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Exchange Rate "Fundamentals" versus Speculation: The Case of Lebanon

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Abstract

The exchange rate for the Lebanese pound experienced a protracted period of depreciation from end-1982 to November 1987, followed by a marked appreciation over the following six months. This paper investigates the competing hypotheses that the exchange rate over these two periods was driven by a speculative bubble versus "fundamental" economic variables. Reduced-form and time series models for the exchange rate are estimated and tested for nonstationarity. The results of these test suggest that the pound's volatility in recent years was consistent with an excessive growth in domestic versus foreign currency denominated liquidity rather than speculation.

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I. Introduction

The exchange rate for the Lebanese pound has exhibited an extreme degree of volatility in recent years, undergoing an extended depreciation over the 1983-87 period followed by a substantial appreciation after November 1987. In particular, over the 1983-85 period the pound depreciated in nominal effective terms at an average annual rate of nearly 40 percent; in 1986, the rate of depreciation increased to some 80 percent; and, during the first eleven months of 1987 the pound depreciated by nearly 90 percent. Between November 1987 and June 1988, however, the pound strengthened dramatically, appreciating by some 50 percent in nominal effective terms, erasing a significant proportion of its previous weakness.

This paper examines the extent to which the behavior of the Lebanese exchange rate can be ascribed to underlying fundamental economic variables versus market psychology or speculation. Besides the armed hostilities that have been a constant factor in the Lebanese economic environment since 1975, a number of traditional economic developments were associated with the pound's depreciation, including a fiscal deficit that had grown to a substantial proportion of GDP, the accelerated growth of government interest- and non-interest-bearing debt (e.g., high-powered money), and intense inflationary pressures. However, since the appreciation of the pound after November 1987 was not accompanied by an apparent change in the trend of the aforementioned variables, the possibility suggests itself that the pound's weakness had been at least partly related to a "bubble" phenomenon.

A speculative "bubble," refers to an exchange rate (or any asset price) driven by self-fulfilling expectations rather than by the fundamental variables (or their expectations) that normally determine its value. The policy relevance of the distinction between exchange rates driven by fundamentals versus speculative forces is paramount. In the former case, the prescription for undesired appreciation or depreciation is to address the economic fundamental factors using policy tools under the monetary authorities' control. However, if the exchange rate is being driven by speculative pressures independent of the authorities' economic policies, the prescription may well be to adopt either a laissez-faire or interventionist approach in the exchange market, the choice depending on the perceived costs of responding to the speculative activity. It is to these competing hypotheses that this paper is addressed.

Section II reviews recent economic developments in Lebanon as they relate to the evolution of the exchange rate for the pound. Section III develops a structural model of the exchange rate and defines the concept of an asset price bubble. Section IV investigates the evidence for the existence of a bubble over the period leading up to November 1987 by testing the structural model introduced in the previous section and univariate and multivariate time-series models of the exchange rate. Section V discusses the resultant conclusions.

II. Recent Economic Developments 1/

Following the intensification of armed hostilities during 1975, the Lebanese pound underwent an extended period of severe depreciation, especially after 1982, depreciating by 99 percent from a level of US$1 = LL 3.84 in January 1983 to US$1 = LL 498.21 by November 1987. The pound underwent a remarkable turnaround thereafter, appreciating by nearly 40 percent to reach a level of US$1 = LL 355.96 by June 1988. The pound's movement in nominal effective terms was just as spectacular over the two subperiods, depreciating and subsequently appreciating by 99 percent and 51 percent, respectively.

Besides the psychological impact of the deteriorating political and security situation, the erosion of the Lebanese pound's exchange rate up to November 1987 was consistent with adverse economic developments. In particular, the Lebanese economy was characterized by an extended period period of declining real GDP, an extremely rapid growth in the Government's fiscal deficit and domestic financing requirements, especially as a share of total GDP, and relatively low domestic interest rates. These factors contributed to exceptional increases in liquidity and inflation, lending further impetus to the depreciation of the pound.

Lebanon's real GDP, which had grown by some 6 percent annually over 1965-75, fell at an average annual rate of 2 percent during 1975-81. 2/ Since then, output is estimated to have declined further to 1986 but may have increased somewhat in 1987. The Government's fiscal position had similarly eroded since 1975. Annual expenditure growth is estimated at 5 percent in real terms over 1975-85, owing to subsidies of staples and petroleum products, growing defense expenditures, and expanding expenditures on reconstruction. The authorities' attempt to maintain public service employment and real wages in the face of recession and inflation and the increased interest burden on public debt were also contributory to expenditure growth. Conversely, revenues declined owing to the effect of the security situation on overall economic activity, the tax base, and tax collection. 3/ As a result, while having exhibited near-budget balance in 1974, the deficit grew to over 25 percent of GDP in 1982 and is estimated to have reached nearly

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1/ A difficulty in describing economic events in Lebanon is the dearth of data. Much of the following discussion relies heavily on details provided by Saidi (1986) and Makdisi (1987), as well as IMF staff estimates.

2/ Both Saidi and Makdisi explain the decline in output in terms of the destruction of productive capital and vital infrastructure, the outward migration of the skilled workforce, sharply curtailed private investment activity, and limitations to the mobility of both goods and services.

3/ For example, customs duty collection was hampered by smuggling and the lack of government control over key ports. Further, a number of indirect taxes were on a specific (versus an ad valorum) basis and lagged significantly behind the inflation rate.

50 percent of GDP in 1985. In real terms, the deficit is estimated to have grown by roughly 70 percent from 1982 to 1985, but may have fallen significantly in 1986 and 1987, primarily owing to reduced domestic subsidies on petroleum products.

Since the bulk of the fiscal deficit was financed internally, significant inflation pressures resulted. Over 90 percent of the budget deficit was financed by the domestic commercial banking sector over 1982-87, in large part through sales of Treasury bills to these institutions induced by frequent adjustments to their secondary reserve requirements. The Central Bank also financed a major proportion of the budget deficits through purchases of government bonds and advances, as well as, at times, by crediting the Treasury's account with some of the revaluation profits from the Bank's foreign asset position. As a result, the rate of domestic currency-denominated liquidity growth averaged 33 percent over 1982-85, fell slightly to 29 percent in 1986, and accelerated to 48 percent in 1987. 1/

These developments contributed to a reduction in the private sector's desire to hold real pound-denominated liquid balances, and to a resultant acceleration in the rate of price inflation and exchange rate depreciation. During 1982 the inflation rate was estimated at 14 percent while the average U.S. dollar/pound exchange rate fell by 9 percent. During 1987, inflation was estimated to have reached as high as 700 percent, well in excess of the growth of nominal pound-denominated liquidity, while the average exchange rate depreciated in nominal effective terms by 88 percent over January-November.

Thereafter, however, the exchange rate rate exhibited a remarkable turnaround, appreciating by over 50 percent, with a concurrent easing of inflationary pressures. Nonetheless, there did not appear to have been a concomitant reversal in the trend of those economic factors which had been associated with the pound's earlier depreciation. For example, the rate of domestic currency-denominated liquidity growth accelerated markedly over the November 1987-March 1988 period as the Central Bank intervened heavily to replenish its foreign exchange reserves. Further, while the fiscal deficit has declined in real terms, it is expected to remain a substantial proportion of domestic output. The obvious question is to what extent the intensification of exchange rate pressures during 1986 and 1987, and therefore the reversal of the pressure after November 1987,can be explained in terms of such "fundamental" economic factors versus speculative pressures.

