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LOOKING FOR RATIONAL EXCHANGE RATE BUBBLES

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Abstract

In this paper simple stationarity and cointegration tests are used to check the empirical relevance of exchange rate bubbles for a class of models that either assume purchasing power parity (PPP) or arrive at a PPP-type relationship. While the possibility of bubbles in the dollar/deutschemark and the dollar/pound exchange rates over the post-1973 free floating period cannot be excluded, the presence of such indeterminacies is not substantiated. Useful extensions of the tests for future research are also suggested.

1. Introduction

The failure of structural asset market models[1] to explain the large variability of exchange rates over the recent free floating period has often been attributed to their abstraction from risk considerations and to expected violent changes in policy which are not observed in the sample (the "peso problem"). The appealing feature of such explanations is that they are compatible with market efficiency. However, the available econometric evidence casts doubts on the ability of a time-varying risk premium to explain the actual exchange rate behavior [see Hansen and Hodrick (1983), Levich (1985), Mark (1985), Cumby (1988)]. Similarly, "peso problem" effects do not appear to be an empirically relevant explanation for a seemingly inefficient market [see Krugman (1989), Engel and Hamilton (1990)].

The ambiguity surrounding the econometric evidence on risk premia and "peso problem" effects has given rise to an approach that relies on the contribution of self-fulfilling prophecies in explaining the observed behavior of exchange

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rates. Self-fulfilling prophecies, or dynamic indeterminacies, occur whenever market participants believe that extraneous events unrelated to market fundamentals drive price changes. Such events, which may contribute to explosive asset price behavior, are termed 'sunspots' or 'bubbles' in the relevant literature.

The basis of the bubble approach to explaining the volatility of exchange rates has been provided by the asset market theory of exchange rate determination. In particular, the dependence of the current spot exchange rate on the expected future spot exchange rate allows for solutions for the current spot rate which may not reflect only market funfamentals. More importantly, the presence of bubbles in asset markets is fully compatible with the assumption of rational agents who make efficient use of all available information. As Dornbusch (1982) and Hardouvelis (1988) point out, in the case of bubbles market participants are aware of the deviation of the current price from its fundamental value. Nevertheless, the bubble may be sustained as long as the bubble premium (i.e. the return in excess of the risk-free return and the risk premium) is high enough to compensate for losses resulting from a bubble crash. In addition, Diba and Grossman (1984) note that if everyone in the market uses the bubble component to form price forecacts, then it would be irrational on the part of an individual not to do so. Thus, cumulative deviations of exchange rates from their fundamental values need not be associated with irrational behavior on the part of market participants, nor do they imply market inefficiency[2].

The ability of the bubble approach to provide a theoretical rationale for events like the stock market crashes of 1929 and 1987 and the sudden depreciation of the U.S. dollar in 1985 has popularized the theoretical and empirical work on the validity of bubble equilibria[3]. However, econometric tests of the no bubbles hypothesis for various markets have produced mixed results. Specifically, the no exchange rate bubbles hypothesis is rejected in studies by Meese (1986), Evans (1986), Woo (1987), and Krugman (1987) but these findings have been challenged by the results of West (1987) and Kearney and MacDonald (1990)[4].

Bubble tests are subject to a number of problems, the foremost of them being their inability to provide conclusive evidence that bubbles exist. Indeed, bubbles are model-specific and therefore the no bubbles hypothesis is tested jointly with the hypothesis that the underlying model that generates the fundamental values of an asset price is correctly specified. Furthermore, bubble paths may be easily confused with asset price paths generated by "peso problem" effects due to anticipated fundamental innovations which are not accounted for by an econometrician. Thus, while bubble tests can provide conclusive evidence against the presence of bubbles, a rejection of the no bubbles hypothesis does not necessarily imply that a buble path exists. For this reason, Flood and Hodrick (1990) contend that "no econometric test has yet demonstrated that bubbles are present in the data" and that "bubble tests are interesting specification tests" of the underlying equilibrium model. Nevertheless, the use of bubble tests as specifications tests of market fundamentals is legitimate only when we are certain that bubbles, "peso problem" effects, or regime changes are not relevant[5].

This paper applies time series methods to test the existence of explosive rational bubbles in the deutschemark/dollar and the dollar/pound exchange rates over the free floating period January 1974 - December 1987. The approach taken here is based on an application of a unit root test and a cointegration test in studying the foreign exchange market. In section 2 we analyze the implications of divergent exchange rate bubbles for the properties of the underlying time series. The empirical methodology is discussed in section 3 and our data and results are presented in section 4. Final remarks and suggestions for future research are given in section 5.

