American Economic Association

A Test for Speculative Bubbles in the Sterling-Dollar Exchange Rate: 1981-84
Author(s): George W. Evans
Source: *The American Economic Review*, Vol. 76, No. 4 (Sep., 1986), pp. 621-636
Published by: American Economic Association
Stable URL: http://www.jstor.org/stable/1806063
Accessed: 20-01-2016 22:42 UTC

Your use of the JSTOR archive indicates your acceptance of the Terms & Conditions of Use, available at http://www.jstor.org/page/info/about/policies/terms.jsp

JSTOR is a not-for-profit service that helps scholars, researchers, and students discover, use, and build upon a wide range of content in a trusted digital archive. We use information technology and tools to increase productivity and facilitate new forms of scholarship. For more information about JSTOR, please contact support@jstor.org.

American Economic Association is collaborating with JSTOR to digitize, preserve and extend access to *The American Economic Review.*

http://www.jstor.org

This content downloaded from 165.123.34.86 on Wed, 20 Jan 2016 22:42:21 UTC
All use subject to JSTOR Terms and Conditions

A Test for Speculative Bubbles in the Sterling-Dollar Exchange Rate: 1981–84

By GEORGE W. EVANS*

The U.S. dollar price of the U.K. pound sterling is tested for a speculative bubble, defined as a period with a nonzero median in excess returns. A nonparametric procedure is developed which controls for data mining over the period of flexible exchange rates and finds a negative bubble in the excess return to holding sterling rather than dollar assets during 1981–84. Possible interpretations are bootstrap equilibria (rational bubbles), asymmetric fundamentals, and nonrational expectations.

Many observers have felt that the foreign exchange markets have been too volatile since the advent of flexible rates, subject to large and persistent speculative movements in exchange rates over substantial periods of time.[1] A salient example of this is the sustained decrease of the U.S. dollar price of the U.K. pound sterling from December 1980 through February 1985, a decline which, toward the latter part of the period, led to a belief among many economists and policymakers in both countries that the pound was undervalued relative to the dollar.[2]

Figure 1 indicates the magnitude of the swing: during this four-year period, the dollar price of the pound fell by over one-half. The extent of this slide cannot be explained by differential interest rates or inflation rates between the two countries. For example, correcting for interest rate differentials, the average excess return to holding dollars rather than sterling over this period was over 19 percent per annum.

The size and persistence of the negative excess returns to holding sterling might appear to constitute decisive evidence of a "speculative bubble" or "abnormality" in the dollar-pound exchange rate during this period. This conclusion, however, requires

*Department of Economics, Stanford University, Stanford, CA 94305-1992. This research was begun while I was visiting the London School of Economics, and was supported by the Center for Economic Policy Research, Stanford. I am indebted to Pauline Andrews for developing the computer programs required for the tests, to Simona Hughes for collecting the data, and to Roger Alford for providing the DIP program used to plot Figure 1. I also thank the participants in the Stanford Macroeconomics Seminar and the managing editor of this journal for their comments. An earlier version of this paper was circulated under the title "Speculative Bubbles and the Sterling-Dollar Exchange Rate: A New Test."

[1]See, for example, Ronald McKinnon (1979, pp. 155–56). There was a related discussion, antedating the 1971–73 shift to flexible exchange rates, over destabilizing and stabilizing speculation under freely floating exchange rates. References can be found in McKinnon.

[2]McKinnon (1984, ch. 5) argued in March 1984 that the dollar was overvalued and the *Economist* (January 5, 1985, p. 54) reported that an OECD study based on purchasing power parity calculations showed the pound substantially undervalued against the dollar in mid-1984. In the United States, the Treasury in 1983 described coordinated "intervention to stem the rise of the dollar" (*Wall Street Journal*, August 2, 1983). In July 1984, Paul Volcker, the Chairman of the Federal Reserve Board, discussed the risk of a fall in the dollar (*Financial Times*, July 26, 1984) and in early 1985 referred to an overvalued dollar, the need for concerted intervention, and the possibility of a "very sharp decline" in the dollar (*Economist*, March 16, 1985, p. 73 and *Financial Times*, March 7, 1985). In the United Kingdom, interest rates were increased in July 1984 "in the face of a further speculative run against the pound" (*Financial Times*, July 7, 9, and 12, 1984). Then in January 1985 the British government, which was "extremely concerned about the recent slide in sterling on the foreign exchanges," engineered a total four and one-half percentage point increase in bank base rates in just over two weeks in an effort to arrest the decline in the pound sterling (*Financial Times*, January 14, 15, and 29, 1985). Intervention in the foreign exchange markets by the central banks of the G5 countries occurred on a number of occasions (for example, see *Financial Times*, July 4, 1984, and March 22, 1985).

This content downloaded from 165.123.34.86 on Wed, 20 Jan 2016 22:42:21 UTC
All use subject to JSTOR Terms and Conditions

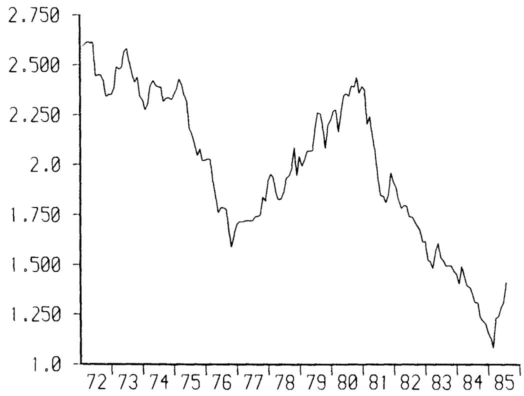

FIGURE 1. U.S. DOLLAR PRICE OF U.K. POUND STERLING

careful statistical support since, for example, an asset price following a random walk will on occasion, purely by chance, exhibit such apparent abnormalities over subperiods. A sound statistical test must therefore take account of data mining. The focus of this paper is the design of a procedure for detecting a precisely defined type of speculative bubble or abnormality in the dollar price of sterling over 1981–84, which protects against data mining from within the flexible exchange rate period.

In apparent contrast to the view that the foreign exchange markets contain periods of bubbles or abnormalities in excess returns, it has been widely assumed in the literature that, under flexible rates, the foreign exchange markets set prices efficiently. If agents are risk neutral, expectations are formed rationally, transactions costs can be neglected and markets are competitive, then "simple efficiency" holds and there should be no expected profit opportunities in the forward exchange markets. This is usually expressed as

$$(1) \qquad E_t x_{t+1} = 0 \quad \text{for} \quad x_{t+1} = s_{t+1} - f_t,$$

where s_{t+1} is the logarithm of the spot price of foreign currency in period $t+1$ and f_t is the logarithm of the one-period forward rate, that is, the price at time t of forward currency to be delivered at $t+1$. Here E_t denotes the mathematical expectation conditioned on information available at time t. Equation (1) is also referred to as the "unbiased expectations" hypothesis since it is equivalent to $f_t = E_t s_{t+1}$.

Equation (1) can be tested empirically in several ways, for example, by regressing x_t, the excess rate of return to holding a forward exchange contract, on variables in the information set and testing the null hypothesis of zero coefficients, or by testing the implied nonlinear restrictions imposed on vector autoregressive models. Although some tests have appeared to support simple efficiency, recent tests using larger sample sizes seem to have clearly detected small but statistically significant deviations from (1).[3] However, the dominant interpretation of these findings is that they are due to time-varying risk premia. Thus it is argued that $E_t \tilde{x}_{t+1} = 0$ for a suitably defined risk-adjusted excess return $\tilde{x}_{t+1}$, so that efficiency holds when allowance is made for risk-averse agents. We return to this issue below in Sections I, Part B, and II. Assuming that (1) holds, or a suitable modification of (1) to allow for risk premia,[4] can this be reconciled with the view of a market subject to large speculative movements? There are several possibilities.