1/ Concomitantly, partly owing to the relatively low real rate of return on domestic savings instruments, the share of the banking sector's balance sheet devoted to foreign currency-denominated assets and liabilities rose dramatically. For example, in 1982 the foreign currency share of private sector liquidity was 24 percent, while in 1986 the share rose to 68 percent. As the pound's rate of depreciation accelerated, the foreign currency share of liquidity rose to nearly 90 percent by end-1987.

III. "Fundamentals" versus Speculative Bubbles

The task of distinguishing between fundamental versus bubble determinants of the exchange rate is complicated by the lack of consensus as to which fundamental relationships, and therefore which fundamental economic variables, are at work in foreign exchange markets (see Boughton (1987) for a discussion of the various alternatives). 1/ In the analysis below it will be assumed that the predominant determinant of the Lebanese exchange rate over recent years was the activity in its financial market, including, as described above, the dollarization of the banking sector's liabilities, and the extreme pressure on the financial market to finance the Government's deficit. Thus, a hybrid version of the monetary-portfolio-balance model proposed by Hooper and Morten (1982) and Frankel (1983) is adopted. 2/ In this case, it is assumed that it is Lebanese residents' attempts to alter the currency composition of their portfolios that is the predominant factor determining the exchange rate. To account for the effect of such portfolio decisions in the exchange market, a currency substitution model of money demand (for recent examples of empirical work in this area see El-Erian (1988), Poloz (1986), and Ramirez-Rojas (1985)) is proposed in which an important role is assumed for residents' demand for foreign currency deposits and Treasury bills.

A simple three-asset model of Lebanon's financial markets is, in log form, as follows:

$$m^d(t)-p(t) = l[r(t), r^*(t), E[e(t+1)]-e(t), r^{tb}(t)] \qquad (1)$$

$$m^{*d}(t)+e(t)-p(t) = l^*[r(t), r^*(t), E[e(t+1)]-e(t), r^{tb}(t)] \qquad (2)$$

$$tb^d(t)-p(t) = b[r(t), r^*(t), E[e(t+1)]-e(t), r^{tb}(t)] \qquad (3)$$

where

$m^d(t)$ = the log of desired nominal domestic liquidity held by residents at t,
$m^{*d}(t)$ = the log of desired nominal foreign currency liquidity, denominated in the foreign currency, held by residents at t,
$tb^d(t)$ = the log of desired nominal Lebanese Treasury bill holdings of non-bank residents at t,
$p(t)$ = the log of the domestic price level at t,

1/ Further, it has been demonstrated (Meese and Rogoff (1983)) that many empirical models cannot outperform simple autoregressive processes in out-of-sample forecasts.

2/ For an example of an application of the pure monetary approach to Lebanon see Spitaller (1980).

$r(t)$ = the nominal interest rate on domestic currency liquidity at t,
$r^*(t)$ = the nominal interest rate on foreign currency liquidity at t,
$r^{tb}(t)$ = the nominal interest rate on treasury bills at t,
$e(t)$ = the log of the domestic currency price of foreign exchange at t, and
$E[\,]$ = the expectations operator.

Equations (1)-(3) represent fairly standard asset-demand equations for domestic currency liquidity, foreign currency liquidity, and domestic treasury bills, respectively. The specification of the response to foreign interest rate shocks is extremely general. It is assumed that the response of asset demands to changes in uncovered foreign exchange rates may differ from that for domestic or foreign interest rates. This is primarily to accommodate data limitations with regard to foreign interest rates, but may also capture possible behavioral asymmetries owing to risk and other considerations. To satisfy standard adding-up constraints, the demand functions must, in general, include as arguments the returns of each asset in investors' choice set (see, for example, Roley (1977) and Niehans (1978)). For example, residents' demand for liquid balances will be a function of the rate of return on those balances, represented by the domestic interest rate r, and one would expect that, ceteris paribus, the demand for domestic liquidity would increase with an increase in those balances. 1/

However, as many authors have noted, there is little that theory can say regarding the response of the individual asset demands to anything but the own rate of interest, since the cross-rate effects are complicated functions of income and substitution effects. Thus, the effect of an increase in the interest rate on foreign currency denominated balances, r^*, the rate on treasury bills, r^{tb}, or the rate of depreciation of the pound, $E[e(t+1)]-e(t)$, on the demand for domestic liquidity is unknown. 2/ Similarly, it would be expected that the demand for Treasury bills and foreign currency balances would be an

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1/ Ideally, the rate of price inflation should be included in the set of asset returns. However, since it is assumed that it is the relative returns on the assets which determines their demand and that r(t) sufficiently proxies the opportunity cost of domestic currency liquidity, the inflation rate is not included. This choice was confirmed by a lack of statistical significance upon inclusion of the inflation rate in the reduced form equation estimates discussed below.

2/ However, in general the wealth constraint implies cross-equation restrictions on the partial derivatives of the asset-demand equations (that they sum to zero). However, since data restrictions imply that the discussion be limited to only a subset of assets available to Lebanese residents, these restrictions are not considered. For the same reason, wealth is not included as an argument of the asset-demand functions.

increasing function of the interest rate on those assets, r^{tb} and r^{*}-E[e(t+1)]+e, respectively, while the cross-rate effects are unknown.

Assuming that equations (1)-(3) can be linearized, they may be expressed in matrix form:

$$\begin{bmatrix} m^{d}(t) \\ m^{*d}(t) \\ tb^{d}(t) \end{bmatrix} - \begin{bmatrix} 0 & 1 & 0 \\ -1 & 1 & 0 \\ 0 & 1 & 0 \end{bmatrix} \begin{bmatrix} e(t) \\ p(t) \\ r(t) \end{bmatrix}$$

$$= B\begin{bmatrix} r^{tb}(t) \\ r^{*}(r(t) \end{bmatrix} + \begin{bmatrix} c_{11} & 0 & 0 \\ c_{21} & 0 & 0 \\ c_{31} & 0 & 0 \end{bmatrix} \begin{bmatrix} E[e(t+1)] \\ E[p(t+1)] \\ E[r(t+1)] \end{bmatrix} \begin{matrix} - \\ - \\ - \end{matrix} \begin{bmatrix} c_{11} & 0 & d_{13} \\ c_{21} & 0 & d_{23} \\ c_{31} & 0 & d_{33} \end{bmatrix} \begin{bmatrix} e(t) \\ p(t) \\ r(t) \end{bmatrix} \qquad (4)$$

Or, more compactly,

$$x^{d}(t)-Ay(t) = Br(t) + CE[y(t+1)] - Dy(t)$$

where $x^{d\prime}=(m^{d}, m^{*d}, tb^{d})$ is the vector of desired asset demands; $r'=(r^{tb}, r^{*})$ is the vector of "exogenous" asset returns; $y'=(e, p, r)$ is the vector of "endogenous" prices and returns; and A, B, C, and D are matrices. As in equations (1)-(3), the left-hand side of equation (4) represents the nominal value of the assets expressed in their original currency, x(t), adjusted by the current domestic price level, and the exchange rate where necessary, to yield an expression in terms of real asset values. Thus, the coefficients of the A matrix are unity along the second column, as applied to the price level, minus one as applied to the exchange rate in the case of foreign liquidity ($a_{21} = -1$), and zero elsewhere. The B matrix coefficients represent the asset demand's response to changes in the domestic bill and foreign deposit rate; the C matrix coefficients equal the asset demand's response to changes in the expected rate of depreciation (the first column). The coefficients in the first column of the D matrix also represent the response of asset demands to an expected depreciation (and thus are equal to the coefficients in the first column of the C matrix), while the coefficients in the third column represent the responsiveness of asset demands to the rate on domestic liquidity.

