence limits. The results in Table I are reported for $\beta = 0.5$, $\beta = 1.0$, and $\beta = 1.5$. The conclusions remain unchanged for other values of β within the interval. The test results of (I1) thus indicate that the magnification of the right-hand side of (I1) given by $\left(\frac{\gamma}{1-\gamma}\right)^2 = 9$ for $\alpha = 3$ is insufficient for the inequality to hold.

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testing the resulting model instead of testing the model as formulated by (3), is that, contrary to intuition, a variance inequality for levels does not imply a variance inequality for changes, so that testing the variance bounds implied by the model obtained by taking the first difference of (2) is not equivalent to testing the variance bounds of (2). To show this, for expositional simplicity, assume the following equilibrium conditions:

$$y_t = E_t[x_{t+1}]$$

$$x_{t+1} = y_t + \eta_t$$

Therefore, if the market is efficient, we would expect

$$\mathrm{Var}(x_{t+1}) \geq \mathrm{Var}(y_t) \tag{F1}$$

But it does not follow that

$$\mathrm{Var}(Dx_{t+1}) \geq \mathrm{Var}(Dy_t) \tag{F2}$$

This is so because

$$\begin{aligned}\mathrm{Var}(Dx_{t+1}) &= 2\,\mathrm{Var}(y_t) + 2\,\mathrm{Var}(\eta_t) - 2\,\mathrm{Cov}(y_t, y_{t-1}) - 2\,\mathrm{Cov}(y_t, \eta_{t-1}) \\ &= \mathrm{Var}(Dy_t) + 2\,\mathrm{Var}(\eta_t) - 2\,\mathrm{Cov}(y_t, \eta_{t-1}).\end{aligned}$$

Therefore, (F2) would be true if and only if

$$\mathrm{Var}(\eta_t) \geq \mathrm{Cov}(y_t, \eta_{t-1}). \tag{F3}$$

A model which violates this is

$$y_t = k\eta_{t-1}$$

where k = constant > 1. The covariance term in (F3) is what Shiller [21] referred to as the cause of fundamental asymmetry in rational expectations models.

[12] See e.g., Bilson [1, 2].

Table I[a]

Inequality Tests With $\alpha = 3$

Date Set	β	I1 LHS	I1 RHS	I2 LHS	I2 RHS	I3 LHS	I3 RHS
United States–United Kingdom	0.5	0.0141	0.00359	0.00068	0.00100	0.00198	0.00091
		(0.603)*				(0.370)*	
	1.0	0.0182	0.00617	0.00068	0.00172	0.00230	0.00157
		(0.501)*				(0.189)	
	1.5	0.0239	0.01071	0.00068	0.00297	0.00279	0.00272
		(0.386)*				(0.013)	
United States–Germany	0.5	0.0399	0.00261	0.00122	0.00072	0.00562	0.00066
		(0.382)*		(0.258)**		(0.794)*	
	1.0	0.0461	0.00441	0.00122	0.00123	0.00627	0.00112
		(0.831)*				(0.700)*	
	1.5	0.0531	0.00757	0.00122	0.00210	0.00700	0.00192
		(0.757)*				(0.571)*	
United Kingdom–Germany	0.5	0.0162	0.00407	0.00088	0.00113	0.00337	0.00103
		(0.603)*				(0.532)*	
	1.0	0.0148	0.00758	0.00088	0.00210	0.00325	0.00192
		(0.329)*				(0.258)**	
	1.5	0.0138	0.01322	0.00088	0.00367	0.00319	0.00336
		(0.022)					

[a] LHS stands for left-hand side of the inequality and RHS stands for the right-hand side of the inequality. The number of observations for each data set equals 72. The numbers in parentheses give the values of Pitman statistics, which are calculated as $P(y_1, y_2) = [\mathrm{Var}(y_1) - \mathrm{Var}(y_2)]/\{[\mathrm{Var}(y_1) + \mathrm{Var}(y_2)]^2 - 4\rho^2_{y_1,y_2}\mathrm{Var}(y_1)\mathrm{Var}(y_2)\}^{1/2}$ where ρ_{y_1,y_2} is the correlation coefficient between y_1 and y_2. Pitman statistic is used in place of an F-statistic when y_1 and y_2 are correlated (see Snedecor and Cochran [26]). *denotes significance at the 1 percent level, and **significance at the 5 percent level.