2. Explosive Exchange Rate Bubbles: Theoretical Issues

Linear asset market models of exchange rate determination assume the validity of either a short-run or a long-run Purchasing Power Parity (PPP) relationship. For example, the simple monetary model assumes short-run PPP [see Frenkel and Mussa (1985)] while Dornbusch's (1976) overshooting model is based on the assumption that only long-run PPP is maintained. Furthermore, exchange rate models which are based on extensions of the Capital Asset Pricing Model (CAMP) [see Lucas (1978, 1982)] arrive at a PPP-type relationship of the form:

$$e_t = \frac{P_t U_c^* (c_t, c_t^*)}{P_t^* U_c (c_t, c_t^*)} , \qquad (1)$$

where e_t is the nominal exchange rate (units of domestic currency per unit of foreign currency), P_t is the domestic price level, P_t^* is the foreign price level, $U(c_t, c_t^*)$ denotes utility derived from consumption of the domestic (c_t) and the foreign (c_t^*) goods, U_c and U_c^* are partial derivatives, and the subscripts denote the time-period.

Under the assumption of finite endowments, equation (1) implies that explosive exchange rate paths should be associated with either a hyperinflation in the home-country or a hyperdeflation in the foreign country or both. The opposite holds true for an implosive exchange rate path. However, in linear models an implosive or deflationary price level bubble is not possible under free disposal (this will be shown below). Also, in general equilibrium monetary models (CAPM) implosive price level bubble paths are not sustainable equilibria under the assumption of finite endowments [see Kirikos (1991)]. Thus, models that either use PPP as a building block or arrive at a PPP-type relationship can be characterized by divergent exchange rate bubbles if and only if the price level of the relevant depreciating currency follows an inflationary bubble path.

The following example shows that divergent price level bubbles are theoretically legitimate paths in linear rational expectations models. Consider the money-market equilibrium condition:

$$m_t - p_t = k_t - \alpha\,(E_t p_{t+1} - p_t), \qquad \alpha > 0; \tag{2}$$

where m_t is the logarithm of the nominal money supply, p_t is the logarithm of the price level, $E_t p_{t+1}$ is the expectation of the logarithm of next period's price level conditional on information available at time t, and k_t represents variables, other than expected inflation, which affect the demand for real money balances. Equation (2) represents a Cagan-type function of the demand for real money balances [see Cagan (1956)]. Assuming that expectations up to time T only affect the current price level, rearrangement and forward iteration of equation (2) gives:

$$p_t = \left[\frac{1}{1+\alpha}\right] \sum_{j=0}^{T-t-1} \left[\frac{\alpha}{1+\alpha}\right]^j E_t\,(m_{t+j} - k_{t+j}) + \left[\frac{\alpha}{1+\alpha}\right]^{T-t} E_t\, p_T \tag{3}$$

Then, if the transversality condition:

$$\lim_{T \to \infty} \left[\frac{\alpha}{1+\alpha}\right]^{T-t} E_t\, p_T = 0 \tag{4}$$

holds, the market fundamentals solution for the price level is given by:

$$p_t^f = \left[\frac{1}{1+\alpha}\right] \sum_{j=0}^{\infty} \left[\frac{\alpha}{1+\alpha}\right]^j E_t\,(m_{t+j} - k_{k+j}) \tag{5}$$

where p_t^f depends only on the expected future path of fundamental determinants. However, if the discounted expected terminal value of the price level does not vanish [i.e. the transversality condition, equation (4), is violated], equation (5) is not the only solution to equation (2). Another solution is:

$$p_t = \pi_t^f + B_t \tag{6}$$

where B_t is the bubble component representing deviations of the current price level from its fundamental value. Clearly, equation (6) satisfies equation (2) as long as the bubble component has the following property:

$$B_t = \left[\frac{\alpha}{1+\alpha}\right] E_t B_{t+1} . \tag{7}$$

Letting $\delta_{t+1} = B_{t+1}\text{-}E_t B_{t+1}$ denote the price level bubble innovation at time t+1, we can write equation (7) as:

$$B_{t+1} = (1 + 1/\alpha)\, B_t + \delta_{t+1} \tag{8}$$

The homogeneous solution to the stochastic difference equation (8) is $B_t = (1 + 1/\alpha)^t B_0$, where B_0 is the starting value of the bubble. Another solution obtains if we iterate equation (8) backwards. Thus, the general solution for the bubble component is:

$$B_t = (1 + 1/\alpha)^t\, B_0 + \sum_{i=0}^{t} (1 + 1/\alpha)^{t-i}\, \delta_i \tag{9}$$

Then, in the presence of a bubble, the solution for p_t obtains by substitution of equations (5) and (9) into equation (6).

