Federal Reserve Bank of St. Louis

Review

March 1982
Vol. 64, No. 3

The Review *is published 10 times per year by the Research Department of the Federal Reserve Bank of St. Louis. Single-copy subscriptions are available to the public free of charge. Mail requests for subscriptions, back issues, or address changes to: Research Department, Federal Reserve Bank of St. Louis, P.O. Box 442, St. Louis, Missouri 63166.*

Monetary Policy and Stock Returns: Are Stock Markets Efficient?

LAWRENCE S. DAVIDSON and RICHARD T. FROYEN

AN efficient market is one that quickly processes all relevant information. For example, if monetary policy affects stock returns, then an efficient stock market rapidly digests and incorporates all news about monetary policy. Consequently, past policy actions will have little value or explanatory power in understanding current stock returns. Previous tests of stock market efficiency have examined the relationship between the timing of the growth of money and stock returns. Although several early studies found that stock returns lagged behind money growth — evidence of stock market inefficiency — the results of recent studies have supported the efficient market hypothesis.[1]

The purpose of this article is to provide further evidence on the timing of the relationship between monetary policy changes and stock returns by estimating models that express stock returns as functions of anticipated and unanticipated monetary policy measures. These models extend previous work in several directions. First, past studies generally have divided money growth into anticipated and unanticipated components in a mechanical or ad hoc fashion.[2] We compare these results with estimates of anticipated money growth measured by the fitted values of previously estimated monetary policy reaction functions. This enables us to determine whether the efficient market findings are robust across differing aggregates and decompositions of monetary policy into anticipated and unanticipated components.

Second, previous studies focused on the relationship between money growth rates and stock returns. But, during much of the period covered by these studies, the Federal Reserve's short-run (month-to-month) operating target was the federal funds rate. Therefore, in addition to estimating relationships between stock returns and money growth rates, we estimate models relating stock returns and both anticipated and unanticipated monetary policy actions using the federal funds rate. Again, anticipated and unanticipated policy actions will be de-

Lawrence S. Davidson, an associate professor of business economics and public policy at Indiana University, is a visiting scholar at the Federal Reserve Bank of St. Louis. Richard T. Froyen is an associate professor of economics at the University of North Carolina.

[1]Examples of studies that indicated a lag in the adjustment of stock returns to changes in money growth rates are: Michael J. Hamburger and Levis A. Kochin, "Money and Stock Prices: The Channels of Influence," *Journal of Finance* (December 1971), pp. 1045-66; Michael W. Keran, "Expectations, Money, and the Stock Market," this *Review* (January 1971), pp. 16-31; and Beryl W. Sprinkel, *Money and Stock Prices* (Richard D. Irwin, Inc., 1964). Recent studies that support the market efficiency postulate include: Michael S. Rozeff, "Money and Stock Prices: Market Efficiency and the Lag in Effect of Monetary Policy," *Journal of Financial Economics* (September 1974), pp. 245-302; John Kraft and Arthur Kraft, "Determinants of Common Stock Prices: A Time Series Analysis," *Journal of Finance* (May 1977), pp. 417-25, and "Common Stock Prices: Some Observations," *Southern Economic Journal* (January 1977), pp. 1365-67; R. V. L. Cooper, "Efficient Capital Markets and the Quantity Theory of Money," *Journal of Finance* (June 1974), pp. 887-908; Richard J. Rogalski and Joseph D. Vinso, "Stock Returns, Money Supply and the Direction of Causality," *Journal of Finance* (September 1977), pp. 1017-30; James B. Kehr and David Leonard, "Monetary Aggregates, the Stock Market and the Direction of Causality," *Journal of the Midwest Finance Association* (1980), pp. 47-57; and J. Ernest Tanner and John M. Trapani, "Can the Quantity Theory be Used to Predict Stock Prices — Or Is the Stock Market Efficient?" *Southern Economic Journal* (October 1977), pp. 261-70.

[2]Rozeff, "Money and Stock Prices," for example, assumes that anticipated money growth in a given month depends on money growth in the past three months.

rived from an empirical reaction function in which the federal funds rate is the dependent variable.

Third, we extend the time period in earlier studies through 1977. This allows us to examine the monetary policy/stock return relationship in both a period of low stable inflation (1954-65) and one of higher and more variable inflation and money growth (1966-77).

Finally, for the period from 1974 through 1976, we estimate models that relate *weekly* stock returns to the anticipated and unanticipated components of *weekly* money growth. Most previous work on this topic used quarterly or monthly data.[3] Estimates with weekly data provide a finer test of the efficient market hypothesis.

DO STOCK RETURNS LAG OR LEAD MONETARY POLICY?

Several recent studies of the relationship between money growth rates and stock returns have found that *future* money growth rates affect current stock returns. Thus, stock returns appear to lead money growth rates.[4] Other studies, however, do not find such effects.[5]

The finding that stock prices lead money growth has been interpreted in several different ways. One interpretation is that stock prices are a causal influence on money growth. However, as Rozeff points out, within the general equilibrium setting of financial markets, it is arbitrary to single out stock returns as a causal variable.[6] Rather, the evidence that future money growth rates affect current returns may be a reflection of the influence of other variables on both stock prices and money growth, with stock prices adjusting more quickly and, therefore, leading money growth rates.

Another interesting interpretation of this finding is provided by the "reversed causation with accurate anticipations" model.[7] In this model, causation runs from currently anticipated money growth to stock returns. The apparent effect of future money growth reflects the accurate anticipations of future money growth by the market. It is these accurate predictions of future money growth that affect current stock returns.

SPECIFICATION OF THE MODELS

This section describes two simple models of equity return determination. Tobin's theoretical model of the financial sector stressed the importance of the return on capital as the link between the real and financial sectors.[8] His model established a potential causal connection between the exogenous variables of the commodities and financial markets and the return on equities (ownership claims on the capital stock). The first of the two models presented here is a simple version of Tobin's, originally estimated by Rozeff.[9] This model stressed the linkage between monetary aggregates and the equity return. It imposed the additional restriction that only unanticipated changes in the growth rate of money (g^u) cause unanticipated movements in the equity return (R^u).

Rozeff's "predictive monetary portfolio" model relates the *unanticipated* current return on equities (R_t^u) to past unanticipated changes in monetary growth rates, that is,

$$(1)\quad R_t^u = f(g_{t-1}^u,\ldots,g_{t-n}^u) + \epsilon_t,$$

where R_t^u is the unanticipated movement in the equity return, defined as the actual return (R_t) minus the expected return conditioned on all available past information ($E[R_t/B_{t-1}]$). Unanticipated money growth in period t-i, g_{t-i}^u, is measured as the change in the money growth rate between t-i and t-i-1. The error term, ϵ_t, is assumed to be a normally distributed random variable with a mean of zero and a constant, finite variance. Rozeff assumed that the expected value of the nominal equity return is constant ($E[R_t/B_{t-1}]=C_0$) and the monthly empirical counterpart of the *predictive* model is:

$$(2)\quad R_t = C_0 + \sum_{i=1}^{16} a_i\, g_{t-i}^u + \epsilon_{2t},$$

where C_0 and a_i are parameters to be estimated.

[3]One recent exception is Neil G. Berkman, "On the Significance of Weekly Changes in M1," *New England Economic Review* (May/June 1978), pp. 5-22.

[4]See, for example, Rozeff, "Money and Stock Prices;" Kraft and Kraft, "Determinants of Common Stock Prices;" and Rogalski and Vinso, "Stock Returns, Money Supply and the Direction of Causality."

[5]See, for example, Kehr and Leonard, "Monetary Aggregates, the Stock Market and the Direction of Causality."

[6]See Rozeff, "Money and Stock Prices."

[7]See Rozeff, "Money and Stock Prices," pp. 275-76.

[8]James Tobin, "A General Equilibrium Approach to Monetary Theory," *Journal of Money, Credit, and Banking* (February 1969), pp. 15-29.

[9]See Rozeff, "Money and Stock Prices," pp. 255-66.

To evaluate the relative importance of the most recent monetary information, Rozeff also estimated the *nonpredictive* monetary portfolio model. In this model, the contemporaneous money surprise is added; the lag on the monetary surprises starts at zero instead of one:

$$(3)\quad R_t = C_0 + \sum_{i=o}^{16} a_i\, g^u_{t-i} + \epsilon_{3t}.$$

A final variant of this model assumes that market participants form expectations of future changes in monetary growth. If these expectations are at least unbiased, then future monetary growth rates would cause changes in current equity returns. Rozeff's empirical nonpredictive monetary portfolio model with anticipations adds eight leads (negative lags) to equation 3:[10]

$$(4)\quad R_t = C_0 + \sum_{i=-8}^{16} a_i\, g^u_{t-i} + \epsilon_{4t}.$$

To test whether past information about unexpected monetary growth influences current stock returns, we examine the statistical significance of the lagged unanticipated money growth terms in the *predictive* model (equation 2). If the stock market is efficient, the coefficients on the lagged terms should be equal to zero ($a_i=0$, $i=1,\ldots,n$). An F-test is used to test this hypothesis; an F-value significantly greater than 1.0 would suggest that the stock market was inefficient, since *past* information would affect current stock returns.

On the other hand, a significant F-value for a similar test of the coefficients in the nonpredictive models (equations 3 or 4) does *not* indicate market inefficiency. The finding that only current monetary growth affects returns simply establishes the importance of monetary variables in equity return determination. If future, but not past, money growth affects current returns, this suggests a forward-looking propensity of the market which also is not inconsistent with an efficient market.

The second model of equity returns considered here is referred to as the Fama approach.[11] In this model, the nominal return on stocks (R_t) is assumed to be composed of the real return (r_t) and a premium for expected inflation (π_t) — a Fisher effect for stock returns:

$$(5)\quad R_t = r_t + \pi_t.$$

From equation 5, we can write the expected value of the *nominal* return conditioned on information available from period t-1 (B_{t-1}), as

$$(6)\quad E(R_t/B_{t-1}) = E(r_t/B_{t-1}) + E(\pi_t/B_{t-1}).$$

If we assume a constant real mean of stock returns (c_o), we can rewrite equation 6 as

$$(7)\quad E(R_t/B_{t-1}) = c_0 + E(\pi_t/B_{t-1}).$$

Since $E(R_t/B_{t-1})$ is equal to the actual nominal return on stocks (R_t) minus its unanticipated component (R^u_t), we can transform equation 7 into an expression for the actual nominal stock return:

$$(8)\quad R_t = c_0 + R^u_t + E(\pi_t/B_{t-1}).$$

Equation 8 then can be converted into a relationship between money growth and nominal stock returns if we express (as in equation 1) the unanticipated component of stock returns as a function of unanticipated changes in money growth and if, further, we express the expected inflation rate as a function of expected money growth. With these assumptions, our expression for nominal stock returns becomes

$$(9)\quad R_t = c_0 + f(g^u_t,\ g^u_{t-1},\ldots,g^u_{t-n_1}) + h(g^*_t,\ g^*_{t-1},\ldots,g^*_{t-m_2}) + v_t,$$

where g^*_t is the expected rate of growth of the money stock, and h is the function relating expected money growth to expected inflation. The empirical counterpart to equation 9 used in our estimation is

$$(10)\quad R_t = c_0 + \sum_{i=0}^{n_1} b_i\, g^u_{t-i} + \sum_{j=0}^{n_2} d_j\, g^*_{t-j} + v_t,$$

where various lag lengths and several different measures of anticipated and unanticipated money growth are employed.

Additionally, one test uses the federal funds rate rather than a monetary aggregate as the monetary policy variable. The effects of this substitution on the theoretical interpretation of our models of equity return are discussed below.

Using the Fama (or Fisher) model of stock returns, we can also test for market efficiency. Market efficiency implies that lagged *unanticipated* changes in

[10]Future values of *unanticipated* money growth should not cause current stock market returns to change. However, the exact interpretation of g^u_{t+i} is not unambiguous. It could be reinterpreted as the perfectly correct anticipated future change in money growth. In that case, it would be an indicator of the forward-looking propensity of the market.

[11]This approach is set out in Eugene F. Fama, "Short-Term Interest Rates as Predictors of Inflation," *American Economic Review* (June 1975), pp. 269-82.

money growth rates would not affect current stock returns ($b_i = 0$ for $i > 0$ in equation 10). In the Fama approach, however, lagged *anticipated* changes in money growth rates might affect current stock returns through an effect on expected future inflation. This result would not violate market efficiency; it would simply be an element of $E(R_t/B_{t-1})$ and would not provide a basis for any profitable trading rules.[12] This effect of anticipated monetary policy on stock returns is another channel by which monetary policy may affect stock prices — even in an efficient market — an effect we test for in the following section.

ESTIMATES OF THE MODELS

Five sets of model estimates are presented. In all five, the measure of the nominal equity return is the percentage change (measured from the last business day in each month or week) in the overall index of all stock prices on the New York Stock Exchange.[13] These tests employ a variety of monetary policy measures.[14] These include: 1) percentage changes in actual, anticipated and unanticipated M1 and the monetary base, and 2) anticipated and unanticipated values of the federal funds rate.