Note that the model described above only relates asset returns to desired asset holdings rather than to actual asset supplies. Further, the dynamics of the model are confined to the rate of expected appreciation. To close the model, it is assumed that owing to transactions or other costs (see Neihans (1978), Chapter 11 for a discussion), the private sector's portfolio may only be adjusted slowly in response to changes in asset returns for each period. Thus, the change in the investors' portfolios between the current and previous periods will only be a "fraction" of the difference between the desired change in asset stocks. The adjustment of each asset is assumed to be according to

$$[x(t)-Ay(t)] - [x(t-1)-Ay(t-1)]$$
$$= \Gamma\{[x^d(t)-Ay(t)] - [x(t-1)-Ay(t-1)]\} \qquad (5)$$

where the Γ matrix is the adjustment coefficient matrix relating the difference between actual and desired stocks of each asset j to the change in asset i. To the extent to which portfolios are adjusted instantaneously, Γ approaches the identity matrix and desired stocks equal actual stocks. As above, the wealth constraint will impose adding-up conditions on the adjustment coefficients. 1/

Solving for the y vector using equations (4) and (5) yields

$$y(t) = (\Gamma D-A)^{-1}[\Gamma CE[y(t+1)] - x(t)$$
$$+ \Gamma Br(t) + (I-\Gamma)(x(t-1)-Ay(t-1))] \qquad (6)$$

so that equation (6) represents three semi-reduced-form equations that relate the endogenously determined exchange rate, price level, and domestic interest rate on domestic currency liquidity to the expected future exchange rate and to the market "fundamentals", the domestic bill and foreign deposit rates, the current nominal asset stocks and the previous period's real asset stocks. The first equation of the system, which explains the evolution of the exchange rate, is

$$e(t) = \alpha E[e(t+1)] + \lambda Z(t) \qquad (7)$$

where Z(t) is the vector of fundamentals (as defined above) and α and λ are the coefficient and vector of coefficients, respectively, reflecting the responsiveness of the current exchange rate to changes in the expected rate and the fundamentals, respectively.

In the case where α is less than unity in absolute value, the familiar rational expectations solution is the forward solution

$$e(t) = \lambda Z(t) + E[\sum_{1}^{\infty}\alpha^{i}\lambda Z(t+i)] = f(t) \qquad (7')$$

where f(t) represents the "fundamental" exchange rate solution at time t, which depends solely on the current and expected future values of the market fundamentals Z(t).

1/ Clearly, a deficiency in this approach is the implicit assumption of the exogeneity of the three assets in question, i.e., there is no role for a behavioral response of the monetary authorities. However, to the extent that such a response was evident over the period in question equation (5) may be viewed as an amalgam of both the private sector's demand and the monetary authorities' supply response.

In contrast, however, it is easily demonstrated that this rational expectations solution is not unique. Alternate solutions exist of the form

$$e(t) = f(t) + b(t) \tag{7''}$$

where b(t) is defined as the bubble component of the exchange rate solution, which is required only to satisfy $E[b(t+1)] = \alpha^{-1}b(t)$ or that $b(t+1) = \alpha^{-1}b(t)+z(t+1)$ where $E[z(t+1)] = 0$. 1/ This is to say that, if market participants come to expect the exchange rate will exhibit an explosive episode independent of the fundamentals that had previously driven both the exchange rate and expectations, this new expectation could become self-fulfilling and not violate the (by now) usual assumptions of market efficiency (or rationality). 2/ While the magnitude of b(t) is unrestricted a priori, its expected future value must conform to the above first-order difference equation. Since the parameter α is less than unity, it is clear that the expected value of the bubble component evolves in an explosive manner. It is this characteristic of the bubble that results in an unstable time path of the asset price in question.

Unfortunately, little else can be said regarding the nature of the bubble path. Neither deterministic bubbles, ($z(t) = 0$ for all t), in which case the implicit assumption is that the market expects the bubble to last forever, nor stochastic bubbles, in which case the bubble may be expected to collapse, are ruled out by the assumptions above. Further, the distribution of the stochastic component may involve a complicated function of the size of the bubble. The example proposed by Blanchard and Watson (1982) is of a bubble that has a higher probability of bursting (going to zero) as it increases in size. While a number of attempts have been made to rule out the existence of bubble solutions by appealing to consumer theory, 3/ the consensus appears to be that bubbles may represent an important reason why fundamental models of

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1/ To confirm this assertion, simply substitute the solution for e(t+1) from equations (3) and (4) into equation (1).

2/ For a recent discussion and review this issue see Singleton (1987) and Flood (1987).

3/ For example, it has been argued that deflationary bubbles (in the case of price equilibria) imply negative prices (if e(t) is defined in levels), or an explosive future real value of nominal assets (if e(t) is defined in log form) contravening infinitely lived consumers' transversality conditions. Tirole (1985) resolves this problem by assuming away the existence of such price paths. Farmer (1984) has demonstrated, however, that explosive price bubbles may also be ruled out if consumers' expenditures are not completely divisible since the bubble may reduce real assets below a minimum level required to finance consumption. Further, Diba and Grossman (1988) argue that if the bubble is nondeterministic, by limiting the distribution of z(t+1) to be one-sided, the condition that $E[z(t+1)] = 0$ would be violated, in turn violating the assumption of rationality.

asset market prices have performed so poorly in empirical tests (see Singleton (1987) for a useful survey of this issue applied to exchange markets).

IV. The Evidence

Both univariate and multivariate tests of asset price bubbles have been undertaken. Univariate tests (such as those performed for exchange rates and stock prices by Meese (1986) and Diba and Grossman (1988a) respectively) involve making use of the fact that since the asset price is composed of both a fundamental and a bubble component (assuming that the latter exists), its time-series properties will be dependent on the time-series properties of both components. Since the bubble component evolves according to a nonstationary stochastic first-order autoregressive process, except in the unlikely event that the fundamental (as defined above) exchange rate is following an identical time path that exactly offsets the instability engendered by the bubble, a necessary condition for the existence of a bubble is that the exchange rate also exhibit the properties of a nonstationary stochastic process. Further, the nonstationarity cannot not be relieved through differencing; since $b(t) = \alpha^{-1}b(t-1)$ then $\Delta b(t) = \alpha^{-1}\Delta b(t-1)$.