The validity of the above model is predicated on the assumption of a well-defined demand for real balances, which implies that agents must not expect the real value of their money holdings to increase without bound. Then, as we can see from equation (6) and (7), a negative price level bubble cannot exist since otherwise the price level would be expected to become negative in finite time. Diba and Grossman (1988) have pointed out that the exclusion of negative bubbles and the zero-mean property of the bubble innovation imply $\delta_{t+1}= 0$ with probability one whenever $B_t= 0$. This means that a bubble cannot start in period t+1 if there is no bubble in period t. In addition, the above model implies that if the bubble crashes, a new bubble cannot start concurrently or in a later period[6].

In linear asset market models agents' decision rules turn out to be functions of the first moments of the conditional distribution of fundamental determinants. However, if asset price volatility is to be of concern, agents' decision rules should be expected to have second moments as arguments. Such rules emerge from intertemporal optimization problems of rational risk-averse agents. Depending on the transaction technology, monetary equilibria may be characterized by divergent price level bubbles in these models [see Obstfeld and Rogoff (1983, 1986), Singleton (1987), Kirikos (1991)]. Thus, price level bubble equilibria are possible in a more complex general equilibrium setting and, as equation (1) shows, they must be associated with divergent exchange rate bubbles. The link between price level bubbles and exchange rate bubbles provides the basis of the bubble tests that we discuss in the next section.

3. Testing for Bubbles

The task of uncovering rational explosive bubbles presents particular difficulties due to the fact that a bubble component cannot be observed. Indeed, what may appear to be a bubble on the basis of information on asset prices and observed variables of the fundamental component may in fact be a path generated by anticipated fundamental innovations that are observed and accounted for by market participants but that cannot be seen by an econometrician. More technically, Hamilton and Whiteman (1985) have shown that solutions to linear rational expectations models in terms of fundamental innovations and bubble solutions are observationally equivalent and that bubble tests impose untestable restrictions on the dynamics of omitted fundamental variables. If such restrictions are relaxed, the only statistically valid test for bubbles is a test of stationarity for the asset price series and the underlying fundamentals.

The tests conducted in this paper are based on Diba and Grossman's (1984) characterization of a divergent bubble. Specifically, they define an asset price path as an explosive bubble whenever the order of integration (i.e. the number of times a series must be differenced to induce stationarity[7]) of the relevant asset price series exceeds that of the underlying fundamentals. Thus, if a bubble does not exist, the order of integration of the price series is equal to the order of integration of the driving exogenous variables. However, if a price series does not exhibit stationary behavior after differencing as many times as necessary to induce stationarity of the underlying fundamentals, we cannot conclude that a bubble is present. Alternative interpretations include a misspecification of the underlying model, "peso problem" effects, a bubble path, irrationality of expectations, and the failure of a small sample to reveal the actual order of integration of the asset price series. Clearly, the hypothesis to be tested here is the no bubbles hypothesis because the observable implications of a bubble are not unique. Moreover, the test is appropriate for testing the relevance of explosive indeterminacies. Non-explosive bubbles [see Miller and Weller (1990)] generate stationary deviations of asset prices from their fundamental values and thus they cannot be detected by a stationarity test (i.e. the test has no power in the case of non-explosive bubbles).

Our tests for explosive exchange rate bubbles are indirect in the sense that we do not use a particular model of exchange rate determination. Instead, we employ the stationarity test to examine the relevance of price level bubbles for the underlying currencies. Given the assumption of PPP (e.g. linear monetary models) or a PPP-type relationship [see equation (1)], evidence of explosive price level bubbles should be taken as evidence of explosive exchange rate bubble paths. Thus, not only are our tests general enough to include a variety of models but also they are not conditional on the validity of a particular exchange rate model. The latter reduces the number of maintained hypotheses that such a dependence entails.

