The Impact of Financial Deregulation on the Relationship Between Stock Prices and Monetary Policy

Unro Lee
University of the Pacific

This paper examines whether aggregate stock prices fully incorporated all available information on the money supply in the United States from January 1978 to September 1990, a period that witnessed the deregulation of deposit rates and several changes in the Federal Reserve's short-run operating procedure. A Granger causality test is performed in the context of a bivariate error correction model (ECM) on the monthly data. I find that stock prices did not fully digest and reflect changes in money supply during this period. This phenomenon can be attributed partly to the existence of a relatively volatile relationship between money supply and overall economic activity engendered by financial deregulation.

INTRODUCTION

Since 1978 the U.S. economy has witnessed the gradual elimination of regulations on the deposit yields offered by financial institutions. During the late 1970s when the interest rates were escalating to high levels, Regulation Q (a Federal Reserve regulation that imposed a ceiling on deposit yields) triggered a large loss of deposits for mortgage-issuing institutions such as savings and loans and mutual savings banks.

In order to stimulate the supply of mortgage credit to prospective home buyers, the regulatory authorities permitted these financial institutions to offer market rates on small denomination time deposits beginning in July 1978. The ceiling subsequently was lifted for other time deposits by the end of 1982.

The continuing rise in interest rates also led Congress to introduce the Depository Institutions Deregulation and Monetary Control Act (DIDMCA) of 1980. DIDMCA authorized all depository institutions to offer interest-bearing checkable deposits, such as NOW and ATS accounts, beginning in January 1981 and Super-NOW accounts in January 1983. Furthermore, DIDMCA mandated the elimination of Regulation Q by 1986.[1]

Prior to deregulation, M1 was regarded primarily as a transactions monetary aggregate, whereas M2 was regarded as a savings aggregate. The introduction of NOW, ATS, and Super-NOW accounts, which contain features of both the transactions and savings funds, and the elimination of the ceiling on the deposit yields for time deposits, however, have blurred the distinction between M1 and M2. The relationships between both M1 and M2 and various measures of economic performance became more volatile in the 1980s.[2]

---

[1] DIDMCA also eliminated usury ceilings on mortgage loans and imposed uniform reserve requirements on all depository institutions. For an excellent description of various financial regulatory changes since the late 1970s, please refer to Gilbert (1986).

[2] Marlow (1991) finds that the increased volatility of the velocity of money supply (both M1 and M2) during 1983-1989 can be explained partly by more variable money growth experienced during this period.

0747-5535/94/1400-0037 $01.00

The purpose of this study is to examine whether the aggregate U.S. stock prices, long considered a leading indicator of the economy, effectively have incorporated all available information on the money supply (both M1 and M2) since the deregulation of deposit yields. The time period investigated is from January 1978 to September 1990.

A Granger causality test procedure is employed in the context of a bivariate error correction model (ECM) to examine the money supply-stock price relationship. I find that the aggregate U.S. stock prices did not fully digest all available information on both M1 and M2 from January 1978 to September 1990. Moreover, the market inefficiency primarily is due to the subperiod October 1982 to September 1990.

COINTEGRATION AND CAUSALITY TESTS

Empirical studies such as Hall (1978) and Nelson and Plosser (1982) document nonstationary patterns of many economic variables.[3] A nonstationary process may require differencing to induce stationarity. A variable that needs to be differenced d times to reach stationarity is referred to as an *I(d) or integrated-of-order-d process*. I(d) also is said to contain d unit roots.

If two nonstationary variables that are integrated of the same order maintain a long-run equilibrium relationship (i.e., they move together over time), then the fundamental economic forces that induce nonstationary behavior of one variable also have a similar impact on the other variable. Therefore, the magnitude of the deviation from the equilibrium path of these two variables as a result of a random shock should decline over time or be a stationary process.