Multivariate tests of asset price bubbles have generally followed one of two strategies; the first has been to explicitly include the bubble process in a structural/semi-reduced-form model of the asset price in question (for example, see Flood and Garber (1980) or Borensztein (1987)), while the second has been to estimate multivariate models without the inclusion of a bubble term and to examine the resultant parameter estimates for evidence of misspecification attributable to a bubble (for example, Hamilton and Whiteman (1985), Meese (1986), and West (1987)). The first strategy suffers from the requirement that both the structure of the fundamental relationships and the bubble process be specified upon estimation, so that a test of the bubble hypothesis is a test of the joint hypothesis regarding the model's structure. Rejection of the bubble hypothesis cannot be distinguished from rejection of the assumed structure, while acceptance of the bubble hypothesis cannot be distinguished from the possibility that the bubble proxies an excluded variable whose time series mimics that of the supposed bubble. 1/

1/ This is exactly the criticism raised by Hamilton and Whiteman (1985) and Hamilton (1986). They argue that the only possible test for rational bubbles is to adopt the strategy advocated by Diba and Grossman (1988a), to examine the stationarity properties of the asset price in question and the fundamental variables to determine whether differencing removes the nonstationarity of the fundamentals but not the asset price. However, as Hamilton notes, this does not remove the possibility that remaining asset price nonstationary may be due to a "peso problem," i.e., the expectation of a (possibly unrealized) catastrophic event or policy shift.

An alternate approach adopted by Meese (1986) and West (1987) has been, first, to estimate the structural equation explaining the exchange rate using McCallum's instrumental variable approach to substitute for the expected value of next period's asset price, which yields consistent parameter estimates even in the presence of a bubble term as long as the bubble is correlated with the instruments. Secondly, the equation is reestimated replacing the expected future value of the asset with its rational expectation (based on the structural model and the time-series process driving the fundamental variables), which will lead to inconsistent estimates if the exchange is being driven by a bubble. The test is to compare the parameter estimates under the two estimation procedures for consistency, using a form of a Hausman specification test. However, as West notes, this test procedure suffers from similar defects as the strategies described above -- that the joint hypothesis of the model's structure is also tested. 1/ As an adjunct to the above tests, Meese (1986) and Diba and Grossman (1988a), test for the cointegration of the asset price with the fundamentals. 2/ However, rejection of the hypothesis of cointegration cannot be distinguished from the rejection of the implicit assumption regarding the set of fundamentals included in the cointegrating equations.

Notwithstanding the difficulties described above, a structural model of the exchange rate will be estimated below, using the consistent-estimator approach suggested by McCallum. While no explicit bubble term will be included, examination of the reduced-form coefficients will provide informal evidence as to whether the forward solution to the exchange rate, a necessary condition for a bubble, is appropriate. Further, comparing the estimates over two subperiods delineated by the November 1987 turnaround in the exchange rate, and examining the residuals for a structural break, will provide some evidence for or against a bubble. Following a discussion of the reduced-form estimates, the univariate and multivariate time-series properties of the exchange rate and fundamental variables (as suggested by the previous analysis) will be examined, and the hypothesis of cointegration between the exchange rate and the fundamentals will be tested.

1/ Further, West also reports that the test statistic itself is not consistent and may tend to incorrectly reject the presence of bubbles.

2/ A vector of variables is defined to be cointegrated of order (d,b) if all the component series of the vector are integrated of order d, i.e., have a stationary ARMA representation after differencing d times, and there exists a nontrivial linear combination of the series that is integrated of order (d-b) (see Granger and Engle (1987)). The significance of the asset price series being cointegrated with the fundamentals is that, if the order (d-b) is sufficiently low, then the residuals of the regression of the asset price on the fundamental variables cannot contain a bubble component.

1. Reduced-form estimates

Given the zero restrictions on the C matrix in equation (6), it is clear that the only expected future variable that will appear in the reduced-form equations will be the expected exchange rate. Since it is the exchange rate's response to the fundamentals that is of interest here, only equation (7) was estimated. Because the expected exchange rate is unobservable, the realization of y(t+1) is substituted for its expected value and the equation parameters is estimated consistently using instrumental variable techniques (see Chow (1983), for a discussion). As West (1987) has noted, even in the presence of a bubble term in the expected exchange rate, this estimation technique provides consistent estimates of the coefficients, provided that the bubble is correlated with the instruments. 1/ The current exchange rate was initially regressed on the instrumented next-period exchange rate, the current domestic bill rate, the foreign deposit rate, current nominal (with foreign liquidity denominated in the foreign currency) asset stocks, and the previous period's real asset stocks. A parsimonious specification was then derived by testing the significance of the estimated coefficients.

Table 1 provides the coefficient estimates for equations explaining the pound per U.S. dollar (LL/$) and the nominal effective exchange rate (NEER). 2/ The R^2 coefficients and the standard errors (both not reported), which in the case of a log dependent variable proxies the average squared percentage error, were approximately unity and 0.05, respectively. As regards residual autocorrelation, the Durbin-Watson statistics were inconclusive but suggestive of a first-order process. However, since the lagged real foreign currency liquid asset variable contains the lagged dependent variable this test statistic is likely to be biased toward rejection of the hypothesis of autocorrelation. Lagrange multiplier tests for residual autocorrelation are also reported; these indicate that the null of first-order autocorrelation can be rejected at the 95 percent level for the nominal effective exchange rate but not for the LL/$ rate. Similarly, the null hypothesis of a fourth-order process cannot be rejected at the 95 percent level for the LL/$ model. Examination of the residuals indicated the possibility of increased error variance in 1986 and 1987. While Engle's ARCH (autoregressive conditional heteroscedasticity) test rejected the hypothesis of heteroscedasticity at the 95 percent level, the value of White's F test for heteroscedasticity (related to the levels and squares of the regressors) also indicated a misspecification of the reduced

1/ In each of the estimates reported, instruments included the other regressors in the equation and lagged values of the exchange rate.

2/ Initial estimates were performed for monthly data from April 1982 to September 1987 (using the computer program PC-GIVE 5.0) since the data for the interest rate rate series did not extend past that date.

Table 1. Reduced-Form Exchange Rate Estimates, 1982(4)-1987(9) 1/

Regressors	Dependent Variable: LL/US$ Rate		Dependent Variable: Nominal Effective Rate	
E[e(t+1)]	0.66 (12.78)	0.68 (13.06)	0.68 (19.45)	0.69 (20.71)
$r^{tb}(t)$	--	--	--	--
$r^*(t)$ 2/	--	--	--	--
m(t)	0.32 (5.12)	--	-0.26 (6.41)	-0.25 (7.14)
$m^*(t)$	-0.34 (6.07)	--	0.39 (7.87)	--
$m(t)-m^*(t)$	--	0.30 (5.47)	--	-0.39 (7.86)
tb(t)	--	--	--	--
m(t-1)-p(t-1)	-0.31 (4.97)	--	0.32 (6.74)	--
$m^*(t-1)+e(t-1)-p(t-1)$	0.34 (6.17)	--	-0.29 (6.79)	--
$m(t-1)-m^*(t-1)+e(t-1)$	--	-0.29 (5.05)	--	0.30 (7.76)
tb(t-1)-p(t-1)	--	--	--	--
Diagnostic Statistics 3/				
Sum of squared residuals	0.17	0.18	0.14	0.14
Durbin-Watson	2.26	2.10	2.28	2.27
AR(1)	5.32	.23	2.34	2.20
AR(4)	3.28	1.58	3.53	3.26
ARCH(4)	1.66	1.23	1.23	1.23
X^2	3.05	2.32	3.19	3.09
FCST/6	4.89	4.64	5.66	6.60

1/ Equations were estimated using instrumental variables estimates for the expected future value of the dependent variable E[x(t+1)]. In none of the four equations was the constant term significantly different from zero. Instruments used were the independent variables, any other current independent variables that were deleted owing to insignificance, and lagged values of the dependent variable. T-statistics are reported in brackets under the parameter estimates.