B. *Tests of Implied Variance Bounds II*

Since (I1) is so dramatically violated, it is of interest to know if inequalities based on other smoothing properties of (3) are violated. In particular, we focus on the different characterizations of the behavior of exchange rates which have been issues of considerations in the literature on the monetary approach. These characterizations concern the volatility of changes in spot exchange rates, and the movement in s_t due to arrival of new information between time $t-1$ and t or the volatility of the innovation in s_t.

In order to derive these further inequalities tests, we define an innovation operator $\Delta_t \equiv E_t - E_{t-1}$ which captures the change in expectation due to arrival of new information between time t and $t-1$. Hence

$$\Delta_{t+1} s_{t+1} = (s_{t+1} - s_t) - \mathrm{E}_t(s_{t+1} - s_t), \tag{6}$$

which upon substitution into (1) yields

$$(s_t - x_t) = \alpha(Ds_{t+1} - \Delta_{t+1}s_{t+1})$$

or

$$\Delta_{t+1}s_{t+1} = Ds_{t+1} - \frac{1}{\alpha}(s_t - x_t). \tag{7}$$

Therefore, given the model, the innovation in the spot rate is an observable variable. As Samuelson [19] has observed, $E_t\Delta_{t+n} = 0$ for $n \geq 1$; therefore, Δ_{t+n} for $n \geq 1$ must be uncorrelated with all the information known at time t. In particular, in deriving the following two inequalities, we will use the assumption

$$\mathrm{Cov}(\Delta_{t+1}s_{t+1},\ s_t - x_t) = 0. \tag{8}$$

In order to derive the upper bound for $\mathrm{Var}(Ds_t)$ in terms of $\mathrm{Var}(Dx_t)$, we substitute (8) into $\mathrm{Var}(Ds_{t+1})$, and then maximize with respect to $\mathrm{Var}(s_{t+1} - x_{t+1})$ for given $\mathrm{Var}(Dx_{t+1})$ and $\rho((s_{t+1} - x_{t+1}), Dx_{t+1})$. The details of this derivation and the derivation for the next inequality are available from the author upon request. The resulting inequality is:

$$\mathrm{Var}(Ds_t) \leq \left(\frac{1}{2}\right)\left(\frac{2-\gamma}{1-\gamma}\right)\mathrm{Var}(Dx_t). \tag{I2}$$

We can likewise derive an upper bound for the variance of the innovation in terms of $\mathrm{Var}(Dx_t)$. Analogous to the derivation above, we can derive the maximum fluctuations $\Delta_{t+1}s_{t+1}$, can experience for a given $\mathrm{Var}(Dx_{t+1})$ by substituting (8) into $\mathrm{Var}(\Delta_{t+1}s_{t+1})$, and then maximizing with respect to $\mathrm{Var}(s_{t+1} - x_{t+1})$ holdin $\mathrm{Var}(Dx_{t+1})$ and $\rho((s_{t+1} - x_{t+1}), Dx_{t+1})$ constant. This leads to:

$$\mathrm{Var}(\Delta_t s_t) \leq \frac{1}{1-\gamma^2}\mathrm{Var}(Dx_t). \tag{I3}$$

The test results of (I2) and (I3) are shown in Table I. In general, inequality (I2) is not violated, but (I3) is. It appears that in the case of (I2), the monetary approach implies a magnification of the $\mathrm{Var}(Dx_t)$ which is big enough so that it is not exceeded by the $\mathrm{Var}(Ds_t)$. On the other hand, there is so much variability in the arrival of "new information" in the market that the magnification implied by the monetary approach of $\mathrm{Var}(Dx_t)$ for (I3) is exceeded by $\mathrm{Var}(\Delta_t s_t)$. Since the three inequalities are different characterizations of the same model, violations of any one of them is sufficient to reject the model.

One can tie the variances on the left-hand side of the three inequalities by observing that (7) implies:

$$\mathrm{Var}(Ds_{t+1}) = \mathrm{Var}(\Delta_{t+1}s_{t+1}) + \left(\frac{1}{\alpha}\right)^2\mathrm{Var}(s_t - x_t). \tag{9}$$

A casual examination of Table I shows that (9) does not hold so that (8) is contradicted by sample data. In the following subsection, we use the more familiar efficient market tests of forecastability for which the power of the tests are different from the variance inequality tests above, to examine our model, and in particular (8).