Since to a close approximation the markets obey covered interest arbitrage, we have $f_t = s_t + i_{t,1} - i^*_{t,1}$ where $i_{t,1}$ and $i^*_{t,1}$ are the domestic and foreign one-period interest rates from period t to $t+1$, known at time t, on comparable assets such as Eurocurrency deposits. Then x_{t+1} is also given by

$$(2) \qquad x_{t+1} = s_{t+1} - s_t + i^*_{t,1} - i_{t,1},$$

[3]A partial list of the large and expanding literature on tests of simple efficiency includes Jacob Frenkel (1977, 1981), John Geweke and Edgar Feige (1979), Lars Peter Hansen and Robert Hodrick (1980), Craig Hakkio (1981), Sebastian Edwards (1983), and Richard Baillie, Robert Lippens and Patrick McMahon (1983), as well as several of the references cited later in this paper.

[4]To simplify discussion of the other issues involved, I ignore risk premia throughout the remainder of the introduction.

This content downloaded from 165.123.34.86 on Wed, 20 Jan 2016 22:42:21 UTC
All use subject to JSTOR Terms and Conditions

and (1) may be equivalently written as

$$i_{t,1} = i^*_{t,1} + E_t s_{t+1} - s_t, \tag{3}$$

which is sometimes known as the open interest parity condition. Note that x_{t+1} can be interpreted as the excess rate of return, in domestic currency, from holding an uncovered foreign interest-bearing asset rather than a comparable domestic asset. It was shown in Rudiger Dornbusch (1976) that when (3) is incorporated in models with predetermined goods prices, it is possible for the exchange rate, in response to new information about fundamentals, to overshoot its long-run equilibrium value and thence to follow a systematic path back toward equilibrium. However, the predictable component of the movement in s_t would be entirely captured in interest rate differentials, which we net out when computing excess returns.

Another possible explanation of large systematic variations in exchange rates is that the rate follows a solution to (3) other than the one dictated by fundamentals.[5] The existence of a wide variety of solutions for asset or goods prices satisfying a condition analogous to (3) has been emphasized by John Taylor (1977), Olivier Blanchard (1981), Robert Flood and Peter Garber (1980), Blanchard and Mark Watson (1982), and Behzad Diba and Herschel Grossman (1983). Solutions other than the "market fundamentals" solution have been variously called "rational bubble," "sunspot," and "bootstrap" equilibria.[6] I will use the name "bootstrap" for such solutions and reserve the term "bubble" for a related but distinct phenomenon. The possibility of bootstrap solutions in the foreign exchange markets was explicitly noted by Flood and Garber (1982) using a simple complete model incorporating a continuous time version of (3), and I give examples in Section III below.

Indirect tests for bootstrap equilibria have been of three types.[7]

(*i*) Variance bounds tests of Robert Shiller (1981) and Blanchard and Watson which, when applied to equities, find too much volatility in equity prices given the behavior of dividends. The principal problem with this type of test is that there are several alternative explanations, discussed by Stephen LeRoy (1984), which could explain such excess volatility without the necessity of assuming prices deviating from their fundamental values. However, Kenneth West (1985a) has recently argued that at least some of these problems can be overcome and has provided additional evidence of excess volatility.

(*ii*) Specification tests, for example, West (1985b), which compare two estimates of parameters required to calculate the "fundamentals" solution. Like (*i*), this test requires an assumed model for pricing the asset.

(*iii*) Runs tests, applied by Blanchard and Watson to gold prices. Here the tests failed to detect an excess of long runs which would have suggested bubbles. However, the tests may have had low power for reasons discussed by Blanchard and Watson,[8] and for other reasons mentioned below. It should also be noted that a tendency toward runs can result from certain distributions of fundamentals.

This paper proposes a new test for speculative bubbles, closer in spirit to the runs test, and applies it to the exchange rate between the U.K. pound sterling and the U.S. dollar. In order to clarify what null hypothesis is being tested, I provide a precise definition of the type of abnormality in excess returns which will be referred to as a

[5] Which variables constitute fundamentals depends upon the specification of the complete model.

[6] The definition of the fundamentals solution is discussed in Edwin Burmeister, Flood, and Garber (1983) and Bennett McCallum (1983). See also Section III below. The entire class of solutions to such models is given in C. Gourieroux, J. J. Laffont, and A. Monfort (1982). The disequilibrium stability of bootstrap equilibria has been considered in my paper (1985).

[7] Direct tests have been conducted by Flood and Garber (1980) and Diba and Grossman (1984), with negative results.

[8] Suppose there is a bubble caused by a bootstrap solution which has probability π of continuing. For π near one or one-half, the tendency toward long runs becomes negligible.

This content downloaded from 165.123.34.86 on Wed, 20 Jan 2016 22:42:21 UTC
All use subject to JSTOR Terms and Conditions

speculative bubble. For the purposes of this paper, I define a speculative bubble as a subperiod during which there is a nonzero *median* in the distribution of x_t, the excess return to holding foreign currency (if an allowance is made for risk premia, I alter the definition to require a nonzero median in the distribution of the risk-adjusted excess returns, $\tilde{x}_t$).

This definition has a number of advantages. First, there is some basis in common usage, since it implies a tendency towards an abnormal number of positive or negative excess returns during the subperiod. Second, if efficiency holds, then the existence of a speculative bubble must correspond to a skewed distribution of excess returns. The low probability of large excess returns of opposite sign to those that predominate can be interpreted as the possibility of a crash. Finally, this definition of speculative bubble is well suited to testing, since it is stated in terms of one measure of the central tendency of excess returns.

It must be emphasized at the outset, however, that detection of a speculative bubble so defined does not imply the detection of a bootstrap equilibrium or "rational bubble" in the sense of Flood and Garber (1980) or Blanchard and Watson. Other possible causes of a speculative bubble are a skewed distribution of fundamental innovations and nonrational solutions.[9]

In the next section I describe and carry out an appropriate nonparametric test which finds evidence of a speculative bubble in the sterling-dollar exchange rate over 1981–84. The following section shows that the evidence for a bubble remains strong when an allowance is made for risk premia. The remaining sections discuss at length the possible interpretations of the results.

I. A New Test for Bubbles

One of the major difficulties in testing for speculative bubbles is that there may be only one or two bubbles covering only part of the sample period. Furthermore, if there are two or more bubbles, they may be of opposite signs. We cannot, therefore, assume a uniform pattern over the whole sample period, and a powerful test must be designed to reflect this.

A. *Description of the Test*

Let x_t be the stochastic process of excess returns to holding foreign currency for $t = 1,\ldots,T$, and let m_t be the median of x_t in period t conditional on information through $t-1$. The null hypothesis H_0 is that $m_t = 0$ for $t = 1,\ldots,T$, and we desire a test designed to detect alternatives H_1 under which $m_t = m \neq 0$ for $t = T_1,\ldots,T_2$ where $1 \leq T_1 < T_2 \leq T$.

A suitable basis for a test of H_0 vs. H_1 is the sign test, a nonparametric test which looks at the difference between the number of positive and negative signs for x_t over the period.[10] This test would be directly applicable if the alternative were uniform, that is, $m_t = m \neq 0$ for $t = 1,\ldots,T$. The sign test has the advantages that it explicitly focuses on the conditional median and that no assumptions are required about the distribution of x_t other than continuity. In contrast, the usual t-test is clearly not an appropriate starting point, since it is based on the sample mean and since the validity of the test in small samples depends on assumptions of constant variance and normal kurtosis, whereas it is widely believed that changes in asset prices typically are heteroscedastic and leptokurtic.