2/ The foreign interest rate was defined as the three-month rate on SDR deposits in the latter two equations.

3/ The AR(i) statistic is the Lagrange multiplier test of first to i'th order autocorrelation and is asymptotically distributed F(i,T-i). The ARCH(i) statistic tests for autoregressive conditional heteroscedasticity of order i and is distributed asymptotically F(i,57-i). The X^2 statistic tests for heteroscedasticity related to the squares of the regressors and is distributed F(n,T-2n), where n is the number of regressors. The FCST statistic tests the hypothesis of no structural change to the estimated equation over the forecast period, which, in this case, is 1987(10)-1988(3). It is distributed asymptotically chi-squared with N degrees of freedom, where N=6 is the number of forecast periods.

form. 1/ However, examination of estimates of White's heteroscedastic consistent standard errors, derived using the instrumented value of the expected exchange rate, suggested only a limited loss of efficiency.

As indicated in Table 1, the coefficients of the expected exchange rate were highly significant in both cases; further, the coefficient estimates were largely indistinguishable regardless of the definition of the exchange rate. Under most circumstances, the model formulation would predict that the coefficient be positive, since it is $E[e(t+1)]-e(t)$ that enters the asset demand equations; further, owing to the effect of the change in the current exchange rate on real foreign currency assets and the partial adjustment of asset demands, it would also be expected to be less than unity. For example, if the current and expected future exchange rate fell by the same amount, asset demands would be unchanged. However, the depreciation would imply an increase in the domestic currency value of foreign currency assets and, therefore, an excess supply of those assets. By reducing the current depreciation below that expected for the next period, the demand for foreign currency assets would be increased while reducing their supply. The effect of the partial adjustment would be to reduce the effect of such shocks on asset demands, further implying that the coefficient on the expected exchange rate would be less than unity.

The current nominal domestic and foreign-currency denominated money stocks were highly significant, as were the lagged real stocks. Further, the sign of the coefficient estimates accorded with intuition; the effect of an "exogenous" increase in domestic liquidity, given expectations, was to depreciate the exchange rate, while a similar increase in the foreign currency holdings of residents tended to cause the pound to appreciate. Taking account of the effect of changes in the money stock on the instrumented value of the expected exchange rate tended to damp the current exchange rate response somewhat, but left it positive. 2/ The current nominal Treasury bill stock and the previous month's real Treasury bill stock were not found to be significant, regardless of the exchange rate definition, implying that the response of asset demands to discrepancies between actual and desired Treasury bill stocks is weak.

Despite the significance of the expected future exchange rate, neither of the exogenous interest rate series were found to be significant determinants of the exchange rate, suggesting that only the rate of exchange rate appreciation and the interest rate on domestic

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1/ A possible source of the misspecification is the proxy for $E[e(t+1)]$. However, Sargan's specification test for the appropriateness of the instruments did not confirm this conjecture.

2/ Of course, since the estimated equation determining the instrumented value of $E[e(t+1)]$ does necessarily represent the "true" expectations process, simulation is meaningless.

liquidity are relevant for determining asset demands. 1/ The lack of an interest rate response would tend to reject the hypothesis of full uncovered interest rate parity since, in that case, the coefficient on the foreign rate would have to equal that on the expected exchange rate. This is not to say that the interest rate response of asset demands, and therefore policies, are not important in the determination of the exchange rate, since the reduced-form coefficients on the money stock are a function of the structural coefficients of the D matrix, which contain the demand response to domestic interest rate changes.

Given the similarity between the coefficients on both the current and lagged domestic and foreign currency liquidity stocks, the equations were reestimated to test the null hypothesis of equal coefficients. In the case of the pound/dollar rate, the null hypothesis could not be rejected, and the equation estimates under the accepted restriction are also reported in Table 1. However, with regard to the nominal effective exchange rate, only the restriction on the lagged real stocks was accepted. In both cases, reestimation subject to the restriction improved the diagnostic statistics substantially, especially regarding the autocorrelation of the errors.

As regards the implications of the parameter estimates for the possibility of exchange rate bubbles, in the simple first-order case with no lagged exchange rate terms, as discussed above, if the coefficient on the asset price expected for the next period is less than unity, the forward solution to the stochastic difference equation is required, in turn admitting the possibility of a bubble. Thus, estimation of the difference equation provides an indirect test for bubbles; if the coefficient on the expected future exchange rate is greater than unity, bubbles are precluded. Similarly, to apply this concept to equation (6), a system of second-order stochastic difference equations, would require examination of the characteristic roots of the lag process for the y(t) vector. Fortunately, the estimates of the exchange rate equation indicated that the exchange rate was recursive with respect to the (unestimated) price and interest rate equations; since the coefficients on the lagged real domestic and foreign currency stocks were indistinguishable except for sign, the lagged price term may be excluded. Therefore, noting that the coefficient on the lagged real foreign currency asset (in the restricted equation estimates) applies to the lagged exchange rate, a test of the bubble hypothesis is whether or not at least one root of the estimated difference equation is within the unit circle. The calculation revealed similar dynamics in both the (restricted) exchange rate equations, which were consistent with forward solutions to the exchange rate, and therefore with the existence of bubbles; in the case of the pound/dollar exchange rate, the roots were 0.40 and 1.07, while in the case of the nominal effective exchange rate

1/ In the case of the Treasury bill yield, this was likely owing to the only recent significance of the Treasury bill in investors' portfolios. The rejection of the significance of the foreign interest rate may have reflected difficulty with the proxy used.

the roots were -0.42 and -1.03. 1/ Thus the necessary condition for a rational expectations speculative bubble over the estimation period were satisfied.

The estimation results reported above cover only the period up to September 1987, just prior to the reversal in the trend of the exchange rate. It is clearly of interest to examine the performance of the estimates for the following months. While both equations predicted the reversal in trend, they both tended to overpredict its magnitude. Table 1 reports the standardized average forecast error, which when multiplied by the number of forecasts is distributed asymptotically chi-squared. Its size (well in excess of two) indicated the possibility of parameter instability between the two subperiods. However, reestimating all four equations over the longer period resulted in largely unchanged parameter estimates.