The long-run equilibrium relationship of two variables can be tested in a rigorous manner by the cointegration technique that was introduced by Granger (1983). Two variables, say X and Y, are considered to be cointegrated if the variables are nonstationary and integrated of same order and if the error term, ξ_t, in the following equation is stationary:[4]

$$(1)\ Y_t = \alpha + \beta X_t + \xi_t.$$

The test of the null hypothesis of first order nonstationarity can be performed by using a statistical procedure proposed by Dickey and Fuller (1979, 1981). This test can be performed by first estimating the following regression equation for each variable:

$$(2)\ \Delta Y_t = \alpha + \gamma t + \rho Y_{t-1} + \sum_{i=1}^{k} \beta_i \Delta Y_{t-i} + \xi_t$$

where Y is the variable under consideration. The order of k should be set large enough to ensure the estimated residual series, ξ_t, becomes white noise. The null hypothesis that Y

---

[3] Hall (1978) shows that the aggregate consumption pattern follows a random walk process, whereas Nelson and Plosser (1982) document that 14 other major macroeconomic variables exhibit nonstationary behavior over time.

[4] Granger (1983) formally defines the components of the vector X_t as being cointegrated of order d and b only if all components of X_t are integrated of order d and if there exists a vector α ($\alpha \neq 0$) such that $Z_t = \alpha' X_t$ is integrated of order d - b, b > 0.

contains a unit root (i.e., Y is I(1)) is rejected if the estimated coefficient of the lagged variable, ρ, is significantly less than zero. The estimated t-statistic for ρ is referred to as the *augmented Dickey-Fuller (ADF) statistic.*

One potential problem that may arise in performing Dickey-Fuller tests is that the substantive results can be overly sensitive to the presence of drift (α) and time trend (γ) terms in the regression equation (2). Therefore, two alternative specifications of the ADF statistic, t_u and t_t, can be used for the unit root test, where the t_u statistic is relevant for the regression equation with just the intercept (drift) and the t_t statistic is relevant for the equation with both the intercept and time trend. Because both test statistics for ρ do not necessarily follow the usual Student t-distribution under the null hypothesis of a unit root, the tabulated values provided in Fuller (1976) should be used.

The null hypothesis that Y_t is I(d), where d > 1, can be tested in the same manner after differencing Y by d times.

After the variables X and Y are determined as being integrated of same order, then the hypothesis of cointegration between these variables can be tested after a regression equation of the following form is estimated:

$$(3)\ Y_t = \alpha + \beta X_t + z_t$$

where z_t is the error term. Ordinary least squares (OLS) can be employed to estimate equation (3) to yield a reliable estimate of β.

Suppose that both X and Y are I(1), as are many nonstationary macroeconomic aggregates. If X and Y are not cointegrated, then the residual series z in the equation (3) also will be I(1). If X and Y are cointegrated, however, then z will be I(0).

The process of ascertaining whether two variables are cointegrated is methodologically equivalent to testing for the existence of a unit root for the residual series z. Accordingly, the ADF test statistic (t_u) can be examined after estimating the equation (2) with Y replaced by the estimated residuals, z, of equation (3). Because these residuals are obtained from a regression equation, the critical values in Engle and Yoo (1987) are appropriate in this case.

If two variables, X and Y, are found to be cointegrated, then the causal relationship between these two variables can be examined by estimating the following bivariate ECM (Sims, Stock, and Watson, 1988):[5]

$$(4)\ \Delta X_t = \alpha_1 + \gamma_1(X_{t-1} - \beta_1 Y_{t-1}) + \sum_{i=1}^{k1} \mu_{1i}\Delta X_{t-i} + \sum_{i=1}^{k2} \theta_{1i}\Delta Y_{t-i} + \varepsilon_t$$

$$(5)\ \Delta Y_t = \alpha_2 + \gamma_2(Y_{t-1} - \beta_2 X_{t-1}) + \sum_{i=1}^{c1} \mu_{2i}\Delta Y_{t-i} + \sum_{i=1}^{c2} \theta_{2i}\Delta X_{t-i} + \varepsilon_t.$$

Bivariate ECM relates the changes in one variable to the past equilibrium error and to the past changes in both variables. That is, ECM simply assumes that a portion of dis-

[5] If X and Y are found to be cointegrated, then β_2 in the equation (5) should be equal to $1/\beta_1$.

equilibrium at a given period will be corrected in subsequent periods. According to the Granger representation theorem (Engle and Granger, 1987), the concept of cointegration and error correction are equivalent. That is, if two variables are found to be cointegrated, then their relationship can be represented by ECM.