The policy measures in all the tests, except those with weekly data, are changes in average monthly values. Returns are changes between the last business day of each month. This specification relates the cumulative stock price change from the end of one month to the next to the average month-to-month change in the monetary policy variable. As a result, the stock return variable is more sensitive than the policy variables to last-day-of-the-month activity. Changes in the average monthly value would appear to be the proper measure of the shift in monetary policy from month to month. We relate this to the cumulative change in stock prices for the month. This does mean, however, that while the dependent and independent variables pertain to the same time period, they weight *daily* observations within the time period differently. Our tests with weekly data therefore provide more intra-month precision.

Unanticipated Money Growth and Stock Returns: Alternative Specifications of the Basic Models

The models in equations 2-4 specify that unanticipated money growth affects the unanticipated stock return. Rozeff's tests make the following two explicit assumptions:

i) $R_t^u = R_t - C_o$, and

ii) $g_t^u = g_t - g_{t-1}$.

The unanticipated return is a deviation from a mean ($R_t - C_0$), while the unanticipated money growth rate is a first difference ($g_t - g_{t-1}$). This section compares the results based on these assumptions with two alternative specifications. The first of these we call the differenced model:

iii) $R_t^u = R_t - R_{t-1}$,

iv) $g_t^u = g_t - g_{t-1}$.

The second is called the mean deviation model:

[12]See Rozeff, "Money and Stock Prices," p. 260.

[13]An alternative measure includes dividends, but because its variance is so dominated by stock price changes, it performs almost identically to the index which contains only prices. This alternative measure is not used in our tests.

[14]These measures of monetary policy each have limitations for the testing of the efficient market hypothesis. Tests of this hypothesis must distinguish between information which is currently *known and used* by market participants and that which is not. In fact, we do not know what information was available to and used by these agents. In this research, we have limited the monetary policy measures to those listed above. We have not tried narrower or broader measures of money like nonborrowed reserves or M2, nor have we used seasonally unadjusted versions of M1 or the monetary base. Our tests have selectively employed both revised and initially announced seasonally adjusted versions of M1. Since seasonally adjusted data are revised several times, it would seem preferable to use the initially announced numbers since those were the ones available to market participants. Furthermore, Courtenay C. Stone and Jeffrey B. C. Olson, "Are the Preliminary Week-to-Week Fluctuations in M1 Biased?" this *Review* (December 1978), pp. 13-20, have shown with weekly data that the revised seasonally adjusted series is largely independent of the unrevised series and therefore is a poor proxy for that data. Our weekly aggregate tests, therefore, employ the unrevised growth rates of seasonally adjusted M1.

This use of initially announced data is not without drawbacks. For example, since initial announcements have been shown to be unreliable indicators of how money is performing, market participants may either ignore seasonally adjusted data or they may modify it. One useful modification would discount the announcement with what agents think is the true seasonal adjustment. If they do this correctly, then they are using what turns out to be the actual revisions. If they use seasonal adjustment factors that are different from the true ones, they are using an unobservable series. Our monthly aggregate tests use the revised, seasonally adjusted growth rates of M1.

The monetary reaction function tests do not rely totally upon either revised or unrevised data. For example, the consumer price index and the unemployment rate, which are used to predict the monetary base, are not regularly revised. However, the monetary base itself, like M1, is revised frequently. Finally, the tests with the federal funds rate have no data revision problems since this series is not revised.

Table 1

Summary Statistics for Lead-Lag Money Growth (g) and Equity Return (R) Models[1]

		1954 — 1965					1966 — 1977				
					Number of significant coefficients					Number of significant coefficients	
Model	Lag (lead) specification	F	$\overline{R}^2$	DW	Lags	Leads	F	$\overline{R}^2$	DW	Lags	Leads
Mixed											
2	16 to 1	1.382	.149	1.78	2	—	.889	.100	1.80	0	—
3	16 to 0	1.296	.150	1.79	1	0	.859	.103	1.83	0	0
4	16 to (9)	1.488	.250	1.92	1	2	2.790*	.381	1.87	3	5
Differenced											
2	16 to 1	1.849*	.178	2.83	0	—	1.480	.147	2.90	0	—
3	16 to 0	1.888*	.192	2.83	0	0	1.720	.177	2.88	0	1
4	16 to (9)	1.285	.214	2.85	0	0	2.990**	.386	2.84	0	2
Mean Deviation											
2	16 to 1	1.748*	.170	1.80	4	—	1.150	.119	1.85	0	—
3	16 to 0	1.649	.172	1.82	4	0	1.070	.119	1.86	0	0
4	16 to (9)	1.720*	.267	1.95	1	1	2.940**	.384	1.91	2	3

Note: In all cases the dependent variables are some transform of the equity return, R. $\overline{R}^2$ is the adjusted coefficient of determination. F is the F-value, and DW is the Durbin-Watson statistic. A * (**) implies rejection of the null hypothesis at the 95% (99%) level. The null hypothesis states that the estimated coefficents of the independent variables equal zero. The Leads columns include the contemporaneous terms.

[1]Data are monthly observations.

v) $R_t^u = R_t - C_0$,

vi) $g_t^u = g_t - g_0$,

where C_0 and g_0 are the sample-period means of R and g, respectively.

Since the original Rozeff specification mixes deviations from means ($R_t - C_0$) with first differences ($g_t - g_{t-1}$), we refer to this as the mixed model. None of the three versions inherently makes more sense than the others. Our intent here is to see how sensitive the original specification is to these minor changes.

Table 1 provides estimates of the original empirical specifications of the three models: the *mixed model,* given by equations 2, 3 and 4, and the modified specifications which we term the *differenced model* and the *mean deviation model.* The estimates in the table cover two subperiods, 1954-65 and 1966-77.

The results in table 1 offer no clear rejection of Rozeffs specification. All three models explain more of the variance of equity returns when current or future money growth is included in the regressions. In the 1966-77 time period, individual coefficients on past monetary information are never significant, nor are they ever significant as a group. In this period, the effect of future money is highly significant, tripling the explanatory power of the estimated models.

In the earlier period, there are no unambiguous differences among the models. The $\overline{R}^2$ reveals relatively equal explanatory power. The differenced model shows a statistically significant effect of the 16 lags of money growth, yet no single coefficient is statistically significant. This model exhibits a high degree of autocorrelation; therefore, the F-tests should be interpreted with caution.[15] The mean deviation model also shows an apparent significant effect of past money growth in the early period. However, when future terms are added to the equation, the number of lagged significant coefficients falls to only one. As a whole, these results offer no clear rejection of stock market efficiency. The effects of future money growth on stock returns are also robust with respect to the type of specification changes we have made.

[15]As is well known, autocorrelation leads to a bias in the standard error of the regression. With negative autocorrelation, the direction of the bias could be positive or negative.

Table 2

Reaction Function Estimates of Unanticipated Monetary Policy ($\tilde{g}_1^u$, $\tilde{g}_2^u$) and Equity Returns (1954:7 to 1972:3)[1]

Independent variable	Model	Lag (lead) specification	F	$\bar{R}^2$	DW	Number of significant coefficients: Lags	Number of significant coefficients: Leads
$\tilde{g}_1^u$	2	16 to 1	.655	.051	1.78	0	—
	3	16 to 0	.621	.051	1.78	0	0
	4	16 to (9)	.946	.117	1.87	7	7
$\tilde{g}_2^u$	2	16 to 1	.896	.068	1.79	0	—
	3	16 to 0	.969	.078	1.83	0	0
	4	16 to (9)	2.660**	.271	1.95	0	3

[1]See note, table 1. Data are monthly observations.

Unanticipated Monetary Base Growth and Stock Returns: Estimates Employing Monetary Reaction Functions

The basic mixed model is retained in this section, but two different proxies for unanticipated monetary policy actions (g^u) are tried. In these tests, we assume that agents are rational and act as if they know the appropriate function guiding monetary policy.

Table 2 presents the results of estimating equations 2, 3 and 4 using two different proxies for unanticipated money growth. The first of these, denoted $\tilde{g}_1^u$, comes from Froyen's monetary policy reaction function for the monetary base.[16] This function, which we assume is used to forecast future growth rates of the monetary base, relates the latter to past values of the Federal Reserve's assumed goal variables: the unemployment rate, inflation rate, balance of payments and the outstanding government debt held by the public. The estimated function is used to predict the level of the monetary base.

If M_t^* is the prediction of the monetary base based on the estimated reaction function, then we can define the anticipated monetary base growth rate as[17]

$$\tilde{g}_{1,t}^* = (M_t^* - M_{t-1}^*) / M_{t-1}^*.$$

Therefore, a first proxy for unanticipated monetary base growth is

$$\tilde{g}_{1,t}^u = \tilde{g}_{1,t}^* - g_t.$$

The second proxy for unanticipated growth ($\tilde{g}_{2,t}^u$) is based on a simple third-order autoregressive process similar to the specification used by Rozeff:

$$\tilde{g}_{2,t}^u = \tilde{g}_{2,t}^* - g_t,$$

where

$$\tilde{g}_{2,t}^* = \hat{\alpha}_0 + \hat{\alpha}_1 g_{t-1} + \hat{\alpha}_2 g_{t-2} + \hat{\alpha}_3 g_{t-3}.$$

The results in table 2 again support the efficient market hypothesis. There is no clear evidence that past unanticipated monetary base growth significantly affects current stock returns using any of the proxies tested here. While there are numerous significant lag coefficients in the $\tilde{g}_1^u$ equation, they are not significant until leads are added, and even then the F-value is not significant. With regard to the effects of future monetary base growth on current stock returns, the pattern of the results in table 2 is interesting. When anticipated monetary base growth is measured by the simple autoregressive specification, and future "unanticipated" monetary base growth is taken to be money growth that cannot be predicted with that specification, $\tilde{g}_{2,t}^u$, our results show a significant effect for these future terms. However, for the proxy constructed on the basis of the estimated monetary policy reaction function, $\tilde{g}_{1,t}^u$, future unanticipated monetary base growth has no significant effect on current stock returns.

[16]See Richard T. Froyen, "A Test of the Endogeneity of Monetary Policy," *Journal of Econometrics* (July 1974), pp. 175-88.

[17]Alternatively, we tried a variant of this form where

$$g_{3,t}^* = (M_t^* - M_{t-1})/M_{t-1}.$$

The results were not different enough to warrant further discussion.

Table 3

Anticipated vs. Unanticipated Monetary Base Growth and Equity Returns (1954:7 to 1972:3)[1]

Anticipated variable	Lag specification $\tilde{g}^u$	Lag specification $\tilde{g}^*$	F	$\bar{R}^2$	DW	Number of significant coefficients $\tilde{g}^u$	Number of significant coefficients $\tilde{g}^*$
$\tilde{g}_2^*$	16	—	.969	.078	1.83	1	—
$\tilde{g}_1^*$	16	—	.621	.051	1.78	0	—
$\tilde{g}_2^*$	16	0	1.014	.081	1.85	1	0
$\tilde{g}_1^*$	16	0	.614	.054	1.80	0	0
$\tilde{g}_2^*$	16	6	1.001	.113	1.83	0	0
$\tilde{g}_1^*$	16	6	.890	.102	1.85	0	1

[1]See note, table 1. Data are monthly observations.

One interpretation of these results is that future "unanticipated" monetary policy actions based on the autoregressive proxy are not in fact unanticipated. Information other than past monetary base growth — information that is available to the public and, if the reaction function specification is correct, information that does affect future money growth — may enable the public to correctly anticipate such future monetary base growth. Since the prediction of the reaction function already incorporates such available information, the public cannot forecast future unanticipated monetary base growth as measured by reaction function residuals; therefore, these future residuals do not affect current stock returns. Our results then are consistent with Rozeff's "reversed causation with correct anticipations" model, where the apparent effect of future monetary base growth on stock returns reflects the public's correct forecasts of future monetary base growth on the basis of currently available information.

Anticipated and Unanticipated Monetary Base Growth and Stock Returns

We discussed previously the Fama version of the model (equation 9), where both anticipated and unanticipated values of monetary policy should affect equity returns. In this section, we again use monetary policy reaction functions to differentiate anticipated and unanticipated policies. The model tested here is the empirical specification of the Fama model given by equation 10.

These estimates are presented in table 3. We use the same proxies for unanticipated money growth and, in this case, the corresponding measure of anticipated monetary base growth, as for the estimates in table 2. The table is divided into three parts: The first two lines include only unanticipated monetary base growth. The second two add only the concurrent anticipation of monetary base growth. The third pair allows up to six months lagged values of anticipations of future monetary base growth. In each of these, unanticipated monetary policy has the current as well as 16 lagged values.

The results are not inconsistent with the efficient market hypothesis, since unanticipated monetary base growth, current or lagged, has no significant effect on stock returns. According to equation 9, however, anticipated monetary base growth *should* have a positive effect on stock returns, if there is a constant expected real return and if anticipated monetary base growth affects money growth and, thereby, anticipated inflation. Our results do not show this effect and would seem to indicate that the expected real return on stocks is negatively affected by expected inflation that results from anticipated monetary base growth. This follows since the expected real return declines with anticipated inflation, unless there is an offsetting increase in the nominal return.[18]

[18]Fama uses a general equilibrium approach and concludes that real returns vary with expectations of future real economic activity. He also argues that apparent correlations between real stock returns and expected inflation or money growth rates are spurious. See Eugene F. Fama, "Stock Returns, Real Activity, Inflation, and Money," *American Economic Review* (September 1981), pp. 545-65.