If the turnaround in the pound's exchange rate after September 1987 was due to the collapse of a speculative bubble, the model's parameters would not likely be stable if post-September 1987 data were included. Examining the LL/$ model between the 1982(4)-1987(9) and 1982(4)-1988(3) sample periods for evidence of a structural break yielded Chow test statistics (which measure the increase in the sum of squared residuals resulting from an increase in the sample period) of 3.95 and 3.53 for the unrestricted and restricted models, respectively, indicating modest evidence of a structural break. 2/ Similarly, the nominal exchange rate model produced evidence of a structural break in both its unrestricted and restricted form (Chow test statistics of 3.31 and 3.87, respectively). To identify the point of structural break, the recursive residuals for the four models were estimated for the 1982(4)-1988(3) period and were used to calculate the cumulative sum (W(t)) and cumulative sum of squares (S(t)) statistics. 3/ While in all of the four cases the W(t) statistic (not reported) remained well within the upper and lower 95 percent confidence bounds throughout the sample

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1/ This involves solving for ρ_1 and ρ_2 in the difference equation $(1-\rho_1 L)(1-\rho_2 L) = 1-bL-cL^2$ (Hanson and Sargent, 1980). Note that in each case one of the roots was close to unity suggesting that the model could have been expressed in first difference form.

2/ The Chow statistic is asymptotically distributed F(n,T-k), where n is the number of extra observations. Note, however, that the validity of this text in the context of an instrumental variables estimator, is not well-established.

3/ These tests were originally proposed by Brown, Durbin, and Evans and are discussed in Chow (1983), Chapter 10. If w(t) is defined as the standardized forecast error at t from regression estimates based on data to t-1, then the cusum (W(t)) and cusumsq (S(t)) test statistics are

$$W(t) = \sum_{j=k+1}^{t} w(j)/s, \text{ and } \quad S(t) = \sum_{j=k+1}^{t} w(j)^2 / \sum_{i=k+1}^{T} w(i)^2$$

for t>k+1 and where s is the standard error of the regression based on the whole data set.

period, some evidence existed (an inflection point) to suggest a break at 1987(9). This conjecture was more directly supported by examination of the cumulative sum of squares statistic. The statistics (also not reported) for the four exchange rate models broke the lower 95 percent confidence bound at mid-1987, indicating the possibility of structural break at that point and, therefore, of a bubble up to September 1987. However, the degree to which the bounds were breached was not excessive (in the unrestricted models the lower 99 percent bound was not breached and was only barely breached for the restricted models) and may have been due to outliers.

Thus, while the estimates discussed above did not directly indicate the existence of bubbles, the results admitted the possibility. The existence of significant autocorrelation and heteroscedasticity is symptomatic of a time-series process that is not adequately explained by the explanatory variables. Further, as discussed above, the coefficient estimates were consistent with the forward solution to the exchange rate and therefore with bubble solutions. The univariate and multivariate time-series properties of the exchange rate series will be examined below for further evidence of speculative bubbles.

2. Univariate analysis

As discussed above, the exchange rate's evolution over time will be dependent on the time-series properties of both the fundamental exchange rate and the bubble component (assuming it exists) so that a necessary condition for the existence of a bubble is for the exchange rate to exhibit the properties of a nonstationary stochastic process, which will not be relieved by differencing. 1/

In this regard, a well-known property of nonstationary time series is that their sample autocorrelations do not damp to zero as the lag length increases (for example, see Diba and Grossman (1988a) and Chow (1983)). As regards the log-level exchange rate series, the sample autocorrelations indicated severe nonstationarity; the correlations to the tenth lag were all above 0.97 and were significantly different from zero (the critical value at a 95 percent confidence level is approximately 0.24). This was the case regardless of the choice of sample or exchange rate definition, suggesting that the data did not reject the hypothesis of nonstationarity and the existence of a bubble. However, the autocorrelations of the log-differences of both the U.S. dollar and effective exchange rates were inconsistent with the

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1/ Note, however, that the existence of the type of nonstationary consistent with bubble phenomena does necessarily imply the existence of bubbles since it may in fact be that the fundamental determinants of the exchange rate are themselves nonstationary (this is exactly the criticism made by Hamilton (1986) and Hamilton and Whiteman (1986) of recent empirical tests for bubbles). Nonetheless, a useful preliminary test for the nonexistence of an exchange rate bubble is the examination of the time-series properties of the exchange rate.

existence of a bubble, damping considerably after the first lag (Table 2). 1/

Evidence provided by Meese (1986) suggests that this simple test may not distinguish between unit roots and the nonstationary alternative when the root is small. As an alternative to the above evidence, Tables 3 and 4 contain the results of Dickey-Fuller (1981) tests for unit roots in the level and first differenced exchange rate series, respectively, for the June 1982-April 1988 and June 1982-September 1987 sample periods. 2/ In all but one case the hypothesis of unit roots (nonstationarity) could not be rejected (ϕ_3 too small) for the level model of the Lebanese pound/U.S. dollar exchange rates. 3/ Similar results were found as regards the stationarity of the first differences; the hypothesis of unit roots (nonstationarity) for the first differenced U.S. dollar exchange rate could not be rejected in either sample period, whereas in the case of the first differenced nominal effective exchange rate, the null hypothesis was rejected for both sample periods. Thus, these tests would seem to reject the hypothesis of nonstationarity of the first differences of the nominal exchange rate series and, therefore, the existence of an exchange rate bubble, while not precluding a bubble in the Lebanese pound/U.S. dollar series. 4/

However, it is important to recall that the Dickey-Fuller test is not one-sided, i.e., the null of unit roots is tested against the alternate hypothesis of either stationary or nonstationary roots. Thus, rejection of the unit root hypothesis does not necessarily a reject the

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1/ Nonetheless, the Box-Pierce Q test, indicated that the differenced series retained a significant degree of autocorrelation.

2/ These consist of estimating the first difference of the variable in question x(t) on its own lagged first differences a time trend and its lagged value in levels and testing the null hypothesis of a unit root by testing the equivalent hypothesis that the coefficient on the lagged level is zero. As Evans and Savin (1984) demonstrate the more direct approach, a regression on the level of x(t), would yield biased results if indeed there were a unit root.

3/ The null hypothesis of unit roots is rejected over the shorter sample period for the nominal effective exchange rate. Note, however, that the coefficient estimates indicate that this discrepancy is largely a result of the significance of the time trend variable; in none of the four cases is the a_2 coefficient significantly different from zero.

4/ This, in turn, would seem to suggest that the source of the apparent nonstationarity was not speculation against the Lebanese pound but speculation in favor of the U.S. dollar.

Table 2. Sample Autocorrelations of the First Differences of Lebanese Pound/U.S. Dollar (LL/$) and Lebanese Pound Nominal Effective Exchange Rates (NEER)

Lag	1982(2) - 1988(4)		1982(2) - 1987(9)	
	ΔLL/$	ΔNEER	ΔLL/$	ΔNEER
1	0.537	0.563	0.521	0.587
2	0.266	0.316	0.214	0.311
3	-0.014	0.031	-0.056	0.044
4	-0.031	0.016	0.098	0.208
5	0.079	0.062	0.293	0.303
6	0.048	0.139	0.258	0.402
7	0.069	0.173	0.208	0.333
8	0.094	0.218	0.152	0.313
9	0.187	0.297	0.145	0.247
10	0.133	0.260	0.108	0.244
Q 1/	27.410	42.492	32.561	67.366

1/ The Box-Pierce statistic rejects the null hypothesis of stationarity for Q greater than 18.307 at the 95 percent confidence level. While this statistic's power is known to be poor in small samples, calculation of the modified statistic suggested by Box and Ljung did not change the results.