The statistical problem is to adapt the sign test so that it is sensitive to a departure from 0 median over only part of the sample period. Although the sign test applied to the full-sample period might detect such a deviation, this test is likely to have very low power for such alternatives. On the other hand, the application of the sign test to a suspect subperiod, ignoring data from the rest of the full period, will be an invalid test as a result of data mining. For my particular applica-

[9] Those readers who object to this use of "speculative bubble" may substitute for it another phrase such as "empirical bubble" or "abnormality in excess returns," or choose their own terminology.

[10] A description of the sign test can be found in W. J. Dixon and A. M. Mood (1946).

This content downloaded from 165.123.34.86 on Wed, 20 Jan 2016 22:42:21 UTC
All use subject to JSTOR Terms and Conditions

tion, the question is whether the behavior of the exchange rate over 1981–84 was too extreme to have plausibly occurred by chance, under the null hypothesis, controlling for data mining from the post-1972 period of flexible exchange rates.

There is no simple procedure for obtaining a properly adjusted significance level for the apparent negative median of x_t over 1981–84. At the time of writing, the pound sterling has floated against the dollar for twelve years and eleven months,[11] giving a full sample period of 155 months.[12] There are thus just over three independent four-year subperiods and it is straightforward to adjust the significance level for choosing the "most significant" of several independent tests. In general, if a is the smallest significance level of n independent tests, then the true overall significance level α is given by $\alpha = 1-(1-a)^n$. But, of course, there are 108 four-year periods if we are permitted to start in any month. Indeed, there is no reason to limit ourselves to four-year periods, since this choice was dictated by the data, and thus we should allow for tests calculated from one-year, two-year, and so on up to twelve-year periods.

Since test statistics based on partially overlapping subperiods will be correlated, there is no simple way of adjusting the nominal significance level to obtain the true significance level. It can be shown that an upper bound for the overall significance level of n tests at nominal level a is na. However, in this case, where n is very large and where the tests under consideration are obviously highly correlated, the use of this bound would severely overinflate the true significance level.

The general procedure adopted in this paper is to directly estimate the overall significance level using a Monte Carlo study. This technique is particularly suited to the problem at hand since, in the case of the sign test, there are no unknown parameters on which the test statistic depends. In order to carry out the technique, however, I must first choose an appropriate test statistic, and again this is not obvious. For any particular subperiod, the absolute value of the deviation of the number of positive signs from half the number of observations in the subperiod is the appropriate statistic. But, given this statistic for numerous subperiods of varying lengths from the overall sample, how do we combine them to obtain an overall test statistic? This also is a question best answered by a Monte Carlo experiment which leads to the two-step procedure which I now describe.

In step 1, 10,000 random samples of 155 plusses and minuses were constructed (corresponding to the number of monthly data points over the flexible rate period), where each plus or minus was generated with probability of one-half.[13] For each sample of 155 observations, the statistics $z_1,\dots,z_{12}$ were calculated where, for $k=1,\dots,12$ years,

$$z_k = \text{Max } N_k,$$

where N_k is the absolute value of the number of plusses less the expected number of $6k$ and where the maximization is carried out over every k year subperiod in that sample of 155 months. Thus z_k is the most extreme deviation from the null hypothesis found in any k year subperiod. The restriction to whole year subperiods was made to reduce the computational burden, but the subperiods were allowed to begin in any month. This step of the Monte Carlo study

[11] The pound floated against the dollar in June 1972. By May 1973 the dollar was floating against most other major currencies. It should also be noted that, since 1947, exchange controls in varying degrees have been applied to U.K. residents until controls were finally abolished in October 1979. However, non-U.K. residents have been free to switch between sterling and dollars since 1958 (A. R. Prest and D. J. Coppock, 1984, pp. 133, 152, and 169). Hansen and Hodrick (1983) argue that until January 1976 there existed a transitional period in which agents may have anticipated a return to fixed exchange rates. Any such arguments, which reduce the effective period of freely floating exchange rates, strengthen the test results found in this paper.

[12] The choice of monthly data was motivated by the need to find matching data for the exchange rate, forward rate, interest rates, and commodity prices.

[13] Random numbers were computed using the NAG library pseudo-random number generator. This has a cycle length of 2^{57}. A total of 3.1 million random numbers were required for this study.

This content downloaded from 165.123.34.86 on Wed, 20 Jan 2016 22:42:21 UTC
All use subject to JSTOR Terms and Conditions

TABLE 1—CUMULATIVE DISTRIBUTION OF z_k

z_k	Subinterval Length: k Years 1	2	3	4	5	6	7	8	9	10	11	12
0	10000	10000	10000	10000	10000	10000	10000	10000	10000	10000	10000	10000
1	10000	10000	10000	10000	10000	10000	10000	10000	10000	10000	10000	10000
2	10000	10000	10000	10000	10000	10000	9999	10000	9998	9987	9936	9663
3	9978	9996	9987	9992	9984	9980	9964	9927	9842	9667	9333	8561
4	8348	9694	9770	9744	9735	9656	9585	9376	9152	8729	8179	7273
5	3033	7789	8682	8818	8854	8731	8539	8314	7941	7418	6838	6043
6	353	4460	6384	7050	7257	7270	7203	6975	6570	6134	5581	4913
7	–	1808	3883	4957	5409	5568	5616	5420	5170	4875	4466	3957
8	–	566	1950	3037	3620	3967	4110	4063	3985	3802	3538	3100
9	–	119	801	1685	2290	2662	2863	2947	2916	2863	2673	2348
10	–	13	278	807	1307	1657	1902	2046	2081	2073	1977	1756
11	–	2	97	363	687	1000	1210	1368	1450	1429	1386	1268
12	–	0	28	132	334	584	748	868	961	973	971	891
13	–	–	11	42	139	292	423	520	604	648	653	588
14	–	–	3	20	58	145	218	294	348	392	429	400
15	–	–	0	4	23	60	111	158	199	243	268	260
16	–	–	0	1	7	22	51	80	112	144	158	163
17	–	–	0	0	2	7	23	37	62	84	93	98
18	–	–	0	0	0	3	8	19	27	48	58	50
19	–	–	–	0	0	0	4	13	15	25	37	30
20	–	–	–	0	0	0	3	4	6	12	15	21
21	–	–	–	0	0	0	1	3	3	6	10	8
22	–	–	–	0	0	0	0	1	3	3	4	6
23	–	–	–	0	0	0	0	1	1	1	2	4
24	–	–	–	0	0	0	0	0	1	1	1	3
25	–	–	–	–	0	0	0	0	0	0	1	1
26	–	–	–	–	0	0	0	0	0	0	0	1
27	–	–	–	–	0	0	0	0	0	0	0	0

Note: For each subperiod of k years, the table provides $a_k(z_k)$, the number of times the value z_k was attained or exceeded in 10,000 simulations of 155 months each.

thus generated 10,000 values of z_k for each $k = 1,\dots,12$, and the cumulative distribution a_k, where for each k, $a_k(z_k)$ is the number of times z_k was obtained or exceeded, is provided in Table 1. Thus $a_k(z_k)$ represents the nominal significance level of test statistics based on subperiods of length k years.

Finally, for each sample, let

$$Y = \operatorname*{Min}_k a_k(z_k),$$

that is, where the minimization is carried out over the nominal significance levels of the z_k statistics for $k = 1,\dots,12$. The statistic Y is the natural overall test statistic to use to look for alternatives H_1. The statistic Y represents the most extreme value for the 576 subperiods considered, with low values of Y indicating greater deviations from the null hypothesis. To calculate the true significance level of a test based on Y we need the distribution of Y. This is found in step 2. Thus 10,000 new random samples of 155 plusses and minuses were generated. Using Table 1, for each random sample the value of Y was calculated. The cumulative distribution of Y was then calculated and this is given in Table 2.