Stock Returns and the Federal Funds Rate

If the monetary authority pegs the federal funds rate, the money supply becomes endogenous, and changes in the setting of the rate may be taken as an exogenous variable. In practice, the federal funds rate may change for reasons other than policy, especially over short intervals. Consequently, these tests may reflect not only how efficiently the market absorbs information about monetary policy but also the impact of other information embodied in movements in the federal funds rate. Nevertheless, they are useful in ascertaining how changes in the federal funds rate are internalized by the market during a period when the expressed policy was to maintain that rate within a narrow range.

In the model with monetary aggregates, anticipated inflation was approximated by anticipated monetary growth. It is less appropriate to think of anticipated changes in the federal funds rate as a proxy for anticipated inflation. However, changes in the anticipated federal funds rate that signal changes expected in financial markets will still provide important information in efficient markets. The tests in this section remain, therefore, as tests of market efficiency. They do, however, have less explicit theoretical development that explains exactly how monetary policy affects stock returns.

To split movements in the federal funds rate into anticipated and unanticipated components, we use the monetary policy reaction function estimated by Abrams, Froyen and Waud in which the federal funds rate is the dependent variable.[19] The fitted values from the estimated reaction function provide a measure of the anticipated federal funds rate (RF_t^*). The unanticipated portion of the federal funds rate (RF^u) is simply the actual federal funds rate minus the anticipated rate. The models we estimate using the federal funds rate as a measure of monetary policy again are those given by equations 2, 3, 4 and 10, where the unanticipated (g^u) or anticipated monetary policy variables (g^*) are now in terms of the federal funds rate. The results of these tests are given in table 4.

The results of estimating equations 2, 3 and 4 are shown in part A of the table. These results, using the interest rate as a measure of monetary policy, are less favorable to the efficient market hypothesis than our estimates using monetary aggregates. As can be seen from the first two lines of the table, lagged values of the unanticipated portion of the federal funds rate (lagged errors in forecasting the monetary authority's funds rate setting) appear to affect stock returns significantly. This evidence supports the view that stock returns lag *monetary policy* — even though our results in the previous section would indicate that stock returns do not lag *money growth*. The addition of current or future federal funds rate prediction errors does not increase the explanatory power of the equation (see estimates of equation 4 in the table).

In part B of the table, we report estimates of the model that allows both anticipated and unanticipated monetary policy to affect stock prices. Our estimates indicate that lagged values of both unanticipated and anticipated monetary policy as measured by the federal funds rate have significant effects on stock returns. Both here and in part A of the table, all the significant coefficients on the federal funds rate variables are negative (the signs of these coefficients are not reported in the table). This accords with the conventional expectation that a tightening of monetary policy, as measured by an increase in the federal funds rate setting, lowers stock prices and, hence, stock returns. In part B, as in part A of the table, however, the finding that past available information significantly affects stock returns raises questions about market efficiency.

This is not to say that the results in table 4 directly contradict the efficient market hypothesis. One interpretation of these results that is potentially consistent with the efficient market view is that the federal funds rate is a determinant of the *expected real* return on stocks, which is not a constant. With this interpretation, the *excess* return on stocks would still be independent of past available information, the condition for an efficient market. Still, the results in table 4 do suggest the possibility that while the market efficiently absorbed data on monetary aggregates, information carried by observations on the federal funds rate was not immediately reflected in stock prices and, hence, affected future stock returns.

[19]The anticipated federal funds rate is a function of 1) consistent forecasts of future values of the unemployment rate, the inflation rate and external balance variables and 2) lagged values of deviations of actual M1 from its target values. See Richard K. Abrams, Richard Froyen and Roger N. Waud, "Monetary Policy Reaction Functions, Consistent Expectations, and the Burns Era," *Journal of Money, Credit, and Banking* (February 1980), pp. 30-42.

Table 4

Anticipated vs. Unanticipated Monetary Policy (The Federal Funds Rate) and Equity Returns (1971:7 - 1976:6)[1]

A. Unanticipated Federal Funds Rate ($\tilde{R}F^u$)

Model	Lag (lead) specification	F	$\overline{R}^2$	DW	Number of significant coefficients: Lags	Number of significant coefficients: Leads
2	16 to 1	2.237*	.254	1.75	4	—
3	16 to 0	2.097*	.243	1.73	4	—
4	16 to (9)	1.580	.206	1.59	4	0

B. Anticipated ($\tilde{R}F^*$) and Unanticipated ($\tilde{R}F^u$) Lead Values and Lags of the Federal Funds Rate

Lags (leads): $\tilde{R}F^u$	Lags (leads): $\tilde{R}F^*$	F	$\overline{R}^2$	DW	Number of significant coefficients: $\tilde{R}F^u$	Number of significant coefficients: $\tilde{R}F^*$
9	—	2.661*	.223	1.76	2	—
9	0	3.355*	.309	2.00	3	1
9	6	3.677*	.440	2.08	2	2
9	(3)	3.058*	.332	2.02	4	0
9	3 to (3)	3.150**	.387	2.08	3	2

[1]See note, table 1. Data are monthly observations.

Tests with Weekly Money Stock Data

Earlier tests that split money growth into anticipated and unanticipated components are redone using weekly data. The measures of anticipated and unanticipated weekly money growth are taken from Naylor.[20] The time period for these tests is August 1974 to March 1977.

As noted, the use of weekly data provides a finer test of possible lead-lag relationships between money growth and stock returns. Data on the money supply generally were announced during our sample period on Thursday afternoons. Therefore, we assume an injection of monetary information occurs Thursday, which is new information to Friday's stock market transactions. By moving to a weekly model, we better capture these events. All money stock data used are the values originally announced on the Thursday of each week. The equity returns are derived from the stock prices recorded at market closing on the next day, Friday.

Table 5 presents the summary data from our weekly regression tests. The top of the table (part A) reveals that up to 16 lags and nine leads of unanticipated money growth explain very little of the variance in weekly stock returns. None of the individual coefficients are statistically significant at the 5 percent level of confidence. The F-values suggest that none of the three lag specifications leads to a rejection of the null hypothesis of market efficiency.

The bottom half of table 5 (part B) specifies past values of both anticipated ($\tilde{g}_1^*$) and unanticipated monetary growth ($\tilde{g}_1^u$) as determinants of the weekly equity returns. Adding six past weeks of anticipated monetary growth improves the explanatory power of the equation (with 16 lags of unanticipated money), doubling the $\overline{R}^2$ to .166. The main contribution in statistical significance comes from the current value of $\tilde{g}_1^*$ with less added by the one week lag (t-value equal to about −1.7). The signs of the estimated coefficients are negative, implying an *inverse* relationship between anticipated money growth and equity returns.[21]

[20]Naylor's forecasts are from a 52-week autoregressive scheme. This model is re-estimated one week at a time over the entire sample period and generates one-week-ahead forecasts. For details, see John A. Naylor, "Do Short-Term Interest Rate Expectations Respond to New Information on Monetary Growth?" *Southern Economic Journal* (January 1982), pp. 754-63.

[21]This finding would agree with the federal funds rate results if expectations of increased monetary growth are at least partially caused by earlier below-target growth. In this case, both higher expected money and higher federal funds rates would correlate with future falling stock returns.

Table 5

Anticipated and Unanticipated Monetary Growth and Equity Returns (1974:8 - 1977:3)[1]

A. Unanticipated Money Growth ($\tilde{g}^u$)

Model	Lag (lead) specification	F	$\bar{R}^2$	DW	Number of significant coefficients: Lags	Number of significant coefficients: Leads
2	16 to 1	.947	.084	2.02	0	—
3	16 to 0	.897	.085	2.02	0	0
4	16 to (9)	.890	.115	2.03	0	0

B. Anticipated ($\tilde{g}^*$) and Unanticipated Money Growth ($\tilde{g}^u$)

Lags: $\tilde{g}^u$	Lags: $\tilde{g}^*$	F	$\bar{R}^2$	DW	Number of significant coefficients: $\tilde{g}^u$	Number of significant coefficients: $\tilde{g}^*$
16	—	.897	.085	2.02	0	—
16	0	1.179	.115	2.03	0	1
16	6	1.302	.166	2.04	0	1

[1]See note, table 1. Data are weekly observations.

Overall, the results of weekly data indicate that information about money growth is quickly reflected in stock prices, as one would expect if the market is efficient.

CONCLUSIONS

The results of our study can be summarized as follows: Estimates of the relationship between stock returns and money growth rates, using monthly data, support the notion that stock markets are efficient. Even from week to week, the market seems to quickly utilize the most recent information on monetary aggregates. Our estimates of the relationship between stock returns and monetary policy actions as measured by the federal funds rate, however, suggest a possible violation of the conditions for market efficiency.

On the question of whether stock returns lead money growth, our results indicate that when anticipated money growth is a fitted value from a reaction function, future unanticipated money growth does not significantly affect current stock returns. But when future changes in money growth rates are based only on past money (using a third-order autoregressive scheme), they do significantly affect returns. This finding supports the hypothesis that the market uses information other than past money growth rates (information embodied in the reaction function prediction) to forecast future money growth and that such anticipations affect current stock returns.

This research has uncovered very little about how one can use monetary policy information for profit in the stock market. Information about aggregates is quickly assimilated by markets. The monthly estimations show little effect of anticipated or unanticipated aggregates (base or M1) upon stock returns. The weekly tests suggest that stock returns tend to fall within a week after the market anticipates a rise in the week's monetary aggregate. The most useful information seems to come from the monthly federal funds rate. We found that increases in that rate tended to lower stock returns over a six- to nine-month period. Since the federal funds rate is an imperfect indicator of monetary policy, this finding may say little about how monetary policy affects stock returns. It does, however, reveal that for our 1971-76 sample period, months when the federal funds rate fell were followed by periods of rising stock returns. Had market participants been aware of this relationship, they might have profited by it. Since the expressed policy of the Federal Reserve today allows the federal funds rate to float within a wide band, there is no indication that this relationship continues. The relationship between monetary growth or movements in the federal funds rate and stock returns in the post-October 1979 period is a subject for future research.

Monetary Policy and Short-Term Real Rates of Interest

R. W. HAFER and SCOTT E. HEIN

TEXTBOOK descriptions of the channels of monetary policy's impact on the economy usually outline a two-step procedure: "The first is that an increase in real balances generates a *portfolio disequilibrium* — at the prevailing interest rate and level of income, people are holding more money than they want. This causes portfolio holders to attempt to reduce their money holdings by buying other assets, thereby changing asset yields. In other words, the change in the [real] money supply changes [real] interest rates. The second stage of the transmission process occurs when the change in interest rates affects aggregate demand."[1]

The rational expectations literature, however, has raised serious questions about this description, especially the first stage wherein an increase in real money balances lowers expected real interest rates. Shiller, for example, drawing from previous work in rational expectations, hypothesizes that the expected real interest rate is unaffected by changes in monetary policy.[2]

While Shiller found little support for this hypothesis, other recent empirical work supports it. Fama, for instance, is unable to reject the hypothesis that the expected real rate on short-term financial assets was constant over much of the post-Accord period in the United States.[3] This hypothesis is even stronger than Shiller's. It holds that monetary actions, as well as everything else, have had no systematic effect on expected real interest rates.

This article re-evaluates the evidence suggesting that the expected (*ex ante*) real interest rate on short-term financial assets is constant. Evidence is provided that allows us to reject this hypothesis for the 1955-79 period. Following this, data are examined to determine whether evidence supports the typical textbook description in which changes in expected real interest rates are associated with changes in real money growth.

THE FRAMEWORK OF ANALYSIS

Consider first the relationship between nominal interest rates and inflation expectations embodied within the so-called Fisher relationship,[4]

$$(1)\quad i_t = r_t^e + \dot{P}_t^e,$$

where i_t is a nominal (or market) rate of interest (the rate measuring how many *dollars* must be repaid in the future for a given dollar loaned today), r_t^e is the expected real interest rate (the rate measuring how

[1]Rudiger Dornbusch and Stanley Fischer, *Macroeconomics* (McGraw-Hill, 1978), p. 120.

[2]Robert J. Shiller, "Can the Fed Control Real Interest Rates?" in Stanley Fischer, ed., *Rational Expectations and Economic Policy* (The University of Chicago Press, 1980), pp. 117-56. Shiller also outlined two other (non-exclusive) hypotheses: (1) the Fed can affect real rates only through *unexpected* policy moves and (2) Fed policies known far enough ahead of time have no effect on real rates. These hypotheses are not as stringent as the hypothesis considered in this paper.

[3]Eugene F. Fama, "Short-Term Interest Rates as Predictors of Inflation," *American Economic Review* (June 1975), pp. 269-82.