The null hypothesis of no causality from Y to X can be formulated as:

$$H_0: \gamma_1 = 0 \text{ and } \theta_{1i} = 0, \text{ for all i.}$$

Accordingly, the null hypothesis of no causality from X to Y can be expressed as:

$$H_0: \gamma_2 = 0 \text{ and } \theta_{2i} = 0, \text{ for all i.}$$

These null hypotheses can be tested with the usual F-statistic (Sims, Stock, and Watson, 1988).

The cointegration procedure described above can be used to examine the long-run relationship between money supply and stock prices. The causal relationship between money and stock prices is tested in the context of a bivariate ECM

DATA AND EMPIRICAL RESULTS

The empirical analysis is conducted on monthly data from January 1978 to September 1990. The measure of the aggregate stock prices is the Standard & Poor's Composite Stock Index for 500 stocks (SP500) at the end of the month.

The data on M1 and M2 are obtained from various issues of the *Federal Reserve Bulletin*.[6] These monetary aggregate data are unrevised, seasonally adjusted, and end-of-the-month figures.

The Dickey-Fuller test determines whether SP500, M1, and M2 are nonstationary processes and integrated of the same order. Two alternative specifications of the ADF test statistic, t_u and t_t, on ρ in the equation (2) are examined. Because the results of the causality test can be sensitive to the choice of the lag length k (see Thornton and Batten, 1985), I use the Akaike (1974) information criterion (AIC) to determine the optimal lag-structure specification of the equation (2). AIC searches for the lag length k that would minimize

$$(6)\ \text{AIC}(k) = T \ln\left(\frac{\text{SSR}(k)}{T}\right) + 2p,$$

where:

T = The number of observations;
SSR(k) = The sum of squared residuals of the regression with k as the lag length; and
p = The number of regressors which are equal to k+1.

[6] The data for both measures of money supply are from the monthly issues of the *Federal Reserve Bulletin* published between April 1978 and December 1990.

The appropriate order of the lag length for each variable can be obtained by computing AIC(k) over the range of values for k from 1 to 12. The k^*th lag that minimizes AIC is provided for all three variables in Table 1.

Part A of Table 1 indicates that the null hypothesis of I(1) cannot be rejected for the three variables at both the 1 percent and 5 percent significance levels. Furthermore, the results are not sensitive to the choice of the ADF statistic.

SP500, M1, and M2 then are tested for second order integration. The results are presented in Part B of Table 1. The t_u and t_t statistics indicate that the null hypothesis of I(2) can be rejected for all three variables at the 1 percent significance level. Therefore,

Table 1—Unit Root Tests for SP500, M1, and M2: January 1978 to September 1990

Panel A H_0: I(1)	Y	k*	$t_u(k^*)$
(i) $\Delta Y_t = \alpha + \rho Y_{t-1} + \sum_{i=1}^{k} \beta_i \Delta Y_{t-i} + \xi_t$			
	SP500	2	1.5610
	M1	12	3.9566
	M2	12	2.4324
(ii) $\Delta Y_t = \alpha + \gamma t + \rho Y_{t-1} + \sum_{i=1}^{k} \beta_i \Delta Y_{t-i} + \xi_t$			
	SP500	2	-2.6653
	M1	3	-1.9775
	M2	3	-2.6394
Panel B H_0: I(2)	**Y**	**k***	**$t_u(k^*)$**
(i) $\Delta^2 Y_t = \alpha + \rho \Delta Y_{t-1} + \sum_{i=1}^{k} \beta_i \Delta^2 Y_{t-i} + \xi_t$			
	SP500	1	-8.1731**
	M1	11	-4.7855**
	M2	11	-5.5259**
(ii) $\Delta^2 Y_t = \alpha + \gamma t + \rho \Delta Y_{t-1} + \sum_{i=1}^{k} \beta_i \Delta^2 Y_{t-i} + \xi_t$			
	SP500	1	-8.1519**
	M1	2	-4.8038**
	M2	2	-5.5609**