To carry out this test using data on x_t, $t = 1,\dots,155$, one can calculate the z_k, $k = 1,\dots,12$, from the signs of x_t and compute Y using Table 1. Table 2 then provides an estimate of the true significance level for the observed Y. Alternatively, one calculates z_k for a suspect subperiod and uses the tables to find the corresponding Y and an upper bound on the true significance level. This test is so constructed that full protection is provided against data mining within the sample. Furthermore, there is no small sam-

This content downloaded from 165.123.34.86 on Wed, 20 Jan 2016 22:42:21 UTC
All use subject to JSTOR Terms and Conditions

TABLE 2—SIGNIFICANCE LEVELS FOR DISTRIBUTION OF Y

y	Significance Level	y	Significance Level
0	0.0002	111	0.0269
1	0.0005	112	0.0283
2	0.0009	119	0.0335
3	0.0011	132	0.0369
4	0.0015	139	0.0396
6	0.0022	144	0.0412
7	0.0027	145	0.0427
8	0.0029	158	0.0452
10	0.0030	163	0.0480
11	0.0037	199	0.0513
12	0.0041	218	0.0534
13	0.0057	243	0.0553
15	0.0061	260	0.0594
19	0.0063	268	0.0620
20	0.0067	278	0.0707
21	0.0073	292	0.0743
22	0.0078	294	0.0776
23	0.0090	334	0.0814
25	0.0094	348	0.0854
27	0.0099	353	0.1070
28	0.0108	363	0.1106
30	0.0110	392	0.1156
37	0.0119	400	0.1213
42	0.0135	423	0.1248
48	0.0143	429	0.1267
50	0.0149	520	0.1312
51	0.0159	566	0.1486
58	0.0182	584	0.1538
60	0.0187	588	0.1602
62	0.0195	604	0.1643
80	0.0205	648	0.1675
84	0.0213	653	0.1708
93	0.0224	687	0.1773
97	0.0247	748	0.1832
98	0.0259	801	0.2012

Note: Significance level is the estimated probability of obtaining a value of Y less than or equal to y.

ple bias since the test is based on a Monte Carlo study rather than asymptotic results.[14]

B. *Discussion of Risk Adjustment*

Before describing the results of the test, I return to the issue of risk. Although, as mentioned in the introduction, there is evidence of a nonzero risk premium, I initially give the results of applying the test just described to the unadjusted series, $x_{t+1} = s_{t+1} - f_t$, for the following reasons:

(i) There is by no means a unanimity of opinion on the existence of risk premia. For example, Jeffrey Frankel (1982) estimated a model using a portfolio balance approach and was unable to reject the hypothesis of risk neutrality.

(ii) Tests rejecting simple efficiency are usually based on asymptotic results. Although large data sets are in fact being used, these typically involve several correlated exchange rates and overlapping data subject to serial correlation, making the effective sample size much smaller. If the underlying distribution has high kurtosis and, as suggested in this paper, is nonsymmetric, one would expect the small sample bias to be exacerbated. Related points are discussed by W. S. Krasker (1980), Craig Hakkio (1983) and Robert Korajczyk (1985).[15]

(iii) The deviations between the forward and expected future spot rates, usually attributed to risk premia, are small compared to the difference between forward and actual future spot rates. For example, Table 1 of Lars Hansen and Robert Hodrick (1983) obtained an $R^2 = 0.095$ in the case of the sterling-dollar exchange rate for a regression of $x_{t+1} = s_{t+1} - f_t$ on eight regressors. Furthermore there appears to be general agreement that the unconditional expected value of x_t is 0.

(iv) There is no agreement on the appropriate empirical model of the risk premium. Thus any specific adjustment for risk will be subject to some criticism.

(v) There does not appear to be any persuasive reason, under the null hypothesis of no bubbles, for believing that the pound sterling was considered to be less risky on average than the dollar over the 1981–84

[14] This may be a significant advantage. The small sample bias of asymptotic approximations to some nonparametric statistics can be important, particularly in the tails, and the continuity correction does not necessarily improve the approximation (see E. L. Lehman, 1975, pp. 16–17).

[15] Another technical problem concerns the treatment of nonstationarity in some tests. Considering the related issue of risk premia in the term structure of interest rates, Gary Shea (1985) found that the simple expectations hypothesis could not be rejected when the model was estimated using fractional difference methods.

This content downloaded from 165.123.34.86 on Wed, 20 Jan 2016 22:42:21 UTC
All use subject to JSTOR Terms and Conditions

subperiod. This is the direction of correction required if risk is to explain the predominance of negative returns to holding sterling during this period.

These considerations suggest that it would be best to apply my test both to the unadjusted series for excess returns and to a risk-adjusted series. In the remainder of this section, I describe the results of the test for the unadjusted data.

C. *Test Results*

Table 3 presents point in time monthly data, 1981–84, for the U.S. dollar price of the U.K. pound sterling and for the realized value of x_t, the excess return to holding a one-month forward contract to buy sterling, expressed in percent per month. The data were collected from the *Financial Times* and refer to the last working day of the month. See the Data Appendix for further details. For comparison, the x_t series was also computed from (2) using one-month Eurocurrency rates. The results (not shown) exhibited a close correspondence and, in particular, the sign of x_t for every month was identical for the two methods of calculation.

During this 48-month period, the x_t series was positive in 9 months and negative in 39 months, providing a value of $z_4 = 15$. From Table 1, we see that this corresponds to a value for the test statistic of $Y = 4$. The significance level of this result, obtained from Table 2, is estimated to be 0.0015. Thus the null hypothesis of a zero median for x_t over the full period of floating can be rejected, against the alternative of a negative median over the 1981–84 period, at very low significance levels.

II. Risk Premia

Can the results of the previous section be explained by risk premia?[16] The modeling of risk premia in foreign exchange has been investigated, inter alia, by Frankel (1982), Hansen and Hodrick (1983), Hodrick and Sanjay Srivastava (1984), Korajczyk (1985), and Nelson Mark (1985). For the purposes of this paper, the most convenient method for correcting for risk is given by Korajczyk. All of the other methods require the estimation of parameters using specific models of risk.[17]

Two-country dynamic general equilibrium models with risk-averse agents of the type examined by Robert Lucas (1982) and others generate expressions specifying how bond risk premia are incorporated into real interest rates and can be used to obtain formulae for the risk premium in the foreign exchange market. Provided deviations from purchasing power parity satisfy the martingale property, it can be shown that the risk premium separating the forward price of foreign currency and the expected future spot price is equal to the difference between the expected real interest rates on default-free nominal bonds denominated in the two currencies.[18]

The simple efficiency condition (1) is thus replaced by the condition

$$E_t(s_{t+1} - f_t) = E_t r^*_{t+1} - E_t r_{t+1}, \tag{4}$$

where $r_{t+1} = i_{t,1} - \Delta p_{t+1}$, $r^*_{t+1} = i^*_{t,1} - \Delta p^*_{t+1}$, and p_t and p^*_t are logarithms of domestic and foreign price indices. Korajczyk found

[16] Another conceivable explanation of the results is transactions costs. Given the size of the exchange rate movements over this period, the existence of transactions costs does not appear to be a credible explanation. The average excess return to holding dollar rather than sterling interest-bearing accounts over 1981–84 was 1.595 percent per month, i.e., an annual rate of 19.14 percent. Since brokerage costs for currency conversion are quite small for large sums of money, and since they need only have been incurred twice, transactions costs would clearly have been quite minor compared to the gain of such a switch for, say, a year or more.