[4]The belief in a positive relationship between expected inflation and nominal interest rates has a long history in economics. Henry Thornton recognized the relationship as early as 1811. Alfred Marshall also acknowledged the link during the latter half of the 19th century. Even so, the intensity with which Irving Fisher examined the relationship during his career has resulted in the distinction of equation 1 being dubbed the "Fisher equation." See Henry Thornton, "Two Speeches of Henry Thornton, esq. on the Bullion Report, May 1811," in F. A. v. Hayek, ed., *An Enquiry into the Nature and Effects of the Paper Credit of Great Britain (1802)*, (August M. Kelley, 1962), pp. 323-62; Alfred Marshall, "Remedies for Fluctuations of General Prices (1887)," in A. C. Pigou, ed., *Memorials of Alfred Marshall* (Kelley & Millman, Inc., 1956) pp. 188-211; and Irving Fisher, *The Theory of Interest* (Kelley & Millman, Inc., 1954), especially Chapter 2.

many more *goods* can be obtained in the future by foregoing consumption today) and $\dot{P}_t^e$ is expected inflation (the rate at which the dollar price of goods is expected to rise.) Equation 1 represents a hypothesized *equilibrium* relationship. It posits that changes in observed nominal rates of interest fully reflect changes in expected inflation, *holding the expected real rate constant.*[5] In other words, nominal rates and expected inflation are positively related and, ceteris paribus, move on a one-to-one basis.

The foundation for this equilibrium relationship is the view that investors have two possible investment opportunities: they can invest either in capital goods that produce a future stream of consumption goods or in financial assets denominated in monetary terms. Investment in capital goods is expected to produce r_t^e percent more consumption goods per year than the amount of consumption goods originally given up to produce the capital good. To make the return on investing in the capital good comparable to the alternative investment (the financial asset), the value of the future stream of consumption goods must be translated into dollar terms. This is accomplished by adding the expected rate of change in the dollar price of consumption goods ($\dot{P}_t^e$) to the rate of increase of consumption goods (r_t^e). The right-hand side of equation 1, therefore, represents the expected dollar return from investing in a capital good.

In equilibrium (and without differential tax rates), the dollar return from investing in capital goods should equal the dollar return from investing in financial assets, measured by the nominal interest rate, i_t. Equation 1 thus states that an individual should not find the dollar yield on financial assets any different from the expected dollar yield on capital goods. We stress that equation 1 is an equilibrium condition: not only are the financial and capital goods markets hypothesized to be individually in equilibrium, but any differential in the expected real yields in these two markets is arbitraged away.

In its present form, equation 1 cannot be examined empirically because the two variables on the right-hand side, the expected real rate of interest and inflation expectations, are not directly observable. While there are many observable nominal interest rates on financial assets, there are no reliable aggregate measures of either the expected real yield on capital goods or the expected future inflation rate.[6]

IS THE EXPECTED REAL RATE OF INTEREST CONSTANT?

To test the relationship specified by equation 1, one can make two assumptions: First, assume that the expected real interest rate is a constant, such that

(2) $r_t^e = \bar{r}$.

Second, to circumvent the problem of measuring inflation expectations, assume that next period's *actual* inflation ($\dot{P}_{t+1}$) is equal to what is currently expected (at time t), plus a random disturbance μ_{t+1}, where μ_{t+1} is independent and distributed $N(0, \sigma^2)$:

(3) $\dot{P}_{t+1} = \dot{P}_t^e + \mu_{t+1}$.

This relationship specifies that one-period-ahead inflation forecasts are unbiased; *on average* the actual inflation rate over the next time period will be the expected rate.

Substituting equations 2 and 3 into 1 yields

(4) $i_t = \bar{r} + \dot{P}_{t+1} - \mu_{t+1}$.

This equation can be arranged to test empirically the hypothesis that today's interest rate accurately predicts tomorrow's inflation as follows:

(5) $\dot{P}_{t+1} = -\bar{r} + \beta_0 i_t + \mu_{t+1}$.

Assuming that financial markets are efficient, we would expect to find β_0 not to be statistically different from *unity* and the estimated constant term to be *negative.* If the estimated coefficient β_0 is not statistically different from unity, the proposition that current interest rates fully reflect the market's anticipations of the future inflation rate cannot be rejected. Similarly, if the estimated constant term is negative, the expected real rate of return is then positive as suggested by the underlying economic theory. More-

[5]This equilibrium relationship also should include the cross-product term $r_t^e \dot{P}_t^e$. Like most empirical analyses of this relationship, we ignore this term, assuming that the magnitude of the variable is sufficiently small.

[6]Some researchers have attempted to investigate the relationship by using directly observed inflation expectations data generated from Joseph A. Livingston's biannual survey of economists. See, for example, William E. Gibson, "Interest Rates and Inflationary Expectations: New Evidence," *American Economic Review* (December 1972), pp. 854-65; David H. Pyle, "Observed Price Expectations and Interest Rates," *Review of Economics and Statistics* (August 1972), pp. 275-80; Kajal Lahiri, "Inflationary Expectations: Their Formation and Interest Rate Effects," *American Economic Review* (March 1976), pp. 124-31; Thomas F. Cargill,"Anticipated Price Changes and Nominal Interest Rates in the 1950's," *Review of Economics and Statistics* (August 1976), pp. 364-67; John A. Carlson, "Short-Term Interest Rates as Predictors of Inflation: Comment," *American Economic Review* (June 1977), pp. 469-75; and Douglas K. Pearce, "Comparing Survey and Rational Measures of Expected Inflation: Forecast Performance and Interest Rate Effects," *Journal of Money, Credit and Banking* (November 1979), pp. 447-56.

Table 1

Empirical Estimates of Equation 5[1]

Coefficient	I/1955-IV/1979	I/1955-IV/1959	I/1960-IV/1969	I/1970-IV/1979
	Ordinary Least-Squares Estimates			
Constant	−0.580	2.686	−1.496	1.393
	(1.46)	(3.10)	(2.65)	(1.42)
β_0	1.056	−0.041	1.073	0.840
	(13.49)	(0.13)	(7.97)	(5.61)
$\bar{R}^2$	0.646	−0.055	0.616	0.439
SE	1.630	1.190	1.116	1.750
DW	1.02	1.63	1.92	1.09
	Generalized Least-Squares Estimates			
Constant	−0.126	2.584	−1.496	1.586
	(0.20)	(2.68)	(2.65)	(1.15)
β_0	0.957	−0.001	1.073	0.797
	(8.00)	(0.00)	(7.97)	(3.93)
$\bar{R}^2$	0.389	−0.056	0.616	0.270
SE	1.424	1.169	1.116	1.573
DW	2.21	2.15	1.92	2.04
$\hat{\rho}$	0.504	0.190	0.000	0.455

[1]$\bar{R}^2$ represents the coefficient of determination adjusted for degrees of freedom, SE is the regression standard error, DW is the Durbin-Watson test statistic and $\hat{\rho}$ is the estimate of the autocorrelation coefficient. Absolute value of t-statistics appear in parentheses.

over, the existence of serial correlation in the residuals would deny the assumption embodied in equation 3 and, consequently, would lead to a rejection of the hypothesis specified in equation 5.[7]

Previous empirical studies generally have not explicitly considered the temporal stability of the expected real rate within this framework. The constant term in equation 5 represents the estimate of the (negative value of the) expected real rate of return. The above theoretical foundation for this specification suggests that, in addition to being negative, this term is statistically time-invariant. Thus, a test of the temporal stability of the constant term is also a test of the constancy of the expected real interest rate.

Table 1 presents estimates of equation 5 for various periods. The inflation data used to estimate equation 5 are based on quarterly observations of the GNP deflator, expressed as annual rates of change.[8] Since the GNP deflator provides an average measure of prices over the quarter, the quarterly average three-month Treasury bill rate is used as the nominal interest rate measure.

Consider first the results obtained by estimating equation 5 over the full sample period, I/1955-IV/1979. The constant term is negative (although not significantly different from zero), and the coefficient on the interest rate variable is not statistically different from unity as suggested by the theory. Unfortunately, the low Durbin-Watson statistic provides evidence of first-order serial correlation.[9]

[7]Fama tested and rejected the hypothesis that the expected real rate was linearly related to expected inflation. Critical examinations of Fama's results are found in Carlson, "Short-Term Interest Rates as Predictors of Inflation"; Douglas Joines, "Short-Term Interest Rates as Predictors of Inflation: Comment," *American Economic Review* (June 1977), pp. 476-77; and Charles R. Nelson and G. William Schwert, "Short-Term Interest Rates as Predictors of Inflation: On Testing the Hypothesis that the Real Rate of Interest is Constant," *American Economic Review* (June 1977), pp. 478-86. Also, see Eugene F. Fama, "Interest Rates and Inflation: The Message in the Entrails," *American Economic Review* (June 1977), pp. 487-96.

[8]The GNP deflator is used to avoid recent problems with the consumer price index. For a discussion of problems with this index, see Alan S. Blinder, "The Consumer Price Index and the Measurement of Recent Inflation," *Brookings Papers on Economic Activity* (2:1980), pp. 539-65.

[9]Eugene F. Fama and Michael R. Gibbons, "Inflation, Real Returns and Capital Investment," Working Paper No. 41 (Graduate School of Business, University of Chicago, 1980), also find evidence of serially correlated disturbance terms when quarterly data are employed. In addition, in that study as well as in his "Stock Returns, Real Activity, Inflation, and Money," *American Economic Review* (September 1981), pp. 545-65, Fama drops the assumption that the expected real rate of interest is constant. Both studies estimate the inflation/interest rate relationship *assuming* that the expected real rate is a random walk.

This result, by itself, is enough to reject the framework in equation 5.[10] Focusing solely on the constancy of the expected real rate, however, the accompanying estimation problem can be corrected by using generalized least-squares estimation. These results appear in the lower half of table 1.

The full sample results reported there again indicate that next period's rate of inflation does mirror, one-for-one, a rise in today's interest rates. Moreover, the constant term remains insignificantly different from zero.

Table 1 further reports estimation results for subperiods arbitrarily truncated at the end of each decade. If the expected real rate of interest is temporally invariant, the constant terms in these subperiods should not differ statistically. Yet, as the table immediately shows, they do differ significantly across the various subperiods shown. In fact, the estimated constant term is *positive* and significant in the first subperiod (late 1950s), while not different from zero in the last decade (1970s). It has the anticipated negative sign only in the decade of the 1960s. Moreover, the coefficient on the interest rate variable is not statistically different from zero in the late 1950s, even though theory suggests that it should equal unity. Thus, the coefficient estimates, as well as summary statistics such as the $\overline{R}^2$ and the standard errors of the equation, vary substantially across subperiods, irrespective of the estimation technique used.

The statistical significance of the variation in the constant term (the estimate of the *ex ante* real interest rate) can be investigated by including dummy variables for possible shifts in the intercept. Thus, equation 5 was re-estimated with two dummy variables: D1 equal to 1 for I/1955-IV/1959 and D2 equal to one for I/1960-IV/1969. Estimating such an equation with ordinary least squares again yielded residuals that were significantly autocorrelated. To improve hypothesis testing, the equation was estimated using a generalized least-squares routine to correct for assumed first-order autocorrelation. The I/1955-IV/1979 estimation results are (absolute value of t-statistics in parentheses):

$$(5') \quad \dot{P}_{t+1} = \underset{(1.62)}{1.40} - \underset{(1.19)}{0.88\,D1} - \underset{(3.46)}{1.88\,D2} + \underset{(6.51)}{0.83\,i_t}$$

$$\overline{R}^2 = 0.55 \qquad SE = 1.37 \qquad DW = 2.07 \qquad \hat{\rho} = 0.35$$

[10]For market efficiency, past values of the disturbance, since they are past inflation forecast errors and are therefore known, should provide no additional help in assessing future inflation beyond that already incorporated in market interest rates. See Fama, "Short-Term Interest Rates," p. 273, for a discussion of this aspect.

These results support the previous subperiod findings: the estimated real interest rate is significantly positive only in the 1960s. The point estimates of the expected real interest rate for the 1950s, 1960s and 1970s, respectively, are −0.52, + 0.48 and −1.40. While the point estimates for the 1950s and the 1970s are negative, they are not significantly different from zero. On the other hand, the positive point estimate for the 1960s is significantly different from zero. Thus, the hypothesis that the expected real interest rate has been constant over the past 25 years must be rejected.