C. *Tests of Equality Restrictions of the Cross Covariance Functions*

It might appear that (4) can be directly tested by regressing A_t on $(s_t - x_t)$ as in familiar tests of market efficiency which involve regressing actual values on the forecasts:

$$A_t = a + b(s_t - x_t) + u_t \tag{10}$$

with the null hypothesis of $b = 1$ against the alternative hypothesis of $b \neq 1$.[13]

[13] The restriction on the constant a, depends on the assumption about the distribution of the error term and therefore, needs not be zero.

Two observations can be made here with regard to (10). First, since (4) implies (5) but not vice versa, rejection of the null hypothesis by the regression tests does not imply the violation of (5). In other words, regression tests are not powerful tests against the alternative hypothesis of $\mathrm{Var}(A_t) \le \mathrm{Var}(s_t - x_t)$. As Geweke and Feige [11] have recently noted, the specification of the alternative hypothesis is more difficult than specification of the null hypothesis of market efficiency. Thus, the regression tests of this subsection, and the tests of implied variance bounds of the previous subsections, offer two different reasons for rejection of the efficient market model as alternative hypotheses. Second, in the specification (10), the u_t's are not serially uncorrelated since A_t is not known at time t, and consequently ordinary significance tests cannot be used. As Shiller [22] has noted, a generalized least squares approach to this problem is to regress $(u_t - \gamma u_{t+1})$ on all the information known at time t. But $(u_t - \gamma u_{t+1}) = \gamma[Dx_{t+1} - \frac{1}{\gamma}(s_t - x_t) + (s_{t+1} - x_{t+1})] = \gamma[\Delta_{t+1}s_{t+1}]$, and therefore is tantamount to testing if the innovation in s_{t+1} can be forecasted using information known at time t. Equation (8) is an example of this restriction, and is the one used to derive both (I2) and (I3). In general, one can test the hypothesis that the innovations in the spot rate at time $t + 1$ are orthogonal to any information known at time t. However, in order to make the strongest possible case against the model, elements of the information set which are more relevant and readily available should be used, since it is not feasible to test the entire information set with finite data. In light of this consideration and the focus of this paper, I confine my regressions to the projections of $\Delta_{t+1}s_{t+1}$ on $s_t - d_t$, and of $\Delta_{t+1}s_{t+1}$ on $\Delta_t s_t$. The latter is the test for serial correlation in the innovations usually performed in market efficiency studies, and the former is a test of (8). It should be noted here that these regressions are different from the projections of Ds_{t+1} on information known at time t as have often been studied in that $\Delta_{t+1}s_{t+1}$ is not the same as Ds_{t+1} if (1) is the model assumed.[14]

If the foreign exchange market is efficient and the monetary approach provides the correct model, then we would expect the regressions of $\Delta_{t+1}s_{t+1}$ on $(s_t - x_t)$ to produce a coefficient insignificantly different from zero or more generally, a low R^2. The regression results are reported in Table II. All the regressions show coefficients which are very highly significantly different from zero with high R^2's. In fact, the signs and magnitudes of the coefficients for the first two data sets indicate that projecting $\Delta_{t+1}s_{t+1}$, as given by (7), on $(s_t - x_t)$ is only capturing the correlation of $-\frac{1}{\alpha}(s_t - x_t)$ on $(s_t - x_t)$, where $-\frac{1}{\alpha} = -\frac{1}{3}$ in the regressions. The results indicate that a martingale model, $E_t(s_{t+1}) = s_t$, at least with respect to an information set containing $(s_t - x_t)$, would describe the data quite well for the first two data sets. This conjecture is supported empirically in Table III where we found, as expected, that the regression of Ds_{t+1} on $(s_t - x_t)$ produce coefficients insignificantly different from zero for the first two data sets. The third data set produce coefficients which are significant but of negligible magnitude. If the null hypothesis of an efficient market is correct, the results in Table II should not be

[14] In a recent paper, Cumby and Obstfeld [6] note a similar distinction between Ds_t and $\Delta_{t+1}s_{t+1}$, but in the context of an open Fisher model, and then proceed to examine the serial correlation of the innovations.

biased due to the presence of $(s_t - x_t)$ on both sides of the equation. Likewise, serial correlation tests for the innovations reported in Table IV show coefficients which are significantly different from zero. The results of tests of implied variance bounds and regression tests of forecastability thus cast doubt on the consistency of the rational expectations version of the monetary approach to exchange rate with sample data.