[17] Furthermore, Frankel's estimates were consistent with risk neutrality; two models considered by Hansen and Hodrick incorporated restrictions rejected by the data; the third model considered by Hansen and Hodrick was found by Hodrick and Srivastava to be rejected by the data; and Mark's estimates of the key risk-aversion parameter show enormous variation in magnitude depending on the detailed implementation.

[18] This result, which is derived in Korajczyk, was also stated in Hodrick and Srivastava, fn. 5. Further references are provided in both papers.

This content downloaded from 165.123.34.86 on Wed, 20 Jan 2016 22:42:21 UTC
All use subject to JSTOR Terms and Conditions

TABLE 3—EXCESS RETURNS TO HOLDING STERLING

Date	$Spot_t$	x_t	$\tilde{x}_t$
12/80	2.3910		
01/81	2.3670	−1.40537	−1.25349
02/81	2.2050	−7.38488	−7.24246
03/81	2.2445	1.51735	2.52236
04/81	2.1405	−4.95351	−2.57442
05/81	2.0700	−3.79193	−3.51709
06/81	1.9305	−7.48292	−7.26054
07/81	1.8415	−5.28804	−5.47266
08/81	1.8490	−0.11892	0.33561
09/81	1.8050	−2.77553	−2.85732
10/81	1.8600	2.93514	3.63439
11/81	1.9550	5.06205	5.72062
12/81	1.9100	−2.00593	−1.96500
01/82	1.8810	−1.37277	−1.33888
02/82	1.8215	−3.22496	−3.53168
03/82	1.7820	−2.26375	−1.24831
04/82	1.7940	0.47493	2.20369
05/82	1.7905	−0.35124	−0.49467
06/82	1.7435	−2.77167	−3.63136
07/82	1.7380	−0.54225	−0.86245
08/82	1.7170	−1.26167	−1.40198
09/82	1.6945	−1.31618	−1.55394
10/82	1.6770	−1.08237	−0.83443
11/82	1.6235	−3.24520	−2.60923
12/82	1.6175	−0.29940	−0.16254
01/83	1.5200	−6.12127	−6.35737
02/83	1.5150	−0.16488	0.00571
03/83	1.4835	−1.85330	−1.96748
04/83	1.5605	5.14451	5.81921
05/83	1.6045	2.90563	2.66080
06/83	1.5340	−4.40917	−4.58661
07/83	1.5210	−0.85433	−0.73273
08/83	1.4940	−1.83383	−1.70438
09/83	1.4970	0.17049	0.14369
10/83	1.4955	−0.09023	−0.02817
11/83	1.4630	−2.23391	−2.02398
12/83	1.4515	−0.86091	−0.67521
01/84	1.4015	−3.56399	−4.12883
02/84	1.4905	6.11761	6.09640
03/84	1.4425	−3.34381	−3.19791
04/84	1.3985	−3.24669	−2.26886
05/84	1.3855	−1.13037	−0.87848
06/84	1.3565	−2.30640	−2.22165
07/84	1.3075	−3.93311	−4.12599
08/84	1.3085	0.12619	0.58454
09/84	1.2350	−5.86507	−6.09918
10/84	1.2210	−1.19270	−0.79855
11/84	1.1980	−1.82793	−1.59381
12/84	1.1590	−3.24697	−3.46695

Notes: Date is last working day of the month; $Spot_t$ is U.S. dollar price of the U.K. pound sterling; x_t is excess return to holding sterling calculated according to (1), expressed as percent per month; $\tilde{x}_t$ is risk-adjusted excess return to holding sterling calculated according to (5), expressed as percent per month.

that deviations from (1) were consistent empirically with the reformulation (4).

To conduct my test for speculative bubbles, allowing for risk premia, I should subtract expected real interest rate differentials from the series for x_t. Unfortunately, expected real interest rates are not observable. While they could be estimated by a time-series model, a simpler procedure is to note that (4) implies

$$(5) \qquad E_t \tilde{x}_{t+1} = 0,$$

where $\tilde{x}_{t+1} = s_{t+1} - f_t + r_{t+1} - r^*_{t+1}$,

and that it is straightforward to apply the test of the preceding section to the risk-adjusted excess returns $\tilde{x}_{t+1}$, calculated using *ex post* real interest rates. The cost of conducting the test using *ex post* differentials in real interest rates rather than *ex ante* differentials is a possible reduction in the power of the test since $\tilde{x}_{t+1}$ includes noise from unforecastable changes in goods prices.[19] However, in view of the strong results of the preceding section, this potential cost would appear to be acceptable.

The final column of Table 3 gives the monthly series for $\tilde{x}_t$ over 1981–84, calculated according to (5) using one-month dollar and sterling Eurocurrency rates for $i_{t,1}$ and $i^*_{t,1}$ and one-month changes in the U.S. Consumer Price Index and the U.K. Retail Price Index for p_t and p^*_t.[20] The risk adjustment changes the sign of excess returns from negative to positive in 2 of the 48 months, so that $z_4 = 13$ for the 1981–84 subperiod.[21] According to Table 1, this corresponds to $Y = 42$. The overall significance level, ob-

[19]Nominal interest rates for the coming period are included in the information set.

[20]As noted by Korajczyk, the price indices provide a better measure of goods prices midmonth, so that there is some mismatch of timing in the calculation of $\tilde{x}_t$.

[21]Again, there are two ways to calculate $\tilde{x}_t$. Using covered interest arbitrage, it follows that $\tilde{x}_{t+1} = s_{t+1} - s_t + \Delta p^*_{t+1} - \Delta p_{t+1}$, so that $\tilde{x}_t$ is simply the change in the real exchange rate. For comparison, $\tilde{x}_t$ was recalculated using this formula and found to match closely the series given in Table 3. For every month the sign of $\tilde{x}_t$ was the same for both methods of calculation.

This content downloaded from 165.123.34.86 on Wed, 20 Jan 2016 22:42:21 UTC
All use subject to JSTOR Terms and Conditions

tained from Table 2, is 0.0135. Thus, although the risk adjustment does lead to a higher significance level than when the test is applied to the unadjusted series, there is still very considerable evidence against the null hypothesis of a zero median for risk-adjusted excess returns and in favor of a negative median over the 1981–84 period.

III. Possible Interpretations

The results of the previous two sections constitute strong evidence, according to my definition, of an empirical speculative bubble in the dollar-sterling exchange rate over 1981–84. Apart from type I error, there are three possible explanations for the finding: (*i*) bootstrap equilibria (rational bubbles); (*ii*) nonsymmetric innovations in fundamentals; and (*iii*) nonrational or disequilibrium expectations.

These points can be made clearly by examining a simple exchange rate model, adapted from a continuous time model of Flood and Garber (1982), and used here for expositional purposes only. The economy is described by the following three equations:

$$m_t - p_t = a + by_t - ci_{t,1}, \tag{6}$$

$$i_{t,1} = i^*_{t,1} + \hat{s}_{t+1} - s_t, \tag{7}$$

$$p_t = p^*_t + s_t, \tag{8}$$

where m is the logarithm of the domestic money supply, y is the logarithm of domestic output, and other variables have been previously defined. Equation (6) is the money market equilibrium equation, and b and c are positive parameters. (7) is the arbitrage equation, and $\hat{s}_{t+1}$ denotes the expected value for the spot exchange rate in period $t+1$ held by agents in period t. Under rational expectations $\hat{s}_{t+1} = E_t s_{t+1}$. For simplicity, I assume zero transactions costs and risk neutrality. (8) states that purchasing power parity holds in every period.