In testing for price level bubbles, we compare the orders of integration of the (logarithm of the) price level (CPI) and the (logarithm of the) money stock (Ml) series for the United States, Germany, and the United Kingdon. If the two series are integrated of the same order, we conclude that the price level does not exhibit explosive bubble behavior and thus the relevant exchange rate is not on a bubble path. While the use of a reduced-form equation for the price level, with the only pre-determined variable being the money stock series, does not affect the ability of the test to provide evidence against the presence of bubbles[8], a rejection of the no bubbles hypothesis may be due to nonstationary omitted fundamental variables and therefore it does not necessarily mean that bubbles exist.

To determine the order of integration of a series, we examine plots of the series against time and of its sample autocorrelation function. A stationary series does not fluctuate extensively and its sample autocorrelation function dies out quickly. However, here we also conduct statistical tests based on the work of Dickey and Fuller (1979). Specifically, given that a nonstationary time series has one or more unit roots in the lag polynomial of its autoregressive (AR) representation, a stationarity test is equivalent to testing for a unit root in the AR polynomial. The Dickey-Fuller test for a unit root extends as follows. Let the series Y_t have the representation:

$$Y_t = \rho Y_{t-1} + \varepsilon_t, \tag{10}$$

where $Y_0 = 0$ and $\varepsilon_t \sim NID(0, \sigma^2)$. If $|\rho| < 1$, Y_t is integrated of order zero, denoted I(0). If $|\rho| = 1$, Y_t is I(1). Finally, for $|\rho| > 1$ Y_t is not stationary and its variance grows exponentially with time. The limiting distribution of the Ordinary Least Squares (OLS) estimator of ρ and of the corresponding t-ratio under the hypothesis $|\rho| = 1$ are derived in Dickey and Fuller (1979) and Tables of the percentiles of the distributions can be found in Fuller (1976). The same limiting distributions apply if ε_t is generated by a stationary autoregressive process of order two or higher [see Dickey and Fuller (1981)]. Suppose, for example, that ε_t follows the p-th order autoregressive process of order two or higher [see Dickey and Fuller (1981)]. Suppose, for example, that ε_t follows the p-th order autoregressive process:

$$\varepsilon_t = \varphi_1 \varepsilon_{t-1} + \ldots + \varphi_p \varepsilon_{t-p} + \eta_t, \tag{11}$$

where $\eta_t \sim NID(0, \sigma^2)$. Then under the hypothesis that $\rho = 1$, equations (10) and (11) give:

$$\Delta Y_t = \gamma Y_{t-1} + \sum_{i=0}^{\tau} \varphi_i \Delta Y_{t-i} + \eta_t, \tag{12}$$

where $\gamma = \rho - 1$ and Δ denotes a differencing factor. The fact that the OLS estimators of ρ-1 and γ in equations (10) and (12) have the same limiting distributions allows us to use equation (12) in testing the null hypothesis $H_0 : \rho = 1$ or, equivalently, $H_0 : Y_t$ is I(1). More precisely, we estimate equation (12) by OLS and calculate the usual t-ratio corresponding to the estimator of γ. If the alternative hypothesis is that Y_t is I(0), H_0 is rejected if the estimate of γ is negative and significantly different from zero. Note that the lag length in equation (12) is selected so that the residuals are approximately white noise.

A supplementary test for bubbles is a cointegration test. The concept of cointegration of economic variables is closely related to equilibrium relationships in the sense that a group of variables may move in a similar fashion under the pressure of market forces so that they do not drift apart even though each one may fluctuate extensively over time. For example two I(1) series will move around extensively, yet they may move together, that is, they stay close to each other over time. Such variables are said to be cointegrated.

Formally, two I(1) series are cointegrated if there is a linear combination of them that is I(0). (The parameter of the linear combination is called cointegrating parameter). Thus, if the price level and its market fundamentals are cointegrated, bubbles cannot be present. However, acceptance of the no cointegration hypothesis may or may not be due to bubbles for the same reasons that a stationarity test does not provide sufficient evidence of bubbles.

At first glance, it appears that stationarity and cointegration tests "contradict" each other since two or more cointegrated variables must be integrated of the same order in the first place. However, given the low power of unit root tests against borderline alternatives[9], a cointegration test appears to be a natural supplement to a stationarity bubble test. If, for example, a stationarity test shows that the price level and the money stock series are both I(1) but a bubble exists, i.e. the price level is actually I(d), d>1, then we should expect a cointegration test to provide evidence against cointegration of the two series.