EX POST REAL RATES: FURTHER CONSTANCY TESTS

Equation 4 can be rewritten as

$$(6) \quad i_t - \dot{P}_{t+1} = \bar{r} - \mu_{t+1}.$$

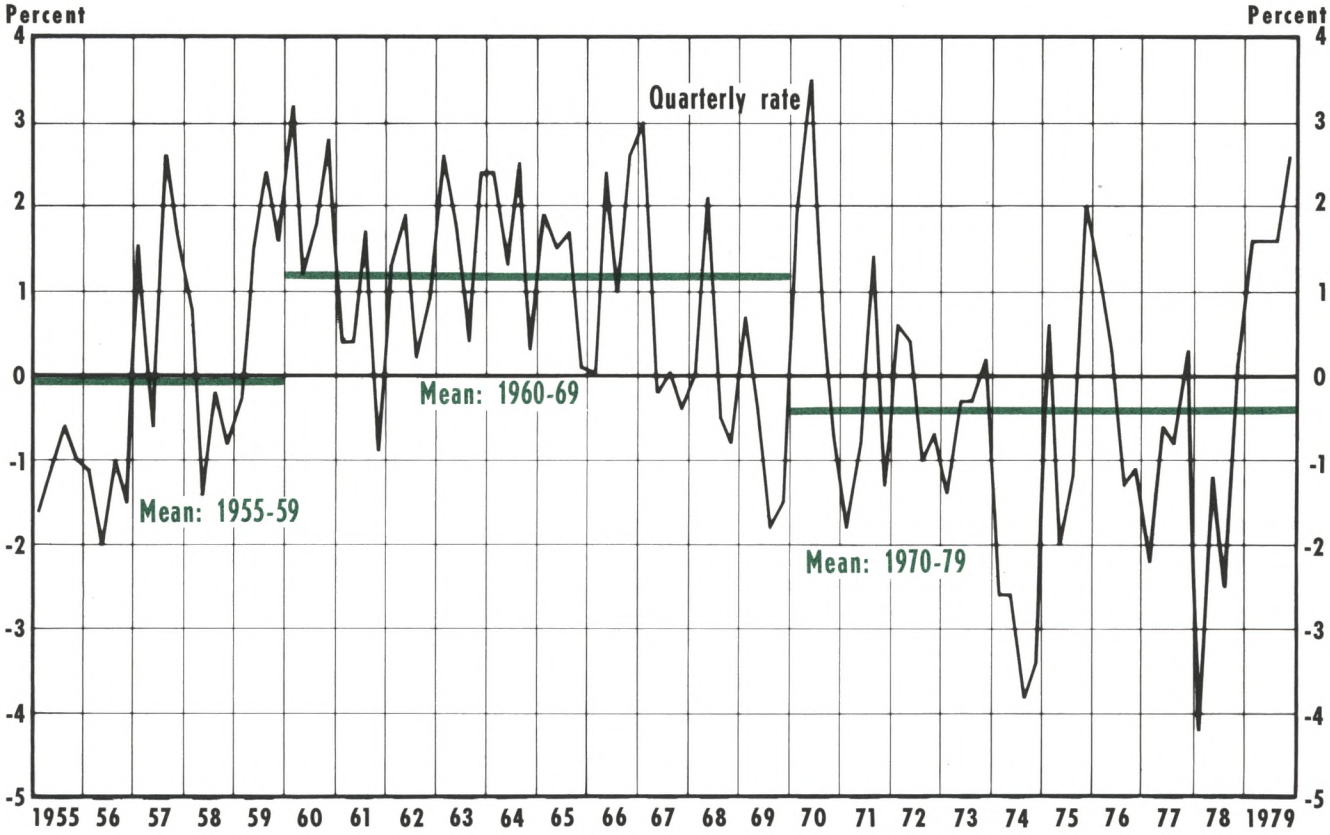
This equation states that the *ex post* real rate should equal a constant (the *ex ante* real rate), minus a white noise random error term.[11] A feel for the statistical variation in the real rate can be obtained by plotting its behavior for our sample period. Chart 1 shows the quarterly *ex post* real rate for the I/1955-IV/1979 period and its mean values for the I/1955-IV/1959 (−0.03), I/1960-IV/1969 (1.21) and I/1970-IV/1979 (−0.39) subperiods. If equation 6 holds for the whole period, the means across subperiods should be equal, since the expected value of the disturbance term in each subperiod is zero.

Tests for equality of the *ex post* real interest rate means across the subperiods provide another investigation of the constancy hypothesis. Such tests again lead to a rejection of this hypothesis. The t-statistic,

[11]This measure of the *ex post* real rate is somewhat different from that used by others. Many take the difference between *today's* interest rates and *today's* inflation rate as an *ex post* real rate measure. Theory suggests, however, that the preferable measure is the difference between *today's* interest rates and *tomorrow's* inflation.

In the test subsequently developed and others which follow, interest rates are assumed to adjust one-for-one with inflation expectations, a hypothesis that can be rejected in equation 5′. The reader should be cautioned that there are counter theoretical arguments and some empirical evidence to suggest that the nature of the U.S. tax system has invalidated this relationship, with interest rates rising more than one-for-one with an increase in inflation expectations. For theoretical discussions, see Michael R. Darby, "The Financial and Tax Effects of Monetary Policy on Interest Rates," *Economic Inquiry* (June 1975), pp. 266-76; and Martin Feldstein, "Inflation, Income Taxes, and the Rate of Interest: A Theoretical Analysis," *American Economic Review* (December 1976), pp. 809-20. For empirical evidence on the matter, see John A. Carlson, "Expected Inflation and Interest Rates," *Economic Inquiry* (October 1979), pp. 597-608.

Chart 1
Short-Term Ex Post Real Rate of Interest



used to test whether the mean *ex post* real rate for the latter half of the 1950s is equal to that of the 1960s, is 3.67, sufficiently large to reject the null hypothesis at the 5 percent significance level. Further, the t-statistic used to test the equality of mean *ex post* real rates in the 1960s relative to the 1970s is 4.86, again allowing rejection of the null hypothesis of constant real interest rates at the 5 percent level. Thus, if one accepts the propositions that interest rates move in direct proportion with expected inflation and that inflation expectations are unbiased, one must reject the constancy of the *ex ante* real interest rate over the subperiods investigated.

MONETARY POLICY AND THE EXPECTED REAL RATE

These findings suggest that the real interest rate has not been constant over the past 25 years. In this light, is there any evidence that links the real rate of interest to monetary policy? After all, the textbook description of monetary policy's transmission mechanism relates changes in the real rate to changes in real money balances. In particular, it maintains that an increase in real money balances lowers expected real rates, at least temporarily.

The previous framework, linking *ex post* and *ex ante* real rates, can be used to address this issue. If inflation expectations are unbiased and financial markets are efficient, then the *ex post* real rate ($i_t - \dot{P}_{t+1}$) is equal to the *ex ante* real rate (r_t^e), minus a random disturbance term (μ_{t+1}) capturing unexpected inflation:

$$(7)\quad i_t - \dot{P}_{t+1} = r_t^e - \mu_{t+1}.$$

The typical textbook relationship can be represented as

$$(8)\quad r_t^e = \beta_0 + \beta_1 (M_t/P_t) + \beta_2 (M_{t-1}/P_{t-1}) + \ldots + \epsilon_t,$$

where M is the nominal money stock, P is the price

level and ϵ is a random error term. This relationship represents the hypothesis that the expected real rate is related to real money balances. Since nothing in macroeconomic theory indicates how long it takes for changes in monetary policy to have an effect, lagged real balances are included in an effort to capture empirically the dynamics of the process. Theory does suggest, however, that some of the coefficients should be significantly negative. While it is impossible to estimate equation 8 because of a lack of observations on r_t^e, equation 7 indicates that we have a close approximation in the *ex post* real rate. Combining equations 7 and 8, we get

$$(9)\quad i_t - \dot{P}_{t+1} = \beta_0 + \beta_1 (M_t/P_t) + \beta_2 (M_{t-1}/P_{t-1}) + \ldots + \epsilon_t - \mu_{t+1}.$$

Equation 9 was estimated initially by arbitrarily trying 10 lags on real money balances in the relationship. Regardless of the sample period considered, however, the only coefficients that were statistically different from zero in any consistent fashion were those for the contemporaneous and first-lagged real money balances. Thus, results including only these two variables are reported.

Estimates of equation 9 over the full sample period (I/1955-IV/1979) and most subperiods provide evidence of significant first-order autocorrelation in the residuals. Consequently, the relationship was reestimated using a generalized least-squares technique to correct for this problem. The resulting full-sample coefficient estimates and summary statistics are (absolute value of t-statistics in parentheses):[12]

$$(10)\quad i_t - P_{t+1} = \underset{(1.73)}{5.00} - \underset{(2.68)}{0.89}\,(M/P)_t + \underset{(2.48)}{0.83}\,(M/P)_{t-1}$$

$$\bar{R} = 0.07 \quad SE = 1.37 \quad DW = 2.14 \quad \hat{\rho} = 0.56 \quad F(2,97) = 4.95$$

While the variation in the *ex post* real rate explained by the equation is small, it is statistically significant. Moreover, the coefficient estimates are consistent with the textbook transmission mechanism. An increase in real money balances is associated with a statistically significant, contemporaneous decline in short-term real rates during this period. Further, the results are consistent with the long-run policy ineffectiveness of increasing real balances to reduce real interest rates.[13] The coefficient estimate for real money balances lagged one period is significantly positive and is not statistically different from the absolute value of the coefficient on contemporaneous real money balances. This finding indicates that a current increase in real money balances will be associated with a current decline in real rates, but followed by a rise in real rates of equal size at time t+1. This suggests that monetary authorities, to the extent that they can change real balances, cannot *permanently* affect real rates of interest.

While earlier evidence showed that the *ex post* real rate ($i_t - \dot{P}_{t+1}$) behaved differently across subperiods, there is little evidence to suggest that its relationship to real money balances has changed over the period. For example, a conventional Chow test evaluating a hypothesized break in the relationship at IV/1969 yields a calculated F-statistic of $F(3,94) = 0.39$, well below the 5 percent critical value of 2.70. Thus, the regression coefficients are not statistically different before or after IV/1969[14] Changes in real balances have the same statistical effect on real interest rates across the sample period.

Finally, it is appropriate to note that the estimated relationship implies a *positive* relationship between the volatility in real money balances and the volatility in real interest rates. If the frequency of change in real money balances increases, the estimated relationship implies an increase in the frequency of change in real interest rates. The evidence presented here suggests that more stable real money growth, even over periods as short as a quarter, will produce a more stable pattern of real interest rate movements.[15]

[12]Money (M/P) is measured (in billions of 1972 dollars) by the adjusted monetary base for all results reported here. Thus, the empirical results indicate that a $1 billion increase in real balances will reduce the real interest rate by 89 basis points in the current period. This decline is offset, however, by an 83 basis-point increase in the real rate in the subsequent period. We also tried the M1 measure and obtained similar results.

[13]We do not mean to suggest that monetary authorities can control real money balances over long periods of time. On this point, see Denis S. Karnosky, "Real Money Balances: A Misleading Indicator of Monetary Actions," this *Review* (February 1974), pp. 2-10.

[14]In addition, we tested the hypothesis that the variance of the error term was larger in the 1970s than in the earlier period. The calculated F-statistic (with 37 and 57 degrees of freedom, respectively) was 1.44, less that the 5 percent critical value of 1.59. Thus, the hypothesis of equal variance across these two periods cannot be rejected.

[15]An interesting investigation into the effects of monetary policy on both short- and long-term real interest rates is provided in Dean W. Hughes and Duane Weimer, "The Impact on Business Investment of the Federal Reserve System's Operating Procedures," Federal Reserve Bank of Kansas City *Economic Review* (February 1982), pp. 14-25.

CONCLUSION

This article has provided evidence counter to the hypothesis that the expected real rate of return on short-term financial assets was constant over the period 1955-79. If such a hypothesis were valid, monetary policy would be powerless in affecting real economic activity through the conventional transmission mechanism. While rejecting the constancy hypothesis, this article also provides evidence consistent with conventional macroeconomic theory whereby increases in real money balances temporarily lower expected real rates. This effect is contemporaneous on a quarterly basis. While such an effect is significant, it is relatively small and is offset in the following quarter by an identical rise in expected real rates. Thus, there is no evidence of a *long-run* effect running from changes in real money balances to changes in real interest rates. Finally, the evidence presented here suggests that more volatile short-run real money growth is likely to produce more volatile real interest rate fluctuations. Thus, contrary to recent claims, stable money growth and stable interest rates are hardly inconsistent policy objectives.[16]

[16]For another view, see Bryon Higgins, "Should the Federal Reserve Fine Tune Monetary Growth?" Federal Reserve Bank of Kansas City *Economic Review* (January 1982), pp. 3-16.

Central Banks' Demand for Foreign Reserves Under Fixed and Floating Exchange Rates

DALLAS S. BATTEN

THE international monetary system has experienced significant changes during the 1970s. The most dramatic of these has been the transformation from a system of pegged exchange rates to one in which central banks make no institutional commitment to maintain a particular exchange rate. Despite this change, central banks have been unwilling, in general, to allow their exchange rates to be completely market-determined and, consequently, continue to hold foreign reserves. The primary focus of this article is to analyze central banks' demand for foreign reserves in light of this institutional change.

Central banks generally are thought to hold stocks of foreign reserves so their economies can avoid incurring the costs of adjusting to every international imbalance that would be transmitted to the domestic economy through changes in exchange rates. In particular, before March 1973, central banks participating in the Bretton Woods Agreement were compelled to hold foreign reserves because they were committed to intervene in foreign currency markets when the value of their currencies moved outside a predetermined range.