Solving (6)–(8) for s_t, we obtain

$$s_t = \theta \hat{s}_{t+1} + v_t, \tag{9}$$

where $0 < \theta = c(1+c)^{-1} < 1$ and $v_t = (1+c)^{-1}(m_t - p^*_t - a - by_t + ci^*_{t,1})$. Under rational expectations, the solution which I have referred to as the "market fundamentals" solution is

$$\bar{s}_t = \sum_{i=0}^{\infty} \theta^i E_t v_{t+i}, \tag{10}$$

providing this sum converges. Assume that v_t is exogenous and has a (stationary or nonstationary) ARMA representation in a symmetric white noise innovation ε_t, that is, $F(L)v_t = G(L)\varepsilon_t$ where L is the lag operator and where ε_t is a white noise disturbance with median 0. The innovation in $\bar{s}_t$ is given by

$$\bar{s}_{t+1} - E_t\bar{s}_{t+1} = \theta^{-1} \sum_{i=1}^{\infty} \theta^i (E_{t+1}v_{t+i} - E_t v_{t+i}).$$

Since the revision in the expectation of v_{t+i} must simply be proportional to ε_{t+1}, that is, $E_{t+1}v_{t+i} - E_t v_{t+i} = \alpha_i \varepsilon_{t+1}$, we have

$$\bar{s}_{t+1} - E_t\bar{s}_{t+1} = \left[\theta^{-1} \sum_{i=1}^{\infty} \theta^i \alpha_i\right] \varepsilon_{t+1}.$$

It follows that $\bar{s}_{t+1} - E_t\bar{s}_{t+1}$ has median 0 and that

$$\Pr[x_{t+1} \geq 0] = \Pr[s_{t+1} - E_t s_{t+1} \geq 0] = 1/2,$$

where Pr stands for the probability conditional on information available at time t. Hence x_{t+1} has zero conditional median providing expectations are rational, the exchange rate is determined by the fundamentals solution, and innovations in the fundamentals are symmetric.

A. *Bootstrap Equilibria*

Consider, now, solutions to (9) other than (10). Blanchard and Watson point out that such solutions, which they term "rational bubbles," may result in a tendency towards runs (a nonzero median) though this is not necessarily the case. The rational expectations solutions to (10) can always be written

This content downloaded from 165.123.34.86 on Wed, 20 Jan 2016 22:42:21 UTC
All use subject to JSTOR Terms and Conditions

in the form $s_t = \bar{s}_t + c_t$ where c_t satisfies $c_t = \theta E_t c_{t+1}$. An asymmetry in the innovation of the bootstrap or bubble term, c_t, can lead to an asymmetry in the innovation of s_t. As an example they discuss the solution given by

$$c_t = \begin{cases} (\pi\theta)^{-1} c_{t-1} + \mu_t & \text{with prob. } \pi \\ \mu_t & \text{with prob. } 1-\pi \end{cases}$$

where μ_t satisfies $E_{t-1}\mu_t = 0$. It follows from their analysis that if μ_t is symmetric, and if the innovations in market fundamentals are symmetric, then the conditional median of x_t will be nonzero for $1/2 < \pi < 1$. Essentially the small probability of a large crash is balancing the likelihood of a continuation of the bubble. It should be noted that μ_t may either be a function of the innovation in the fundamentals, or be a variable which is wholly independent of the fundamentals (in which case the solution is often called a "sunspot" equilibrium).

B. *Asymmetric Fundamentals*

A second possibility is that the fundamentals themselves may have nonsymmetric innovations. Suppose that $v_t = v_t^* + \varepsilon_t$, where ε_t is a symmetric white noise disturbance, and that $\Delta v_t^* = g\omega_t$, where v_0^* is given, where $\omega_0 = 1$ and where

$$\omega_t = \begin{cases} \omega_{t-1} & \text{with prob. } \pi \\ 0 & \text{with prob. } 1-\pi. \end{cases}$$

The $0-1$ variable ω_t may be thought of as modeling a permanent regime change. Since $E_t v_{t+i+1}^* = E_t v_{t+i}^* + gE_t\omega_{t+i+1}$, we have $E_t v_{t+i}^* = v_t^* + g\pi(1-\pi)^{-1}(1-\pi^i)$, from which it follows that the fundamentals solution is given by

$$\bar{s}_t = \begin{cases} (1-\theta)^{-1} v_t^* + g\pi\theta(1-\theta)^{-1} \\ \quad \times (1-\theta\pi)^{-1} + \varepsilon_t & \text{if } \omega_t = 1, \\ (1-\theta)^{-1} v_t^* + \varepsilon_t & \text{if } \omega_t = 0. \end{cases}$$

Once the permanent regime change to $\omega = 0$ has occurred, the distribution of $\bar{s}_t$ is symmetric. However, it is straightforward to show that $\Pr[\bar{s}_{t+1} - E_t\bar{s}_{t+1} \geq 0] > 1/2$ if $1/2 < \pi < 1$ as long as $\omega_t = 1$. Thus if agents believe, for example, that there is a small probability of a regime change in any given period, such as a permanent shift in monetary policy, but a high probability that such a change will eventually occur, then this will induce a highly nonsymmetric conditional distribution to $\bar{s}_t$ and hence a nonzero conditional median to x_t. In this case it would be natural to say that there is a bubble in the fundamentals.

C. *Nonrational Expectations*

The final possibility is that expectations are not rational in the strict sense of being conditionally unbiased.[22] The specific possibilities run the range from transitional deviations as agents learn about the model as in Taylor (1975) to the more dramatic departures of the sort described in Charles Kindleberger (1978).

Suppose that the fundamentals follow a random walk with drift, that is, $\Delta v_t = d + \varepsilon_t$, so that

$$\bar{s}_t = (1-\theta)^{-2}\theta d + (1-\theta)^{-1} v_t.$$

Suppose also that at time $t = 0$, the drift d, which is not directly observable, falls to d'. If expectations initially remain at $\hat{s}_{t+1} = s_t + (1-\theta)^{-1}d$ then $s_t = (1-\theta)^{-2}\theta d + (1-\theta)^{-1}v_t$ where $\Delta v_t = d' + \varepsilon_t$ and it follows[23] that $\Pr[x_{t+1} \geq 0] < 1/2$. If forecasts are revised on the basis of recent data as their bias becomes recognized, then expectations may eventually converge to the new rational forecast function. However, until convergence has been achieved, we can expect x_t to have a negative conditional median and mean.

The existence of such nonrational expectations does not require imperfect knowledge

[22] Allowing for risk, I would restate this last possibility as $E_t\tilde{x}_{t+1} \neq 0$. The views of John Bilson (1981) seem to belong to this category. Bilson has argued that the foreign exchange markets are inefficient in the sense that trading strategies can be developed which are too profitable to be explained by compensation for risk.

[23] I assume that $f_t = \hat{s}_{t+1}$ so that $x_{t+1} = s_{t+1} - \hat{s}_{t+1}$.

This content downloaded from 165.123.34.86 on Wed, 20 Jan 2016 22:42:21 UTC
All use subject to JSTOR Terms and Conditions

of the true model. Suppose that, at $t=0$, d is reduced to $d=0$ as a result of a government policy which is announced, fully credible, and actually adhered to. But, suppose that individually rational agents are not confident that other agents will immediately and fully adjust their expectations, either because they do not believe that other agents are rational (or that other agents believe that *they* are rational), or because they are not confident that other agents find the announced policy credible. Then again expectations will be biased, and, although they may well converge to equilibrium, it is likely that the mean and median of x_t will be negative during the transition. Arguments of this sort are extensively discussed in the volume edited by Roman Frydman and Edmund Phelps (1983).