Table 3. Dickey-Fuller Test Results, 1982(6) - 1988(4) 1/

Coefficient	LL/US$ Rate Δe_t	LL/US$ Rate $\Delta^2 e_t$	Nominal Effective Rate Δe_t	Nominal Effective Rate $\Delta^2 e_t$
a_o	-0.010	-0.021	0.223	0.049
	(0.025)	(0.027)	(0.114)	(0.029)
a_1	0.004	0.002	-0.004	-0.003
	(0.002)	(0.001)	(0.001)	(0.001)
a_2	-0.037	-0.832	-0.031	-1.125
	(0.022)	(0.249)	(0.018)	(0.274)
b_1	0.463	0.295	0.436	0.546
	(0.123)	(0.219)	(0.123)	(0.236)
b_2	0.068	0.355	0.104	0.597
	(0.131)	(0.190)	(0.128)	(0.205)
b_3	-0.288	0.041	-0.349	0.211
	(0.138)	(0.171)	(0.137)	(0.183)
b_4	-0.011	-0.019	-0.052	0.175
	(0.137)	(0.145)	(0.139)	(0.155)
SSR	0.513	0.533	0.495	0.507
Φ_3	4.556	5.763	6.236	8.741

1/ Standard errors are in parentheses. Sample size of second difference equations are 1982(7) - 1988(4). Regressions are of the form $\Delta x(t) = a_o+a_1t+a_2x(t)+\Sigma b_i\Delta x(t-i)$. The test statistic Φ_3 is calculated as an F statistic for the null hypothesis, distributed such that the critical value at the 95 percent confidence level is between 6.73 and 6.49 for sample sizes between 50 and 100.

Table 4. Dickey-Fuller Test Results, 1982(6) - 1987(9) 1/

Coefficient	LL/US$ Rate Δe_t	LL/US$ Rate $\Delta^2 e_t$	Nominal Effective Rate Δn_t	Nominal Effective Rate $\Delta^2 n_t$
a_o	-0.030 (0.027)	-0.038 (0.025)	0.106 (0.110)	0.061 (0.025)
a_1	0.003 (0.002)	0.002 (0.001)	-0.004 (0.001)	-0.003 (0.001)
a_2	-0.013 (0.026)	-0.869 (0.256)	-0.009 (0.019)	-1.088 (0.259)
b_1	0.399 (0.134)	0.241 (0.224)	0.390 (0.132)	0.480 (0.223)
b_2	-0.044 (0.136)	0.223 (0.181)	-0.031 (0.130)	0.403 (0.177)
b_3	-0.390 (0.136)	-0.167 (0.161)	-0.472 (0.130)	-0.072 (0.161)
b_4	0.120 (0.137)	-0.103 (0.137)	0.141 (0.135)	0.097 (0.137)
SSR	0.347	0.342	0.266	0.265
Φ_3	5.672	5.914	7.849	8.871

1/ Standard errors are in parentheses. Sample size of second difference equations are 1982(7) - 1987(9). Regressions are of the form $\Delta x(t) = a_o+a_1t+a_2x(t)+\Sigma b_i\Delta x(t-i)$. The test statistic Φ_3 is calculated as an F statistic for the null hypothesis, distributed such that the critical value at the 95 percent confidence level is between 6.73 and 6.49 for sample sizes between 50 and 100.

hypothesis of nonstationarity. 1/ Bhargava (1986) has developed Von Neuman type statistics that allow for testing the unit root hypothesis against one sided alternatives, e.g., against the alternative of an explosive root. The results of applying these tests to the levels and first differences of the exchange rate series are presented in Table 5. The first statistic (R_1) tests the null hypothesis of a simple random walk (unit root) against the stationary alternative representation of the autoregressive process. As the figures indicate, the null hypothesis was clearly accepted in the case of the log-level series but was rejected at the 95 percent confidence level for the first differenced U.S. dollar and nominal effective exchange rate series, regardless of the sample period chosen. The N_1 statistic tests the null hypothesis of the unit root against the alternative of an explosive root (for small values of the statistic) or a stable root (for large values of the statistic). Again the level series were apparently explosive, while the differenced series were stationary, regardless of the sample period or definition of the exchange rate. Finally, the N_2 statistic allows for a deterministic trend in the autoregressive process and tests the null hypothesis of a unit root against the alternative of explosive roots (for small values of the statistic) or the existence of stable roots (for large values of the statistic). Again, the null hypothesis (of unit roots) was rejected in favor of explosive roots in the case of the level exchange rate series and stationarity in the case of the first differences.

Thus, the time-series properties of the exchange rate series argue against the existence of a speculative bubble over the sample period in question. While some indication of bubble behavior was evident for the Lebanese pound/U.S. dollar rate in the context of the Dickey-Fuller tests, it was not apparent upon calculation of the one-sided tests proposed by Bhargava. Nonetheless, the apparent significance of the time-trend variable in the Dickey-Fuller regressions for both the level and differenced series suggests the possibility that the tests above may be unable to distinguish between a bubble phenomenon and the deterministic time trend. It is to this possibility that the next section is addressed.

3. Cointegration tests

It was suggested above that, if the exchange rate was subject to a bubble process, stationarity could not be achieved regardless of the order of differencing. However, this "test" does not consider the

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1/ Nonetheless, simulating the autoregressive process for the first difference of the nominal effective exchange rate (implied from the coefficient estimates for regressions explaining the second differences) indicates stationarity in first differences. Further, the significance of the time trend in each of the regressions indicates this variable may be proxying the deterministic portion of the solution to the difference equation. There is also some doubt (Evans and Savin (1984)) regarding the power of the Dickey-Fuller test, especially in small samples.

Table 5. Bhargava Tests for Stationarity 1/

Test Statistic	Sample Size 2/	LL/US$ Rate e_t	LL/US$ Rate Δe_t	Nominal Effective Rate e_t	Nominal Effective Rate Δe_t
R_1 3/	1982(6) - 1988(4)	0.007	0.931	0.006	0.888
	1982(6) - 1987(9)	0.010	0.957	0.009	0.834
N_1 4/	1982(6) - 1988(4)	0.004	0.707	0.004	0.743
	1982(6) - 1987(9)	0.005	0.682	0.005	0.577
N_2 5/	1982(6) - 1988(4)	0.014	0.573	0.011	0.714
	1982(6) - 1987(9)	0.009	0.496	0.007	0.629

1/ Asterisks indicate acceptance of the null at 95 percent confidence level.

2/ The sample in the case of the first difference models begins in 1982(7).

3/ The R_1 test statistic tests the null hypothesis of a simple random walk versus a stationary alternative (for large values of R_1, approximately greater than 0.37).

4/ The N_1 test statistic tests the null hypothesis of a unit root versus either an explosive model (for small values of N_1, approximately less than 0.009) or a stable model (for large values of N_1, approximately greater than 0.24).

5/ The N_2 statistic tests the null hypothesis of a unit root versus either a model with roots greater than unity and a time trend (for small values of N_2, approximately less than 0.031) or a stable model and a time trend (for large values of N_2, approximately greater than 0.37).

possibility that some other fundamental (or Granger exogenous) variable could be the source of any nonstationarity. To address this possibility, the concept of cointegration may be applied; the components of a vector X(t) are said to be cointegrated of order (d,b) if all components of X(t) are integrated of order d and there exists a (nontrivial) vector (a) such that z(t) = a'X(t) is integrated of order (d-b) (Granger and Engle (1987).