Engle and Granger (1987) propose 7 different test statistics for testing the no cointegration hypothesis and give tables with critical values generated by a simulation study with samples of size 100. The suggested test, on the basis of the stability of its critical values, is an augmented Dickey-Fuller test whose steps are the following. Suppose that each one of the series Y_t and X_t is I(1). We form the cointegrating regression:

$$Y_t = \alpha + \beta X_t + \varepsilon_t \tag{13}$$

and by using OLS we take an estimate of the cointegrating parameter β. (Note that β can be estimated by OLS only if Y_t and X_t are cointegrated, otherwise ε_t will be I(1) and thus have a very large variance). The next step is to test whether the residuals from equation (13) are I(0) or not. To do this, we estimate the following augmented Dickey-Fuller regression:

$$\Delta e_t = \gamma e_{t-1} + \sum_{i=1}^{\pi} \varphi_i \Delta e_{t-i} + \eta_t, \tag{14}$$

where e_t denotes the OLS residuals from equation (13). Our test statistic is the usual t-ratio corresponding to the estimator of γ in equation (14). This statistic does not have a t-distribution under the hypothesis that Y_t and X_t are not cointegrated or, equivalently, that the error term ε_t has a unit root. Critical values of the limiting distribution of the t-ratio are given in Engle and Granger (1987) but they should be used with caution because they have been constructed for a sample size of 100. In addition, the power of the augmented Dickey-Fuller test, though higher than that of the simple Dickey-Fuller unit root test, diminishes fast for roots close to one[10]. Therefore, borderline results may not be reliable, especially when the cointegrating coefficient is estimated by OLS which minimizes the variance of the residuals and makes it likely that the residual series is stationary.

4. Bubbles in the \$/DM and the \$/Pound Exchange Rates?

Our monthly data are the logarithms of the Consumer Price Index (CPI) and the Money Stock (M1) series of the United States (US), Germany (G), and the United Kingdom (UK) over the period January 1974 - December 1987. The levels of the series are taken from various issues of the *International Financial Statistics.*

We employ the following notation: P, M, and E denote the CPI, M1, and residual series, respectively, while superscripts (US, G, UK) denote the country. A differencing factor Δ^d precedes any differenced series, the exponent d being the order of differencing. Finally, the negative number in parentheses denotes the

TABLE 1
SAMPLE AUTOCORRELATIONS OF P AND M SERIES

	Lag								
	1	2	3	4	5	6	7	8	9
P^{US}	.983	.967	.951	.935	.919	.903	.887	.872	.856
ΔP^{US}	.686	.549	.491	.440	.389	.335	.389	.406	.467
$\Delta^2 P^{US}$	-.290	-.095	-.037	.021	-.009	-.188	.050	-.042	.138
M^{US}	.978	.953	.924	.894	.864	.832	.799	.763	.726
ΔM^{US}	.491	.311	.314	.131	.194	.173	.185	.265	.279
$\Delta^2 M^{US}$	-.324	-.172	.171	-.235	.081	-.040	-.062	.070	.150
P^{G}	.984	.969	.953	.937	.922	.906	.890	.874	.858
ΔP^{G}	.463	.295	.273	.210	.150	-.014	.134	.208	.220
$\Delta^2 P^{G}$	-.322	-.159	.035	.008	.108	-.295	.066	.043	-.010
M^{G}	.969	.935	.911	.888	.866	.844	.821	.796	.773
ΔM^{G}	-.039	-.386	-.068	-.070	.115	.132	.087	-.068	-.082
$\Delta^2 M^{G}$	-.322	-.237	.155	-.095	.083	.033	.052	-.068	.141
P^{UK}	.980	.961	.941	.922	.903	.884	.865	.845	.825
ΔP^{UK}	.455	.327	.321	.164	.249	.348	.235	.140	.225
$\Delta^2 P^{UK}$	-.373	-.136	.154	-.222	-.015	.206	-.018	-.180	.149
M^{UK}	.978	.956	.934	.912	.891	.869	.847	.825	.804
ΔM^{UK}	-.167	-.107	-.135	.154	.000	-.085	.091	.064	-.030
$\Delta^2 M^{UK}$	-.524	.046	-.147	.194	-.027	-.115	.089	.032	.028

order of lag. For example, P^{US} stands for the logarithm of the U.S. price level and Δ^2P^{UK} (-1) stands for the second differences of the logarithm of the U.K. price level lagged one period. Since all variables are in logarithms (except for residuals), we drop the term logarithm in the following discussion.