As a final example, suppose that v_t is symmetric white noise and that agents believe the spot rate is following the path

$$s_t = \begin{cases} \beta + \delta t + v_t & \text{for } t = 0,\ldots,T-1 \\ v_t & \text{for } t \geq T. \end{cases}$$

With $\hat{s}_{t+1}$ computed on this assumption, the actual path followed by s_t will be

$$s_t = \begin{cases} \theta(\beta+\delta)+\theta\delta t + v_t & \text{for } t = 0,\ldots,T-2 \\ v_t & \text{for } t \geq T-1. \end{cases}$$

It is easy to check that the conditional median and mean of x_t will generally deviate from zero for $t=0,\ldots,T-1$. However, for θ near 1, δ small and appropriate choice of β, the expectations of agents will not be badly erroneous except that the price will collapse one period earlier than expected. This example suggests that there may be a wide class of "near rational" solutions.[24]

D. *Discussion*

Is it possible to distinguish between the three types of bubbles just described? It would appear to be very difficult. One might hope to detect nonrational bubbles by tests based on the mean of x_t. However, the detection of a nonzero mean will be difficult for two reasons. First, if the deviation from rationality occurs over only part of the sample period, then techniques analogous to those of this paper must be developed for the sample mean. Second, most tests based on the mean assume a symmetric distribution, yet rationality is consistent with skew distributions. Furthermore, it is clear that a highly skew distribution of x_t, whether the result of a bootstrap solution or of a bubble in the fundamentals, can statistically mimic irrationality. Since an apparent bias over a limited period of time can always be explained by the possibility of a rare large event of opposite sign, it may be impossible to develop a statistically powerful test to detect transitory irrationality.[25] Indeed, for conjectured rare or unique events, it may be impossible to define the objective probability distribution, making it difficult to distinguish conceptually between rational and irrational expectations.

Distinguishing between the types of rational bubbles also presents considerable difficulties. In a somewhat different context, it has been argued by James Hamilton and Charles Whiteman (1985) that apparent evidence of what I have referred to as a bootstrap equilibrium can be reinterpreted as having arisen from a suitable specification of fundamentals which are unobserved by the econometrician. The above examples illustrate this point for the case at hand. The negative median for x_t over 1981–84 could equally well be explained by a bootstrap bubble in which the run of negative returns on holding sterling was compensated by the possibility of a collapse of the dollar or, say, by the discounting of a strong U.K. output boom, expected to eventually occur but not fully realized during the period. Distinguish-

[24]An interesting question is whether there exist near rational solutions in the more precise sense of George Akerlof and Janet Yellen (1985).

[25]There appears to be little point examining the x_t data for skewness over 1981–84. A finding of skewness would still be compatible with irrationality. An apparently symmetrical empirical distribution would still be compatible with theoretical skewness if the latter is due to rare possibilities not actually observed.

This content downloaded from 165.123.34.86 on Wed, 20 Jan 2016 22:42:21 UTC
All use subject to JSTOR Terms and Conditions

ing empirically could be particularly difficult since a bootstrap equilibrium can be a function solely of fundamentals.

Nevertheless, imposing rationality is a restriction and assuming a fundamentals solution is an additional restriction. What is clear from the above discussion is that distinguishing between the competing explanations for a bubble given above can only be attempted within the context of a complete model of the exchange rate.[26]

IV. Conclusions

This paper has developed a test for a type of abnormality in excess returns which we have taken to define a bubble. In particular, I have defined a speculative bubble as occurring during any subperiod in which the (risk-adjusted) excess return on holding an asset has a nonzero median. This definition, while not conventional, is a natural use of the term, since it implies a tendency toward an abnormal number of positive or negative excess returns and has the advantage of being directly testable from data on excess returns. Indeed, with speculative bubbles so defined in terms of this measure of the central tendency of excess returns, it is possible to test for bubbles using nonparametric methods and without assuming any particular model of the exchange rate. The principal statistical complication, treated in this paper, is the necessity to develop a procedure for detecting a nonzero median over only part of the whole sample period. Equivalently, we may think of this as a procedure which controls for data mining.

When applied to the sterling-dollar exchange rate this test detects the presence of a negative median excess return on holding sterling rather than dollar assets over the 1981–84 period. After carefully controlling for possible data mining from the full period of flexible rates, the significance level of the test is 0.0015 when using unadjusted excess returns and 0.0135 when an allowance is made for possible risk premia. The finding of a bubble during 1981–84 can be explained by either nonrational expectations, bootstrap equilibria (i.e., rational bubbles) or nonsymmetric fundamentals. Distinguishing between these alternatives would be difficult and will require a complete model of the exchange rate (which will then become part of the maintained hypothesis). An intriguing open question is whether some variable can be found, whether interpreted as a fundamental or as a sunspot variable, which can explain the bubble found for the sterling-dollar rate during this period.

Many other questions remain to be investigated. The sterling-dollar rate was chosen for analysis because of the considerable discussion among policymakers in each country, at various times, suggesting that they believed the exchange rate between the currencies was out of line with its equilibrium value. Will empirical speculative bubbles be detected by these methods between other currencies and in other assets?[27] It would be desirable to extend this type of test to a multivariate context so that the exchange rate system could be analyzed as a unit. More powerful versions of the test could be devised which make use of the magnitudes of the changes or consider alternatives of multiple bubbles. Such extensions are nontrivial and must await further research.

DATA APPENDIX

Except as noted below, the spot U.S. dollar price of the U.K. pound sterling, its

[26]After completing this paper, I became aware of Richard Meese (forthcoming) and Frankel and Kenneth Froot (1985). Meese considers various evidence for rational bubbles (i.e., bootstrap equilibria) in foreign exchange markets, focusing on a specification test which is conditioned on a particular model of the exchange rate. Frankel and Froot construct an account of the 1981–84 period, partly in terms of irrational speculation, which provides one possible explanation of the empirical findings of this paper.

[27]It should be noted that if bubbles are of short duration, they will be difficult to detect. For example, if bubbles last 12 months, and only appear once every 13 years, then they will be virtually undetectable using monthly data. From Tables 1 and 2, it can be seen that the significance level of a run of twelve negatives is only 0.107. The data spoke clearly in this case because the bubble lasted 4 years.

This content downloaded from 165.123.34.86 on Wed, 20 Jan 2016 22:42:21 UTC
All use subject to JSTOR Terms and Conditions

one-month forward premium, and the sterling and U.S. dollar one-month Eurocurrency interest rates were taken from issues of the *Financial Times* of London and refer to the last working day of the month. Exchange rate data were taken from the table labelled "Pound Spot and Forward" and refer to the market-closing price. Interest rate data were taken from the "Euro-Currency Interest Rates" table and refer to market-closing rates. In each case the figures used were the averages of the ranges given. The *Financial Times* was not published from June 1 to August 8, 1983, due to a printing dispute. For May 31, June 30, and July 29, 1983, equivalent figures were taken from the London *Times*. The U.S. Consumer Price Index for all urban consumers was taken from the April 1985 issue of *Business Conditions Digest*, Table 320, and the U.K. Index of Retail Prices for All Items was taken from Central Statistical Office publication *Economic Trends*, issues for May 1982 (Table 18.8), May 1983 (Table 18.8), and May 1985 (Table 18.1).

REFERENCES

Akerlof, George A. and Yellen, Janet L., "Can Small Deviations from Rationality Make Significant Differences to Economic Equilibria?," *American Economic Review*, September 1985, *75*, 708–20.

Baillie, Richard T., Lippens, Robert E. and McMahon, Patrick, "Testing Rational Expectations and Efficiency in the Foreign Exchange Market," *Econometrica*, May 1983, *51*, 553–63.

Bilson, John F. O., "The 'Speculative Efficiency' Hypothesis," *Journal of Business*, July 1981, *54*, 435–52.