It was commonly believed that the demise of the Bretton Woods Agreement and the concomitant greater flexibility of exchange rates would reduce central banks' intervention in foreign currency markets and, consequently, reduce their demand for foreign reserves. That is, since perhaps the single, most important reason for holding reserves had diminished, central banks would not be expected to hold such large stocks of foreign reserves as they had under the fixed exchange rate system. In spite of this expectation, however, central banks have continued to maintain sizable stocks of reserves since March 1973. This observation has led researchers to conclude that central banks have not changed appreciably their demand for reserves with the transition from a fixed to a floating exchange rate system.[1]

This conclusion, though potentially accurate, is founded on a framework of analysis in which foreign reserves are considered by central banks as a very special type of asset — one held *solely* to enable them to intervene in foreign currency markets. However, there is an alternative framework for analyzing central bank behavior that predicts that, even if all countries had adopted a completely clean-floating exchange rate system in 1973, central banks would have continued to hold a variety of financial assets, some of which would have been classified as foreign reserves under the previous fixed exchange rate system. This article investigates which of these competing frameworks better explains central bank behavior since March 1973.

TWO MODELS OF CENTRAL BANK BEHAVIOR

To analyze whether or not central bank behavior has changed significantly since the introduction of flexible exchange rates, the demand for reserves based on the intervention motive is compared with an alternative one developed within an asset-choice

The author would like to thank John Bilson, Michael Bordo and Ed Ray for their comments on an earlier draft.

[1]See, for example, Jacob A. Frenkel, "International Reserves: Pegged Exchange Rates and Managed Float," in Karl Brunner and Allan H. Meltzer, eds., *Public Policies in Open Economies*, Carnegie-Rochester Conference Series on Public Policy, supplement to the *Journal of Monetary Economics*, Volume 9 (1978), pp. 111-40; H. Robert Heller and Mohsin S. Kahn, "The Demand for International Reserves Under Fixed and Floating Exchange Rates," International Monetary Fund *Staff Papers* (December 1978), pp. 623-49; Nasser Saidi, "The Square-Root Law, Uncertainty and International Reserves Under Alternative Regimes," *Journal of Monetary Economics* (May 1981), pp. 271-90.

framework.[2] Only if the former explanation outperforms the latter for the floating period can one conclude that the changes in behavior since 1973 have been relatively minor and inconsequential.

The first model is the standard one based on the derived demand for foreign reserves for purposes of intervening in foreign exchange markets. Since this model has appeared frequently in the literature, its characteristics are only briefly described.[3] The second model is based on asset-choice behavior and has not been applied, until now, to the analysis of foreign reserve demand. In this model, foreign reserves are treated as one of several assets that appear in a bank's portfolio and are held for the general conduct of monetary policy.

The Intervention Model

The central bank intervention motive has been thoroughly investigated. Earlier studies typically have employed an optimizing approach in determining the demand for foreign reserves. One procedure is to find the stock of reserves at which the marginal costs of holding them equal the marginal benefits of using them to intervene in foreign currency markets (i.e., the avoidance of costs associated with the domestic economy having to adjust to each external shock). A second procedure is conducted in terms of welfare maximization under uncertainty. In particular, a central bank's demand for foreign reserves is the result of its maximizing a societal welfare function which is a positive function of the expected level of real income and a negative function of its variability. Since the holding of foreign reserves diverts resources away from domestic uses, the larger the stock of reserves, the lower the expected level of real income. However, if no reserves are held, the domestic economy would have to adjust to every external shock, resulting in more real income variability.

Employing the intervention motive within this framework, previous studies have identified four major determinants of reserve demand: the variability of international payments and receipts, the propensity to import, the opportunity cost of holding reserves and a scale variable measuring the size of international transactions (usually the value of imports). The variability of receipts and payments measures the likelihood that external disequilibrium will occur, inducing the central bank to intervene in foreign currency markets in order to mitigate the impact of this disequilibrium on domestic markets. The larger the variability of a country's receipts and payments, the more susceptible is that country to external disequilibrium; consequently, the larger is the optimal stock of reserves desired for purposes of intervention.

There are two possible rationales for including the propensity to import as a determinant of reserve demand. First, the average propensity to import can be considered a measure of the degree of openness in an economy, thus indicating the degree to which the economy is vulnerable to an external disequilibrium. A second, alternative rationale stems from the Keynesian model of an open economy in which an external disequilibrium could be corrected, without changing the exchange rate, by a change in output in proportion to the foreign trade multiplier. This output cost of adjustment could be avoided if the central bank used its stock of foreign reserves to finance (or to sterilize) the disequilibrium. Since this output cost is directly related to the size of the foreign trade multiplier, and since this multiplier is inversely related to the marginal propensity to import, the output cost of not holding sufficient reserves necessary to avoid this adjustment and, thus, the central bank's demand for reserves, must also be inversely related to the marginal propensity to import. Because the marginal propensity to import is difficult to measure, most studies have substituted the average propensity as a proxy. However, if the average propensity to import is employed both as a proxy for the marginal propensity and as a measure of

---

[2]See Russell S. Boyer and David Laidler, "A Comment on the Frenkel Paper," in Brunner and Meltzer, eds., *Public Policies in Open Economies*, pp. 141-43.

[3]Examples of this and similar models include Peter B. Clark, "Demand for International Reserves: A Cross-Country Analysis," *Canadian Journal of Economics* (November 1970), pp. 577-94; Peter B. Clark, "Optimum International Reserves and the Speed of Adjustment," *Journal of Political Economy* (March/April 1970), pp. 356-76; T. J. Courchene and G. M. Youssef, "The Demand for International Reserves," *Journal of Political Economy* (August 1967), pp. 404-13; Jacob A. Frenkel, "The Demand for International Reserves by Developed and Less-Developed Countries," *Economica* (February 1974), pp. 14-24; Frenkel, "International Reserves: Pegged Exchange Rates and Managed Float"; H. Robert Heller, "Optimal International Reserves," *Economic Journal* (June 1966), pp. 296-311; Heller and Khan, "The Demand for International Reserves Under Fixed and Floating Exchange Rates"; F. Steb Hipple, *The Disturbance Approach to the Demand for International Reserves*, Princeton Studies in International Finance No. 35 (Princeton University Press, 1974); Milton A. Iyoha, "Demand for International Reserves in Less-Developed Countries: A Distributed Lag Specification," *The Review of Economics and Statistics* (August 1976), pp. 351-55; Michael G. Kelly, "The Demand for International Reserves," *The American Economic Review* (September 1970), pp. 655-67; and Saidi, "The Square-Root Law, Uncertainty and International Reserves."

openness, the sign of its impact on reserve demand is ambiguous.

Since central banks do not hold an infinite stock of foreign reserves, there must be some cost associated with holding them. Conceptually, from society's point of view, holding foreign reserves represents an allocation of scarce resources away from domestic uses. Presumably, for every dollar invested in its stock of foreign reserves (through its central bank), society foregoes a dollar of domestic capital formation. Consequently, a rate of return on domestic capital is the appropriate measure of the opportunity cost to society of its central bank's stock of foreign reserves. On the margin, the optimal stock of reserves is that level at which the cost of holding reserves equals the marginal benefits provided by that stock of reserves. Few studies have included explicitly a measure of opportunity cost. Moreover, those that have included it have not found it to be empirically significant.[4] The hypothesized reason for the overall poor performance of this variable is the strong positive relationship between it and the supply of reserves. In particular, the higher the opportunity cost of holding reserves, the higher also the domestic rate of return on financial capital which motivates capital inflows and, ceteris paribus, increases the supply of reserves. As described below, interest rate differentials are employed as an attempt to circumvent this problem.

Finally, the scale variable and the demand for foreign reserves should be positively related. In fact, if the value of international transactions is used as the scale variable, the elasticity of reserve demand with respect to the value of international transactions should be between 0.5 and 1.0.[5]

An Asset-Choice Model

In formulating an asset-choice model of central bank behavior, foreign reserves are treated simply as one type of asset in a central bank's portfolio held to enable the central bank to conduct domestic monetary policy. It is assumed that the primary objective of monetary policy is to provide an economic environment conducive to the stable, noninflationary growth of real output. To this end, the central bank affects the level of commercial bank reserves (and, subsequently, the money supply) through activity in government securities and foreign currency markets and by making loans directly to the banking sector. Consequently, to conduct monetary policy adequately, its portfolio should contain at least three assets: foreign reserves, government securities and claims on commercial banks.

A central bank typically confronts two types of economic phenomena — expected and unexpected — to which it makes policy responses. In light of this, the specific modeling of the portfolio decision-making process of a central bank involves separating its assets into two categories: committed and uncommitted assets. In response to its anticipations of prospective events, a central bank commits a portion of its portfolio so that it can pursue its monetary policy objective within this "expected" economic environment.

However, since a central bank also is faced with unanticipated economic events to which it may wish to respond, it must hold additional reserves to enable it to respond to these "unexpected" occurrences (or shocks) as well. These "precautionary" reserves may or may not be used for the conduct of monetary policy in any specific period, while the committed portion, is, by definition, fully involved in the monetary control process. Consequently, a central bank is concerned only with the yield (cost) on the potentially idle, precautionary portion. That is, a central bank's demand for the assets that form the committed component is hypothesized to be insensitive to their relative yields, whereas the composition of the precautionary (or uncommitted) reserve component is hypothesized to be sensitive to changes in relative asset yields.

To formalize this discussion of central bank behavior, assume that a central bank (subject to certain constraints) desires to maximize its "ability" to respond to unanticipated events. It accomplishes this by maximizing the *uncommitted* portion of its portfolio.[6] This can be summarized with the fol-

[4]See, for example, Courchene and Youssef, "The Demand for International Reserves"; Iyoha, "Demand for International Reserves in Less-Developed Countries"; Kelly, "The Demand for International Reserves"; and Saidi, "The Square-Root Law, Uncertainty, and International Reserves."

[5]See Ernst Baltensperger, "The Precautionary Demand for Reserves," *The American Economic Review* (March 1974), pp. 205-10; and J. H. G. Olivera, "The Square-Root Law of Precautionary Reserves," *Journal of Political Economy* (September/October 1971), pp. 1095-1104.

[6]A model assuming a wealth-maximizing objective of the U.S. Federal Reserve System has been shown to be a better predictor of Fed behavior than the traditional model of the Fed as an automaton reacting only to political pressures. See Mark Toma, "Inflationary Bias of the Federal Reserve System: A Bureaucratic Perspective," unpublished manuscript (California State University, Northridge, 1981). Consequently, applying a similar assumption to other central banks is not without precedent.

lowing objective function:

$$(1)\quad F(x_1,\ldots,x_n) = \prod_{k=1}^{n} (x_k - \gamma_k)^{\beta_k},$$

where x_k = asset k's maturity value at the end of the time period,

γ_k = the committed or required value of asset k,

$x_k - \gamma_k$ = the uncommitted or precautionary value of asset k,

β_k = asset k's share of the uncommitted portfolio,

and $\sum_{k=1}^{n} \beta_k = 1,$

which the central bank maximizes subject to the following balance sheet accounting constraint:

$$(2)\quad TA = \sum_{k=1}^{n} v_k x_k,$$

where $v_k = \dfrac{1}{1 + r_k}$,

r_k = the yield on asset k within the period,

TA = the present value of the assets in the portfolio.[7]

The resulting system of asset-demand equations is as follows:[8]

$$(3)\quad x_k = \gamma_k + \frac{\beta_k}{v_k}\left(TA - \sum_{j=1}^{n} \gamma_j v_j\right)$$

$$k = 1, \ldots, n$$

It is clear from equation 3 that a central bank's demand for each asset in its portfolio has two primary components. The first is the required or committed portion (γ_k), which is determined regardless of yields. The second, or precautionary, component is the *remainder* of its balance sheet ($TA - \sum_{j=1}^{n} \gamma_j v_j$), which the bank allocates to the various assets (in proportions denoted by β_k) according to relative yields in a manner that maximizes its objective function.[9]

---

[7]All assets are assumed to mature in one period, but longer-lived assets could be included without a substantive change in the analysis. Also, since the issue investigated here is a central bank's allocation of a *given* portfolio among various assets, the determination of the size of the portfolio in any time period (TA_t) is not considered. For some insights into this question, see Toma, "Inflationary Bias of the Federal Reserve System."