Table 1 reports the sample autocorrelations of the CPI and M1 series as well as of their difference. The pattern of the sample autocorrelations shows that the first differences of all series (with the possible exception of ΔP^{US}) may be borderline stationary. To conduct Dickey-Fuller tests, we estimated regressions of the form given in equation (12) for all series. All of them gave positive estimates of γ which is the wrong sign for the series to be stationary in the levels. In Table 2 we report estimates of the regression:

$$\Delta^2 Y^t = \beta \Delta Y_{t-1} + \sum_{i=1}^{p} \theta_i \Delta^2 Y_{t-i} + \varepsilon_t, \tag{15}$$

for all series. (Note that a constant and/or a trend are not included because they were not found to be significant). The t-ratio corresponding to the coefficient of the lagged first differences is the Dickey-Fuller test statistic for testing the hypothesis that ΔY_t has a unit root or, equivalenty, that Y_t is I(2). The critical values are approximately -1.95 and -2.60 at the 5% and 1% significance levels, respectively [see Fuller (1976, p. 373)]. Thus, according to the t-ratios reported in Table 2, we have a borderline rejection of the unit root hypothesis for the series ΔP^{US} [i.e. the series is I(1) in the levels] at the 5% significance level but not at the 1% significance level. All other series turn out to be I(1) in the levels at the 1% significance level. While these results show that explosive bubbles were not present in the price level of Germany and the United Kingdom, the borderline rejection for ΔP^{US} casts doubts on the evidence against divergent bubbles in the U.S. price level (see footnote 9).

The results of cointegration tests are reported in Table 3. The augmented Dickey-Fuller test statistic for testing the hypothesis that the series of the residuals (E) from the cointegrating regression has a uni root, i.e. the no cointegration hypothesis, is again the t-ratio corresponding to the past level of the residuals. The critical value at the 5% significance level is approximately -3.17 [see Engle and Granger (1987, p. 270)]. Thus, on the basis of our estimates, we accept the no cointegration hypothesis for all three countries. Apparently, the lack of cointegration between the CPI and the M1 series shows that bubbles might have been present in the price level of the U.S., Germany, and the U.K. and, in effect, in the relevant exchange rates over our sample period.

TABLE 2
DICKEY - FULLER REGRESSIONS

	Dependent Variable					
	Δ^2P^{US}	Δ^2M^{US}	Δ^2P^{G}	Δ^2M^{G}	Δ^2P^{UK}	Δ^2M^{UK}
Regressor						
ΔP^{US} (-1)	-.065 (-1.96)					
Δ^2P^{US} (-1)	-.313 (-4.07)					
Δ^2P^{US} (-2)	-.176 (-2.33)					
ΔM^{US} (-1)		-.373 (-4.40)				
Δ^2M^{US} (-1)		-.195 (-2.21)				
Δ^2M^{US} (-2)		-.196 (-2.25)				
ΔP^{G} (-1)			-1.72 (-3.04)			
Δ^2P^{G} (-1)			-.299 (-.3.69)			
Δ^2P^{G} (-2)			-.299 (-3.09)			
ΔM^{G} (-1)				-1.32 (-12.7)		
Δ^2M^{G} (-1)				.341 (4.52)		
ΔP^{UK} (-1)					-.157 (-3.01)	

TABLE 2 (continued)

	Dependent Variable					
	Δ^2P^{US}	Δ^2M^{US}	Δ^2P^{G}	Δ^2M^{G}	Δ^2P^{UK}	Δ^2M^{UK}
Δ^2P^{UK} ↗½∺↘					-.375 (-4.79)	
$^{D2}P^{UK}$ (-2)					-.259 (-3.55)	
ΔM^{UK} (-1)						.486 (-3.79)
Δ^2M^{UK} (-1)						-.473 (-4.00)
Δ^2M^{UK} (-2)						-.381 (-3.71)
Δ^2M^{UK} (-3)						-.316 (-4.24)

Note: Numbers in parentheses below the estimated coefficients are t-ratios.