Blanchard, Olivier J., "Speculative Bubbles, Crashes and Rational Expectations," *Economic Letters*, 1981, *4*, 387–89.

______ and Watson, Mark, "Bubbles, Rational Expectations and Financial Markets," in P. Wachtel, ed., *Crises in the Economic and Financial Structure*, Lexington: Lexington Books, 1982.

Burmeister, Edwin, Flood, Robert P. and Garber, Peter M., "On the Equivalence of Solutions in Rational Expectations Models," *Journal of Economic Dynamics and Control*, May 1983, *5*, 311–21.

Diba, Behzad and Grossman, Herschel I., "Bubbles, Rational Expectations and Financial Markets," NBER Working Paper No. 1059, 1983.

______ and ______, "Rational Bubbles in the Price of Gold," NBER Working Paper No. 1300, 1984.

Dixon, W. J. and Mood, A. M., "The Statistical Sign Test," *Journal of the American Statistical Association*, 1946, *41*, 557–66.

Dornbusch, Rudiger, "Expectations and Exchange Rate Dynamics," *Journal of Political Economy*, December 1976, *84*, 1161–76.

Edwards, Sebastian, "Floating Exchange Rates, Expectations and New Information," *Journal of Monetary Economics*, May 1983, *11*, 321–36.

Evans, George W., "Expectational Stability and the Multiple Equilibria Problem in Linear Rational Expectations Models," *Quarterly Journal of Economics*, November 1985, *100*, 1217–33.

Flood, Robert P. and Garber, Peter M., "Market Fundamental versus Price Level Bubbles: The First Tests," *Journal of Political Economy*, August 1980, *88*, 745–70.

______ and ______, "Bubbles, Runs, and Gold Monetization," in P. Wachtel, ed., *Crises in the Economic and Financial Structure*, Lexington: Lexington Books, 1982.

Frankel, Jeffrey A., "In Search of the Exchange Risk Premium: A Six-Currency Test Assuming Mean Variance Optimization," *Journal of International Money and Finance*, December 1982, *1*, 255–74.

______ and Froot, Kenneth, "The Dollar as an Irrational Speculative Bubble: A Tale of Fundamentalists and Chartists," mimeo., University of California-Berkeley, December 1985.

Frenkel, Jacob A., "The Forward Exchange Rate, Expectations, and the Demand for Money: The German Hyperinflation," *American Economic Review*, September 1977, *67*, 653–70.

______, "Flexible Exchange Rates, Prices, and the Role of 'News': Lessons from the 1970s," *Journal of Political Economy*, August 1981, *89*, 655–705.

This content downloaded from 165.123.34.86 on Wed, 20 Jan 2016 22:42:21 UTC
All use subject to JSTOR Terms and Conditions

______, *Exchange Rates and International Macroeconomics*, NBER, Chicago: University of Chicago Press, 1983.

Frydman, Roman and Phelps, Edmund S., *Individual Forecasting and Aggregate Outcomes: 'Rational Expectations' Examined*, New York: Cambridge University Press, 1983.

Geweke, John and Feige, Edward, "Some Joint Tests of the Efficiency of Markets for Forward Foreign Exchange," *Review of Economics and Statistics*, August 1979, *61*, 334–41.

Gourieroux, C., Laffont, J. J. and Monfort, Alain, "Rational Expectations in Linear Models: Analysis of Solutions," *Econometrica*, March 1982, *50*, 409–25.

Hakkio, Craig S., "Expectations and the Forward Exchange Rate," *International Economic Review*, October 1981, *22*, 663–78.

______, "Comment" on L. P. Hansen, and R. J. Hodrick, in J. A. Frenkel, ed., *Exchange Rates and International Macroeconomics*, Chicago: University of Chicago Press, 1983.

Hamilton, James D. and Whiteman, Charles H., "The Observable Implications of Self-fulfilling Expectations," *Journal of Monetary Economics*, November 1985, *16*, 353–74.

Hansen, Lars Peter and Hodrick, Robert J., "Forward Exchange Rates as Optimal Predictors of Future Spot Rates: An Econometric Analysis," *Journal of Political Economy*, October 1980, *88*, 829–53.

______ **and** ______, "Risk Averse Speculation in the Forward Foreign Exchange Market: An Economic Analysis of Linear Models," in J. A. Frenkel, ed., *Exchange Rates and International Macroeconomics*, Chicago: University of Chicago Press, 1983.

Hodrick, Robert J. and Srivastava, Sanjay, "An Investigation of Risk and Return in Forward Foreign Exchange," *Journal of International Money and Finance*, April 1984, *3*, 5–29.

Kindleberger, Charles, P., *Manias, Panics and Crashes: A History of Financial Crises*, New York: Basic Books, 1978.

Korajczyk, Robert A., "The Pricing of Forward Contracts for Foreign Exchange," *Journal of Political Economy*, April 1985, *93*, 346–68.

Krasker, W. S., "The 'Peso Problem' in Testing the Efficiency of Forward Exchange Markets," *Journal of Monetary Economics*, April 1980, *6*, 269–76.

Lehman, E. L., *Nonparametrics*, San Francisco: Holden-Day, 1975.

LeRoy, Stephen F., "Efficiency and the Variability of Asset Prices, *American Economic Review Proceedings*, May 1984, *74*, 183–87.

Lucas, Robert E., Jr., "Interest Rates and Currency Prices in a Two-Country World," *Journal of Monetary Economics*, November 1982, *10*, 335–60.

McCallum, Bennett T., "On Non-uniqueness in Rational Expectations Models: An Attempt at Perspective," *Journal of Monetary Economics*, March 1983, *11*, 139–68.

McKinnon, Ronald, I., *Money in International Exchange: The Convertible Currency System*, New York: Oxford University Press, 1979.

______, *An International Standard for Monetary Stabilization*, Washington: Institute for International Economics, 1984.

Mark, Nelson C., "On Time Varying Risk Premia in the Foreign Exchange Market; An Economic Analysis," *Journal of Monetary Economics*, July 1985, *16*, 3–18.

Meese, Richard A., "Testing for Bubbles in Exchange Markets: A Case of Sparkling Rates?," *Journal of Political Economy*, forthcoming.

Prest, A. R. And Coppock, D. J., *The UK Economy*, 10th ed., London: Weidenfeld and Nicolson, 1984.

Shea, Gary S., "Long Memory Models of Interest Rates, the Term Structure, and Variance Bounds Tests," International Finance Discussion Paper No. 258, Federal Reserve Board, July 1985.

Shiller, Robert J., "Do Stock Prices Move Too Much to be Justified by Subsequent Changes in Dividends?," *American Economic Review*, June 1981, *71*, 421–36.

Taylor, John B., "Monetary Policy during a Transition to Rational Expectations," *Journal of Political Economy*, October 1975, *83*, 1009–22.

______, "Conditions for Unique Solutions in Stochastic Macroeconomic Models with

This content downloaded from 165.123.34.86 on Wed, 20 Jan 2016 22:42:21 UTC
All use subject to JSTOR Terms and Conditions

Rational Expectations," *Econometrica*, September 1977, *45*, 1377–85.

West, Kenneth D., (1985a) "Dividend Innovations and Stock Price Volatility," mimeo., Princeton University, 1985.

______, (1985b) "A Specification Test for Speculative Bubbles," mimeo., Princeton University, 1985.

United Kingdom, Central Statistical Office, *Economic Trends*, London, England.

U.S. Bureau of Economic Analysis, *Business Conditions Digest*, Washington, D.C.

This content downloaded from 165.123.34.86 on Wed, 20 Jan 2016 22:42:21 UTC
All use subject to JSTOR Terms and Conditions