[8]More formally, the system of demand equations represented by equation 3 is derived by setting up the Lagrangean function and maximizing it with respect to each asset as follows:

$$(1')\quad L = \prod_{k=1}^{n} (x_k - \gamma_k)^{\beta_k} + \lambda\,(TA - \sum_{k=1}^{n} v_k x_k)$$

$$(2')\quad \frac{\partial L}{\partial x_j} = (x_1 - \gamma_1)^{\beta_1} (x_2 - \gamma_2)^{\beta_2} \ldots (x_{j-1} - \gamma_{j-1})^{\beta_{j-1}} (x_{j+1} - \gamma_{j+1})^{\beta_{j+1}} \ldots (x_n - \gamma_n)^{\beta_n} \beta_j (x_j - \gamma_j)^{\beta_j - 1} - \lambda v_j = 0$$

$$= \beta_j (x_j - \gamma_j)^{-1} F - \lambda v_j = 0 \text{ or}$$

$$(3')\quad \frac{F}{\lambda} = \frac{v_j}{\beta_j}(x_j - \gamma_j)$$

Solving (3′) for β_j yields:

$$(4')\quad \beta_j = \frac{\lambda v_j (x_j - \gamma_j)}{F}.$$

Since $\sum_k \beta_k = 1$,

$$(5')\quad \sum_j \beta_j = \frac{\lambda}{F} \sum_j v_j (x_j - \gamma_j) = 1$$

$$= \frac{\lambda}{F} (\sum_j v_j x_j - \sum_j v_j \gamma_j)$$

$$= \frac{\lambda}{F} (TA - \sum_j v_j \gamma_j) \text{ from (2) in the text or}$$

$$(6')\quad \frac{F}{\lambda} = TA - \sum_j v_j \gamma_j = \frac{v_j}{\beta_j}(x_j - \gamma_j) \text{ from (3′).}$$

Solving (6′) for x_j yields:

$$(7')\quad x_j = \gamma_j + \frac{\beta_j}{v_j} (TA - \sum_j v_j \gamma_j)$$

which is the system represented by (3) in the text. It can be shown that the own-price elasticity of demand for asset j is

$$(8')\quad \xi_{xv} = -1 + \frac{\gamma_j(1-\beta_j)}{x_j},$$

and that the Allen partial elasticity of substitution between assets i and j is

$$(9')\quad \delta_{ij} = \left(\frac{x_i - \gamma_i}{x_i}\right)\left(\frac{x_j - \gamma_j}{x_j}\right) \sum_{k=1}^{n} \beta_k \left(\frac{x_k}{x_k - \gamma_k}\right).$$

For $(x_k - \gamma_k) > 0$, all assets are Hicksian substitutes.

[9]The value of γ_k is determined by those variables that influence each country's monetary policy decisions (e.g., economic activity, unemployment, inflation). Certainly, interest rates may be included in this group of determinants. However, since γ_k is estimated, the hypothesized interest insensitivity of a portion of a central bank's portfolio can be easily tested. Specifically, if $\hat{\gamma}_k$ is statistically significant, the hypothesis that a central bank holds a portion of its portfolio for reasons other than relative yields cannot be rejected. Also, the hypothesis that *any* part of the portfolio is sensitive to changes in interest rates can be tested by testing the statistical significance of $\hat{\beta}_k$.

ESTIMATION OF THE MODELS

The Intervention Model

The functional form of central bank demand for foreign reserves for the purpose of exchange market intervention is a familiar one:[10]

$$(4)\quad \ln R_{it} = a_0 + a_1 \ln M_{it} + a_2 \ln m_{it} + a_3 \ln \sigma_{it} + a_4 \ln r_{it} + u_{it},$$

where R_{it} = the sum of country i's holdings of gold, convertible foreign exchange, SDRs and reserve position in the IMF at the end of time period t,

M_{it} = imports of i during t,

m_{it} = i's average propensity to import during t (M_{it}/GDP_{it}),

σ_{it} = the trend-adjusted variance of i's stock of foreign reserves in t,

r_{it} = i's opportunity cost of holding foreign reserves during t,

u_{it} = error term.

(All variables denominated in domestic currency units are converted into U.S. dollars using the end-of-period exchange rate.)

The use of imports as a scale variable and the average propensity to import as an indicator of openness have been discussed above. The trend-adjusted variance of country i's stock of foreign reserves is a proxy for the variability of international receipts and expenditures. It is calculated using a method similar to Frenkel's.[11]

The measure of opportunity cost employed is the ratio of the discount rate in each country to the three-month Eurodollar deposit rate. For a given portfolio of assets, the discount rate represents a measure of the foregone earnings of central banks as a result of holding assets in the form of foreign reserves; the three-month Eurodollar deposit rate is a measure of the income earned from invested foreign reserves. The rationale for this is that central banks hold most of their foreign reserves in the form of U.S. dollars. Instead of holding idle balances of dollars, central banks typically invest their reserves in some short-term asset in order to maintain a relatively high degree of liquidity; hence, the ratio (or log difference) measures the net foregone yield. Consequently, an appropriate yield on invested foreign reserves is a short-term interest rate on dollar-denominated assets.[12]

The sample employed consists of seven countries for the time period I/1964 to IV/1979.[13] The countries included are Denmark, France, West Germany, Japan, the Netherlands, Norway and Sweden. The United States is not included because it is considered to be the primary supplier of foreign reserves. The data set consists of a pooling of cross-section and time-series observations.

The possibilities that country-specific variation may be present and that a lagged adjustment process may exist are provided for in the following assumed autoregressive error structure:

$$(5)\quad u_{it} = \rho_i u_{it-1} + \epsilon_{it},$$

where ρ_i = autocorrelation parameter for country i,

ϵ_{it} = white noise random error.

Including a separate autocorrelation parameter for each country captures the country-specific variation

[10]See, for example, Frenkel, "International Reserves"; and Heller and Khan, "The Demand for International Reserves Under Fixed and Floating Exchange Rates."

[11]Frenkel, "International Reserves," p. 136. Our measure of variability is actually Frenkel's divided by the number of degrees of freedom (14 in this case); i.e.,

$$\sigma_{it} = \sum_{m=t-15}^{t-1} (R_{im} - R_{im-1} - \eta_{im})^2 / 14,$$

where η_{im} is the slope of a linear time-trend equation estimated over the period t-15 to t-1.

[12]The discount rate is employed because, even though it is not market-determined, its movement closely parallels market rates in the countries in the sample. Also, since most of the central banks studied use interest rates as a mechanism of monetary control, the discount rate reflects conditions in the respective credit markets. Government securities markets are not sufficiently developed in all of the countries to be able to use an interest rate from that market. The Eurodollar deposit rate is used as the yield on foreign reserve stocks even though other currencies are held as foreign reserves and even though some central banks have refrained generally from investing in the Eurodollar market directly. The justifications for this are: (a) the U.S. dollar is still the major reserve currency, comprising 66 to 75 percent of the foreign reserves held by central banks, (b) some central banks do invest directly in the Eurodollar market while others invest indirectly using the Bank for International Settlements as an intermediary and (c) the major alternative to the Eurodollar market is the market for U.S. Treasury bills. However, since the three-month Eurodollar rate and the three-month Treasury bill rate move very closely together, they yield virtually identical results when employed individually in the estimation of both the intervention and the asset-choice models. Finally, the ratio has been criticized as simply a proxy for the forward discount or premium on the currencies included. However, when the *covered* ratio is substituted for the uncovered one, no significant qualitative changes occur.

[13]The sample period extends to IV/1980 for Japan, West Germany and the Netherlands. Gross domestic product data were not available for the other countries in the sample for this extended period.

Table 1

Estimation of Intervention Model

	I/1964-II/1973		III/1973-IV/1979[1]	
Parameter	Estimate	Standard error	Estimate	Standard error
a_0	−.525	.420	1.102*	.503
a_1	.883*	.066	.644*	.069
a_2	−.614*	.072	−.289*	.070
a_3	.064*	.023	.113*	.038
a_4	−.073	.060	−.217*	.055

Country	ρ	ρ
Denmark	.94	.89
France	.94	.78
Germany	.83	.93
Japan	.92	.93
Netherlands	.92	.54
Norway	.90	.88
Sweden	.96	.74

N	= 259	N	= 187
RMSE	= .118	RMSE	= .141
R^2 between actual values and predicted values	= .96	R^2 between actual values and predicted values	= .87

[1]The sample period extends to IV/1980 for Japan, West Germany and the Netherlands.

*Significantly different from zero at the 5 percent level.

and also provides a means of introducing dynamic behavior into the model.[14]

Finally, the date of the switch from fixed to floating exchange rates must be identified. Since the data are pooled, it is extremely difficult to identify the break point as occurring at *a specific* point in time. It is likely the switch occurred over different intervals for each country analyzed.[15] Experimentation with various break points around the March 1973 collapse of the Smithsonian Agreement yielded no single quarter as the most likely break point for all of the countries in the sample. Consequently, the break is simply assumed to coincide with the actual failure of the Smithsonian Agreement, that is, between the second and third quarters of 1973.[16]

The results obtained from estimating the solution of equations 4 and 5 over the two time periods indicated above are reported in table 1. Several differences in the estimated relationships for the two periods are apparent. First, the import elasticity (a_1) in the fixed exchange rate period is significantly larger than that in the floating rate period. In fact, the import elasticity in the fixed period is not statistically different from one, which indicates that central bank holdings of foreign reserves do not exhibit economies of scale during that period. Second, the magnitude of the response to changes in variability (a_3) is

[14]For further explanation, see John F. O. Bilson and Jacob A. Frenkel, "Dynamic Adjustment and the Demand for International Reserves," NBER Working Paper No. 407 (November 1979), pp. 1-4; and Heller and Khan, "The Demand for International Reserves Under Fixed and Floating Exchange Rates," p. 631. As pointed out by Heller and Khan, when equation 5 is substituted into equation 4, the result is observationally equivalent to an adaptive-expectations or an error-learning process.

[15]This is supported by Frenkel, "International Reserves," pp. 122-25; and Saidi, "The Square-Root Law, Uncertainty and International Reserves," pp. 280-83.

[16]This choice is generally supported by Frenkel, "International Reserves," pp. 124-25, and by Heller and Khan, "The Demand for International Reserves Under Fixed and Floating Exchange Rates," pp. 637-39. The selection of the break point is also constrained by the necessity to choose the same break point for each model so that the performance can be compared over identical sample periods. Also, for each model, the hypothesis that the estimated parameters before this point are equal to those after this point is rejected at the 5 percent confidence level.

larger under floating than under fixed rates. This is somewhat paradoxical since one might expect that the increased exchange rate flexibility during the floating rate period would serve as a buffer and, consequently, reduce central banks' response to changes in variability.[17]

Third, the sensitivity of central banks' reserve holdings to interest rate changes under fixed rates (a_4) is insignificant, a result similar to that of other studies.[18] Alternatively, under floating rates, central banks are found to respond in a significant and conceptually consistent manner to changes in interest rates. When compared with those of previous studies, these results suggest that an interest rate differential is a better measure of the opportunity cost of holding reserves. Finally, a comparison of the intercepts (a_0) suggests that central banks are holding larger stocks of foreign reserves, on average, in the floating rate period than they did in the fixed rate period, indicating that they have actually added to their stocks during the floating period.

The Asset-Choice Model

To estimate the system of asset-demand equations represented by equation 3, it was assumed that normally distributed random errors enter additively with zero mean and constant variance. As a result of introducing a random component in this manner, the sum of the error terms across all equations in the system must equal zero if the system is to be consistent.[19] This restriction on the error structure, by introducing linear dependence across equations, has at least two important implications for estimation. First, single-equation estimating techniques are inappropriate. Efficient estimation requires the use of a system technique. Second, the covariance matrix of the entire system is singular. Because of this, a full-information technique cannot be employed on the entire system of n asset-demand equations simultaneously because the inversion of this covariance matrix is required during the estimation process. Consequently, only n-1 equations can be estimated simultaneously.[20]

The countries and time periods employed here are identical to those used in estimating the intervention model. The assets of the central banks of these countries are aggregated into three categories: foreign reserves, claims on government and claims on commercial banks. The interest rates used for these asset groups are the three-month Eurodollar deposit rate (for foreign reserves), short-term government bond yield in country i (for claims on government) and the discount rate in i (for claims on commercial banks). The three-month Eurodollar rate is used here for the same reason it was used in the estimation of the intervention model. Also, a dynamic specification is employed to capture lagged adjustment of the committed parameters (γ_k) by allowing them to vary over time. This dynamic feature is introduced into the system by assuming that the committed level of each asset is a function of the total holding of that asset during the previous time period as follows:

$$(6)\quad \gamma_{kt} = \theta_k x_{kt-1},$$

with $0 \leq \theta_k \leq 1$ for all k. The parameter θ_k reflects a proportional relationship between the committed level of asset k in the current period to the total holding of that asset in the preceding period. Finally, the date of the switch from fixed to floating exchange rates is the same as in estimating the intervention model.

Substituting equation 6 into equation 3 and recognizing that n=3 in this case, the resulting system of asset-demand equations is as follows:

$$(7.1)\quad x_{1it} = \theta_1 x_{1it-1} + \frac{\beta_1}{v_{1it}}\left(TA_{it} - \sum_{j=1}^{3} \theta_j x_{jit-1} v_{jit}\right) + u_{1it}$$

$$(7.2)\quad x_{2it} = \theta_2 x_{2it-1} + \frac{\beta_2}{v_{2it}}\left(TA_{it} - \sum_{j=1}^{3} \theta_j x_{jit-1} v_{jit}\right) + u_{2it}$$

$$(7.3)\quad x_{3it} = \theta_3 x_{3it-1} + \frac{\beta_3}{v_{3it}}\left(TA_{it} - \sum_{j=1}^{3} \theta_j x_{jit-1} v_{jit}\right) + u_{3it}$$

[17]Frenkel, "International Reserves," p. 120, also obtained this result; however, Saidi, "The Square-Root Law, Uncertainty and International Reserves," p. 285, found smaller responses to changes in variability in the floating-rate period.

[18]See footnote 4.

[19]For the system of asset-demand equations represented by equation 3 to be consistent, the value of the *estimated* portfolio must equal the value of the *actual* portfolio. This condition implies that the error terms across all n asset-demand equations must sum to zero. That is, the error terms across equations are linearly dependent and thus, by definition, correlated. It could also be argued that, for this analysis, the demands for assets are correlated regardless of the consistency condition. In particular, if the impact of foreign exchange market intervention upon the domestic money supply is sterilized (e.g., through an offsetting sale or purchase of government securities), then foreign exchange holdings and government security holdings are necessarily negatively correlated.