TABLE 3

COINTEGRATING REGRESSIONS AND AUGMENTED DICKEY-FULLER REGRESSIONS

	Dependent Variable					
	P^{US}	P^{G}	P^{UK}	ΔE^{US}	ΔE^{G}	ΔE^{UK}
Regressors						
Constant	.925 (.618)	1.59 (21.3)	2.07 (33.7)			
M^{US}	.585 (2.44)					
M^{G}		.558 (40.5)				
M^{UK}			.717 (40.6)			
E^{US} (-1)				-.002 (-1.53)		
ΔE^{US} (-1)				.603 (7.81)		
ΔE^{US} (-2)				.170 (2.23)		
E^{G} (-1)					-.356 (-1.25)	
ΔE^{G} (-1)					.041 (.545)	
ΔE^{G} (-2)					-.357 (-4.71)	
E^{UK} (-1)						-.007 (-.618)

Note: Numbers in parentheses below the estimated coefficients are t-ratios.

Oddly enough, the strong evidence against the no bubbles hypothesis provided by the cointegration test is in sharp contrast with the weak evidence in favor of the same hypothesis given by the stationarity test. Although one might be tempted to place greater reliance on the results of the cointegration test on the basis of power considerations (see footnotes 9, 10), it would be misleading to conclude that the no bubbles hypothesis is not supported by the data. Shortcomings and possible extensions of our tests are discussed in the following section.

5. Concluding Remarks

The existence of rational explosive price level bubbles and exchange rate bubbles has important implications for market efficiency [see Kirikos (1991)], macro policy [see Dornbusch (1982)], and the ability of bubble-augmented structural exchange rate models to explain the observed variability of exchange rates.

The results of the bubble tests conducted in this paper are rather supportive of the possibility of bubbles in the dollar/deutschemark and the dollar/pound exchange rates over the recent free floating period. However, the presence of bubbles is not substantiated because our tests cannot provide sufficient evidence of it.

Our findings should be interpreted with caution for several reasons. First, the ability of unit root tests to discriminate between borderline stationary alternatives has been questioned [see Christiano and Eichenbaum (1990) and Cochrane (1991)]. Second, we conduct the unit root tests under the arbitrary assumption that the inflation rate and the rate of monetry growth are satisfactorily represented by low order autoregressive processes instead of general autoregressive moving average (ARMA) processes[11]. Third, the stability of our test statistic (for the stationarity test) in empirical power studies is not satisfactory [see Dickey, Bell and Miller (1986)]. Fourth, the assumption of constant parameters for the cointegrating regressions may be misleading [see Canarella, Pollard and Lai (1990)]. Future research will address these issues.

Footnotes

1. Asset market models include the simple monetary model, Dornbusch's (1976) fix-price model, Frankel's (1979) real interest rate differential model, currency substitution models, and the portfolio balance model. For a review of these models see Baillie and McMahon (1989, chapter 3). For a description of the asset market approach to exchange rate determination see Mussa (1979).

2. Kirikos (1991) argues that as long as agents' expectations are validated *ex ante* returns to uncovered speculation in the foreign exchange market are zero in the presence of bubbles. However, the possibility of a bubble crash introduces non-zero ex ante returns which imply a misallocation of resources.

3. Theory, econometric evidence, and bubble interpretations of specific events are presented in a Symposium on Bubbles in the Spring 1990 issue of the *Journal of Economic Perspectives.*

4. For a review of the empirical evidence on explosive bubbles and a *discussion of some methodological problems see Flood and Hodrick (1990).*

5. An assessment of the empirical literature on bubbles is presented in West (1988)

6. For Etδt+1 = 0 and B_t*+1>0, equation (8) implies Pr (*δ_t*+i=0) - 1 when* B_t*= 0. Thus, a bubble cannot burst and restart in the same period. Also, for* $B_{t=}$ *0 and* $B_{t+1} >$*0 (i>0)),* $E_t B_{t+1}$ *= (1+1/α)'B,= 0 and thus Pr(*B_{t+i} *= 0; ï>0) = 1, that is, if a bubble bursts, a new bubble cannot start in a later period.*

7. We use the term stationarity in a second order weak sense or covariance stationary sense [see Wei (1990, chapter 2)].

8. The order of integration of omitted fundmental variables cannot exceed that of the price level.

9. A Monte Carlo study, reported in Dickey and Fuller (1979, p. 430), shows that for a sample size of 100 and values of ρ [see equation (10)]. 95, .99 and 1.02 the power of the t-ratio test is .17, .04, and .59, respectively.

10. Engle and Granger (1987, p. 270) estimated the power of the augmented Dickey-Fuller test to be .61 and .90 for a 4th order autoregressive process and roots of the residual series equal to .90 ad .80, respectively.

11. Said and Dickey (1985) and Said (1991) have extended the unit root tests to general ARMA representations.

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