Table II

Regression Test: $(\Delta_{t+1}s_{t+1}) = a + b\,(s_t - x_t) + \epsilon_t$ With $\alpha = 3$

Data Set	β	$\hat{a}$	$\hat{b}$	R^2	DW
United States	0.5	0.021	−0.322	0.655	1.80
		(0.363)[a]	(−11.53)		
	1.0	0.035	−0.315	0.705	1.83
		(0.686)	(−12.92)		
United Kingdom	1.5	0.039	−0.313	0.758	1.84
		(0.911)	(−14.80)		
United States	0.5	−0.004	−0.341	0.784	2.02
		(−0.144)	(−15.92)		
	1.0	−0.003	−0.339	0.806	2.02
		(−0.091)	(−17.05)		
Germany	1.5	−0.001	−0.338	0.826	2.02
		(−0.041)	(−18.24)		
United Kingdom	0.5	0.045	−0.389	0.753	1.67
		(2.50)	(−14.61)		
	1.0	0.052	−0.399	0.749	1.66
		(2.79)	(−14.45)		
Germany	1.5	0.057	−0.409	0.749	1.65
		(3.09)	(−14.45)		

[a] The numbers in parentheses are t-statistics. The number of observations in each data set equals 72.

Table III

Regression Test: $(Ds_{t+1}) = a + b\,(s_t - x_t) + \epsilon_t$

Data Set	β	$\hat{a}$	$\hat{b}$	R^2	DW
United States	0.5	0.0213	0.011	0.002	1.80
		(0.363)[a]	(0.407)		
	1.0	0.035	0.018	0.008	1.83
		(0.686)	(0.737)		
United Kingdom	1.5	0.039	0.021	0.013	1.84
		(0.911)	(0.972)		
United States	0.5	−0.004	−0.007	0.002	2.02
		(−0.144)	(−0.335)		
	1.0	−0.003	−0.006	0.001	2.02
		(−0.091)	(−0.299)		
Germany	1.5	−0.001	−0.005	0.001	2.02
		(−0.041)	(−0.267)		
United Kingdom	0.5	0.045	−0.055	0.058	1.67
		(2.50)	(−2.08)		
	1.0	0.052	−0.066	0.075	1.66
		(2.79)	(−2.38)		
Germany	1.5	0.057	−0.076	0.093	1.65
		(3.09)	(−2.69)		

[a] The numbers in parentheses are t-statistics. The number of observations in each data set equals 72.

Table IV

Regression Test: $(\Delta_{t+1}s_{t+1}) = a + b\,(\Delta_t s_t) + \epsilon_t$ With $\alpha = 3$

Data Set	β	$\hat{a}$	$\hat{b}$	R^2	DW
United States	0.5	0.314	0.548	0.290	2.33
		(4.35)[a]	(5.31)		
	1.0	0.267	0.609	0.356	2.38
		(3.93)	(6.17)		
United Kingdom	1.5	0.220	0.673	0.433	2.43
		(3.49)	(7.26)		
United States	0.5	0.145	0.693	0.456	2.56
		(3.26)	(7.60)		
	1.0	0.129	0.721	0.492	2.60
		(3.03)	(8.18)		
Germany	1.5	0.115	0.747	0.525	2.63
		(2.82)	(8.74)		
United Kingdom	0.5	−0.075	0.658	0.426	2.24
		(−3.69)	(7.16)		
	1.0	−0.081	0.623	0.380	2.19
		(−3.90)	(6.50)		
Germany	1.5	−0.087	0.588	0.336	2.14
		(−4.10)	(5.91)		

[a] The numbers in parentheses are t-statistics. The number of observations in each data set equals 71.

III. Conclusions

In his well-known defense of flexible exchange rates, Professor Friedman states that to consider flexible exchange rates to be unstable is to confuse "the symptom of difficulties with the difficulties themselves. A flexible exchange rate need not be an unstable exchange rate. If it is, it is primarily because there is underlying instability in the economic conditions governing international trade."[15] The tests in this paper can be regarded as an empirical analysis of this statement in the framework of an underlying economic structure denoted by a monetary model in an efficient market, and as such found evidence which shows that exchange rates are too volatile to be consistent with the monetary model, and/or an efficient market.

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[15] Friedman [10], p. 173.

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