This definition's relevance to bubble behavior is twofold. First, by definition, if the asset price is cointegrated with other variables, bubble behavior is precluded since a requirement is that the components must be stationary after differencing. Second, suppose that the exchange rate is related to some fundamental variable (or variables) Z(t) such that

$$e_t = \lambda Z(t) + u(t) \tag{8}$$

where u(t) is a residual and, in the context of the discussion above, $\lambda Z(t)$ is the fundamental exchange rate solution f(t). If e(t) and Z(t) are cointegrated of order (d,b) where the cointegrating vector is $(1,-\lambda)$ then the residual u(t) cannot contain a bubble component since it is integrated of order (d-b). 1/ Granger and Engle develop a number of simple statistics to test for cointegration. Assuming that e(t) and Z(t) are integrated of order 1, the first test requires estimation of the cointegrating regression relating the exchange rate to the current fundamental variables, and testing the hypothesis that the Durbin-Watson statistic (DW) is zero, so that large values of DW reject the null hypothesis of cointegration. A second test requires subjecting the residuals to augmented Dickey-Fuller tests for unit roots. 2/

The estimations described in the previous section provide strong evidence that the exchange series is related to the stock of domestic currency-denominated liquidity relative to foreign currency-denominated liquidity. Evidence was provided in the previous section that the nominal effective and (possibly) the LL/$ exchange rates were stationary in first differences. Dickey-Fuller tests indicated unit roots in the

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1/ It should be clear, however, that despite the application of this concept to empirical tests for bubble phenomenon by the aforementioned authors, its particular relevance is not well-established. An important preliminary to the test for cointegration is to establish that the vector time series is integrated (jointly) of order d. This in itself precludes bubbles.

2/ Granger and Engle report critical values for these statistics only for the two variable, cointegration of order (1,0) case. Therefore, the tests are limited to the simple cases described above. While Meese (1986) performs tests based on greater than two variables, the distributions of the test statistics were assumed to be unchanged from the two variable case. Engle and Yoo (1987) relieve some of these concerns by admitting greater than two variables; nonetheless, the systems they examine are limited to the (1,0) case.

log-level domestic currency money series for the 1982(6)-1988(4) and 1982(6)-1987(9) sample periods (ϕ_3 = 1.697 and ϕ_3 = 1.363, respectively) and were suggestive of no unit roots for the first differences (ϕ_3 = 4.949 and ϕ_3 = 8.618, respectively). As regards the time-series properties of the log differences of the real money stocks, the results of the Dickey-Fuller tests were inconclusive; tests of a unit root in levels did not reject the hypothesis (of unit roots) for the 1982(6)-1988(4) period but did reject the hypothesis for the shorter 1982(6)-1987(9) period (ϕ_3 = 6.967 and ϕ_3 = 7.684, respectively).

Table 6 reports the results of tests for cointegration between the Lebanese domestic currency component of the money supply and the two exchange rate indices, respectively, and the results of similar tests of cointegration between the exchange rate indices and the log difference of the domestic and foreign currency component of the money stock. The test statistics uniformly rejected the hypothesis of cointegration between the exchange rate series and the money supply. Given that the exchange rate and money supply series are jointly integrated of order 1, the residuals of the cointegrating equation must be integrated of order zero, i.e., stationary, if e and m are cointegrated. However, the Durbin-Watson statistics were too small to accept this hypothesis and, similarly, estimating Dickey-Fuller equations for the residuals confirms the nonstationarity of the residuals. The statistics also reject the hypothesis of cointegration between the exchange rate indices and the log differences of the real money stocks. As in the previous case, the Durbin-Watson statistics in the cointegrating equations were too small while the evidence of unit roots in the residual series was too strong. Thus, tests of cointegration between the exchange rate and the fundamentals did not rule out exchange rate bubbles. However, the specific nature of Granger and Engle's test, i.e., two variables cointegrated of order (1,0), limits its usefulness in applications of this sort. The great likelihood, as evidenced by the reduced-form equations estimated above, is that there exists a much richer array of variables and time-series relationships that relate the exchange rate to the fundamentals. Thus, the failure to accept the hypothesis of cointegration may simply be due to the exclusion of relevant variables.

V. Conclusion

Conventional exchange rate theory suggests that exchange rates should be determined with reference to fundamental economic variables or to their expectation. However, as discussed above, theory does not preclude the possibility that exchange rates deviate, at least temporarily, from their so-called fundamental value owing to self-fulfilling speculation solely with regard to the future course of the exchange rate. Such deviations have been termed speculative bubbles. The exchange rate for the Lebanese pound would appear to have exhibited some of the symptoms of an exchange rate bubble; the rate of its depreciation accelerated markedly over 1986 and 1987, well in excess of

Table 6. Cointegration Test Results 1/

Sample size	Pound/U.S. dollar Rate Σ_1	Pound/U.S. dollar Rate Σ_3	Nominal Effective Rate Σ_1	Nominal Effective Rate Σ_3
$e(t) = c + \alpha m(t) + u(t)$				
1982(1)-1988(4)	0.141	-2.186	0.113	2.070
1982(1)-1987(9)	0.077	1.450	0.054	1.242
$e(t) = c + \alpha[m(t)-m^*(t)+e(t)] + u(t)$				
1982(1)-1988(4)	0.108	-2.362	0.152	0.866
1982(1)-1987(9)	0.082	2.785	0.121	2.209

1/ Σ_1 is the DW statistic for the cointegrating equation. Large values (greater than 0.386) reject the hypothesis of non-cointegration at the 95 percent confidence level. Σ_2 is the t-statistic on the coefficient of the lagged value of u(t) in the Dickey-Fuller equation for u(t). Large values (greater than 3.17) reject the hypothesis of non-cointegration. Note that the critical values are approximate since Engle and Granger report statistics for only the 100 observation case.

$e(t)$: The log of the monthly average Lebanese pound/U.S. dollar or the nominal effective exchange rate. Sources are International Monetary Fund, IFS and staff estimates, respectively.

$m(t)$: The log of end-of-period domestic currency liquidity (money plus quasi-money less resident foreign currency deposits). Sources are International Monetary Fund, IFS; Banque du Liban, Bulletin Trimestriel; and staff estimates (for missing observations).

$m^*(t)$: The log of end-of-period residents' foreign currency deposits converted to U.S. dollars or the composite currency using $e(t)$. Sources are Banque du Liban, Bulletin Trimestriel and staff estimates (for missing observations).

$tb(t)$: The log of end-of-period nonbank private sectors' holdings of public sector debt. Source is Banque du Liban, Bulletin Trimestriel.

$r^*(t)$: The monthly average of three-month US$ or SDR LIBOR. Source is International Monetary Fund, IFS.

$r^{tb}(t)$: The average interest rate on public sector debt with less than one year to maturity. After 1985, secondary market yields on banks with one year to maturity were used. Sources are Banque du Liban, Bulletin Trimestriel and staff estimates.

$p(t)$: The log of the monthly average consumer price index. Sources are the Beirut Chamber of Commerce and staff estimates for (missing observations).

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