[20]Robert A. Pollak and Terence J. Wales, "Estimation of the Linear Expenditure System," *Econometrica* (October 1969), pp. 611-28. They prove that if a full-information, maximum-likelihood estimation procedure is employed, the estimated parameters are invariant to whichever n-1 equations are included.

Table 2

Estimation of Asset-Choice Model

	I/1964-II/1973		III/1973-IV/1979[1]	
Parameter[2]	Estimate	Standard error	Estimate	Standard error
θ_1	.987*	.022	.983*	.011
θ_2	.893*	.032	1.001*	.018
θ_3	1.005*	.017	.897*	.049
β_1	.536*	.048	.261*	.038
β_2	.326*	.035	.213*	.047
β_3	.138*	.041	.526*	.058
	RMSE of equation 7.1 = .697		RMSE of equation 7.1 = 1.404	
	R^2 between actual and predicted values for equation 7.1 = .98		R^2 between actual and predicted values for equation 7.1 = .99	

[1]The sample period extends to IV/1980 for Japan, West Germany and the Netherlands.

[2]The subscripts 1, 2 and 3 refer to the three asset categories, foreign reserves, claims on government and claims on commercial banks, respectively.

*Significantly different from zero at the 5 percent level.

where x_{jit} = the value of country i's holding of asset j at the end of time period t,

$v_{jit} = \frac{1}{1 + r_{jit}}$,

r_{jit} = the yield on asset j in country i from beginning to end of time period t,

TA_{it} = the value of i's portfolio at beginning of period t,

u_{jit} = error term.

Table 2 presents the results of estimating the above system omitting equation 7.3.[21] A full-information, maximum-likelihood technique is used to obtain efficient estimates.

All parameter estimates are statistically significant and within conceptually acceptable ranges of values. As before, differences between time periods, but also across assets, are readily apparent. In particular, the estimated committed parameter for foreign reserves ($\hat{\theta}_1$) is relatively constant across time periods, indicating that central banks have not altered the committed portion of their foreign reserves in the move from fixed to floating exchange rates. On the other hand, the estimated committed parameters for claims on government ($\hat{\theta}_2$) and for claims on commercial banks ($\hat{\theta}_3$) have changed significantly with the change in regimes.[22] Furthermore, the percentage of their discretionary portfolio that central banks held in the form of foreign reserves ($\hat{\beta}_1$) fell significantly from the fixed to the floating period. The sensitivity of the demand for foreign reserves to changes in interest rates (as measured by the absolute value of the price elasticity of demand) also fell from .563 in the fixed rate period to .289 in the floating rate period. Nonetheless, the fact that this percentage is statistically significant in both periods indicates that reserve holdings are at least partially sensitive to changes in interest rates.

Taken together, the changes in $\hat{\theta}_1$ and $\hat{\beta}_1$ over the two periods shed some light on why Heller and Khan consistently *overpredict* central bank demand for foreign reserves during the floating period.[23] In their model, central banks hold foreign reserves solely to intervene in foreign exchange markets. Alternatively, in the asset-choice model, intervention is simply one of several motives (where the committed parameter measures the demand for reserves for

[21]Except for β_3 and its variance, all parameters and their variances are estimated directly. Since $\sum_k \beta_k = 1$, $\hat{\beta}_3 = 1 - \hat{\beta}_1 - \hat{\beta}_2$ and $Var(\hat{\beta}_3) = Var(\hat{\beta}_1) + Var(\hat{\beta}_2) + 2\,Cov(\hat{\beta}_1, \hat{\beta}_2)$. The same results as those reported were obtained when either equation 7.1 or 7.2 (instead of 7.3) was deleted.

[22]Even though $\hat{\theta}_3$ in the fixed period and $\hat{\theta}_2$ in the floating period are greater than 1 (the conceptual limit of each), neither is significantly greater than 1 in a statistical sense.

[23]Heller and Khan, "The Demand for International Reserves Under Fixed and Floating Exchange Rates," pp. 639-43.

Table 3

Partial Elasticities of Substitution

Assets	I/1964-II/1973	III/1973-IV/1979[1]
Foreign reserves and claims on government	.116	.028
Foreign reserves and claims on commercial banks	.031	.076
Claims on government and claims on commercial banks	.065	.094

[1]The sample period extends to IV/1980 for Japan, West Germany and the Netherlands.

purposes of intervention).[24] Even though this relationship appears to be relatively stable across time (in the asset-choice model), overlooking the significant decline in the percentage of the precautionary portfolio held in the form of foreign reserves by basing predictions on a intervention model should lead to an overprediction of reserve demand, ceteris paribus.

One final question remains to be answered: Are the assets in central banks' portfolios close substitutes for each other? To answer this question, partial elasticities of substitution are calculated for each of the asset pairs over each time period. Since these elasticities are functions, inter alia, of the committed and uncommitted levels of each asset, the elasticities reported are evaluated using the mean holdings of the relevant assets (table 3). Given the relatively high estimated values of the committed parameters, it is not too surprising to find that none of the assets are close substitutes.

Predictive Abilities of the Two Models

The ultimate test of a structural model is how well it predicts behavior. This section compares the predictive abilities of the two models described above.

Table 4

Residual-Variance Estimates

	I/1964-II/1973	III/1973-IV/1979[1]
Intervention model		
Level	.788	4.577
Log-level	.0139	.0199
Asset-choice model (Equation 7.1)		
Level	.486	1.970
Log-level	.0144	.0167

[1]The sample period extends to IV/1980 for Japan, West Germany and the Netherlands.

Two methods of comparison are employed: The first is the residual-variance criterion developed by Theil.[25] The use of the residual-variance criterion involves calculating a residual-variance estimate (error sum of squares divided by degrees of freedom) for each model and selecting the model with the smallest residual variance.[26] Since the intervention model is estimated in log-level form and the asset-choice model is not, the residual-variance estimates from the two models are not directly comparable. To make these estimates comparable, either the residuals of the estimated intervention model have to be transformed from logarithms to levels or the residuals of the estimated asset-choice model have to be transformed from levels to logarithms.[27] Table 4 presents the results of both of these transformations. Except for the logarithmic specification estimated over the fixed rate period, the asset-choice model appears to outperform the intervention model.

These results, however, must be qualified. The residual-variance method presupposes that one of the specifications is *the correct* one, a somewhat presumptuous supposition. Also, in this case the two

[24]One may infer that, since the asset-choice model does not explicitly contain explanatory variables that represent the intervention motive, it is fundamentally misspecified. However, the estimation of the asset-choice model clearly indicates that the foreign reserve demands of central banks are sensitive to yields on other assets in their portfolio. Since the intervention model ignores these explanatory variables, it is also fundamentally misspecified. Consequently, future research should be directed at combining the features of both of these models to specify correctly a central bank's demand for foreign reserves.

[25]Henri Theil, *Principles of Econometrics* (John Wiley and Sons, Inc., 1971), pp. 543-45, 553-54.

[26]The selection of the specification with the smallest residual variance is justified by the following proposition: if the correct specification has uncorrelated disturbances with zero mean and constant variance and if the explanatory variables are nonstochastic, the residual-variance estimator of the correct specification has an expectation that is never larger than that of an incorrect specification. See Theil, *Principles of Econometrics*, p. 543.

[27]This transformation is accomplished by converting the actual and the predicted values from the level (logarithmic) specification into logarithms (anti-logs), calculating the sum of squared deviations of the predicted value from the actual, then adjusting for degrees of freedom.

models compared are non-nested; that is, the models have separate sets of explanatory variables such that one model cannot be obtained from the other. Consequently, the conventional use of summary statistics and F-tests to discriminate among alternatives can be misleading and even inappropriate.[28]

The second method is an extension of the Cox test developed by Pesaran and Deaton.[29] This procedure for testing non-nested hypotheses is not subject to either of the above qualifications necessary for interpreting the results of the residual-variance method. In particular, Pesaran and Deaton's procedure does not employ a single maintained (null) hypothesis. (No model is considered a priori to be the correct one.) The alternative models are analyzed one at a time. One by one, each is assumed to be the correct one.) The alternative models are analyzed one at a time. One by one, each is assumed to be the correct model (null hypothesis); the alternative has been observed. The notion of absolute goodness of fit plays no role in this procedure. In fact, the possibility exists that all competing models may be rejected. This is not the case for conventional testing procedures.[30]

The test statistics calculated with the intervention model and equation 7.1 of the asset-choice model, respectively, as the null hypothesis are reported in table 5.[31] Under the null hypothesis, this test statistic is asymptotically distributed as a normal random variable with zero mean and unit variance. The results are unambiguous. When confronted with the data and the asset-choice model as an alternative, the intervention model must be rejected. Alternatively, the asset-choice model cannot be rejected. This conclusion is invariant across sample periods. While the rejection of the intervention model for the floating rate period is not unexpected, it is certainly interesting that this model is also rejected for the fixed rate period. This result confirms that the asset-choice model provides a more general explanation of central banks' demand for reserves than does the intervention model.

Table 5

Statistics for Testing Hypotheses Involving Non-Nested Models

H_0: Intervention model
H_A: Asset-choice model (Equation 7.1)

Period	Test statistic
I/1964-II/1973	−18.36*
III/1973-IV/1979[1]	−22.62*

H_0: Asset-choice model (Equation 7.1)
H_A: Intervention model

Period	Test statistic
I/1964-II/1973	−0.85
III/1973-IV/1979[1]	−0.75

[1]The sample period extends to IV/1980 for Japan, West Germany and the Netherlands.

*Statistically different from zero at the 5 percent level.

SUMMARY AND CONCLUSION

The purpose of this article has been to compare central bank behavior before and after the movement to floating exchange rates within the framework of two alternative models of a central bank's demand for foreign reserves. In the first model,

[28]See M. H. Pesaran, "On the General Problem of Model Selection," *The Review of Economic Studies* (April 1974), pp. 153-71.

[29]D. R. Cox, "Tests of Separate Families of Hypotheses," in *Proceedings of the Fourth Berkeley Symposium on Mathematical Statistics and Probability* (University of California Press, 1961), pp. 105-123; M. H. Pesaran and A. S. Deaton, "Testing Non-Nested Nonlinear Regression Models," *Econometrica* (May 1978), pp. 677-94.

[30]A necessary condition for the use of this test is that both models explain the same dependent variable. In this case, the first equation of the asset-choice model explains the quantity of reserves demanded while the intervention model explains the *logarithm* of the quantity of reserves demanded. Consequently, to perform the Cox test, the anti-log of the intervention model (i.e., a non-linear, Cobb-Douglas-type function) is estimated using a maximum-likelihood procedure. The resulting predicted values and estimated parameters are essentially identical to those obtained from a least-squares estimation of the log-linear functional form.

[31]The test statistic (C) is defined as:

$$C = \frac{T_0}{\sqrt{\mathrm{Var}(T_0)}},$$

$$\text{where } T_0 = \frac{T}{2} \ln \frac{\hat{\sigma}_A^2}{\hat{\sigma}_0^2 + \frac{1}{T}[f(\hat{\phi}_0) - g(\hat{\phi}_{A0})]'[f(\hat{\phi}_0 - g(\hat{\phi}_{A0})]},$$

T	=	sample size,
$\hat{\sigma}_0^2$	=	estimated variance of the model under H_0,
$\hat{\sigma}_A^2$	=	estimated variance of the model under H_A,
$f(\hat{\phi}_0) - g(\hat{\phi}_{A0})$	=	the residuals from an auxiliary estimation of the model under H_A using the predicted values from the model under H_0 as the dependent variable,
$\mathrm{Var}(T_0)$	=	the variance of T_0 as defined in Pesaran and Deaton, "Testing Non-Nested Nonlinear Regression Models," p. 687.

foreign reserves are treated as a special type of asset, one demanded solely to enable a central bank to intervene in foreign currency markets. The second model considers foreign reserves to be the same as — and also to be held for the same reasons as — any other asset within a central bank's portfolio.

The estimation of the asset-choice model as an alternative to the intervention model yielded several interesting results. First, a central bank's demand for foreign reserves is sensitive to relative changes in the yields of the assets in the portfolio. Second, central banks consider foreign reserves as substitutes to other assets in their portfolio. Third, the decrease in the percentage of the uncommitted portfolio composed of foreign reserves is identified as a possible reason for the usual overprediction of reserve demand by the intervention model in the floating rate period. Finally, and most importantly, the asset-choice model consistently outperforms the intervention model.

Since the testing procedure employed could lead to the rejection of both models, the fact that the asset-choice model cannot be rejected in either sample period is an extremely robust result. The implication is simply that, regardless of exchange rate regime, central banks hold foreign reserves for a wide variety of purposes — not just for intervention in foreign exchange markets. Consequently, the investigation of whether or not central banks' general behavior has changed with the movement to a system of floating exchange rates within the framework of the intervention model appears to be misdirected. Investigation should focus on the arguments, instead of the parameters, within the demand function.