Note: * and ** denote the significance at the 5 percent and 1 percent level, respectively. The critical values for t_u statistic are -2.57 (10 percent), -2.88 (5 percent), and -3.46 (1 percent) for n = 250. For t_t statistic, the critical values for the same sample size are -3.13 (10 percent), -3.43 (5 percent), and -3.99 (1 percent) (Fuller, 1976, p. 373)

SP500 represents the Standard & Poor's Composite Stock Index for 500 stocks at the end of the month (source: *Survey of Current Business*)

The data on M1 and M2 are obtained from various issues of the *Federal Reserve Bulletin*

SP500, M1, and M2 are found to have only one unit root. That is, these variables can be characterized appropriately as first order, nonstationary processes.

I use the Engle-Granger test to determine cointegration between stock prices and the money supply. The results are summarized in Part A of Table 2. The ADF statistic is obtained after the lag length is chosen by invoking the AIC procedure.

The null hypothesis of no cointegration between either M1 or M2 and SP500 is rejected at the 5 percent significance level. The results are robust with respect to the choice of the dependent variable for the equation (3).

Table 2—Cointegration Tests for SP500, M1, and M2: January 1978 to September 1990

Panel A Engle-Granger test with no time trend: $Y_t = \alpha + \beta X_t + z_t$, H_0: *no cointegration between X and Y* (i.e., $z \sim I(1)$)

Y	X	k*	ADF(k*)
SP500	M1	2	-3.3265*
M1	SP500	2	-3.3124*
SP500	M2	2	-3.3362*
M2	SP500	2	-3.5208*

Panel B Engle-Granger test with a time trend: $Y_t = a + \gamma t + \beta X_t + z_t$, H_0: no cointegration between X and Y (i.e., $z \sim I(1)$)

Y	X	k*	ADF(k*)
SP500	M1	2	-3.9371*
M1	SP500	2	-3.9344*
SP500	M2	2	-3.9542*
M2	SP500	2	-3.8362*

Panel C Johansen's multivariate cointegration test

	$-2\ln Q_r$ statistic for H_0: $r = 0$
i. SP500 and M1	29.07**
ii. SP500 and M2	50.93**

Critical values for test statistic	ADF (part A)	ADF (part B)	Johansen's test
1% level	-3.78	-4.34	24.99
5% level	-3.25	-3.82	20.17
10% level	-2.98	-3.51	17.96

Note: The critical values for ADF in Part A are from Engle and Yoo (1987, Table 3) for n = 200. The critical values for ADF in Part B are from Engle and Yoo (1989) for n = 200, and the critical values for the Johansen's test are from Johansen and Juselius (1990, Table A3)

* Significant at the 5 percent level
** Significant at the 1 percent level

SP500 represents the Standard & Poor's Composite Stock Index for 500 stocks at the end of the month (source: *Survey of Current Business*)

The data on M1 and M2 are obtained from various issues of the *Federal Reserve Bulletin*.

Unfortunately, the validity of the results may be diminished if the true specification of a variable under investigation is a joint process of a linear trend and I(1). (See Engle and Yoo, 1989.) If a variable has a detectable trend, then it should be removed before a cointegration test can be performed. Thus, I perform cointegration tests between SP500 and both M1 and M2 with a trend term included in the cointegration equation (3) as follows:

$$(7)\ Y_t = \alpha + \gamma t + \beta X_t + z_t.$$

The critical values for testing the null hypothesis of no cointegration with the regression equation (7) are provided in Engle and Yoo (1989).[7]

As can be seen in Part B of Table 2, the null hypothesis of no cointegration between stock price and both measures of money supply is rejected at the 5 percent significance level even after the trend term is removed.

Dickey, Jansen, and Thornton (1991) also warn that the Engle-Granger cointegration test may not be robust with respect to the choice of the variable on the left side when a finite sample is used. They suggest a multivariate cointegration test such as that proposed by Johansen (1988) should circumvent this problem. Johansen's multivariate test is performed in the context of a vector autoregressive (VAR) model that explicitly allows for possible dynamic interactions among variables. Therefore, I use Johansen's test procedure to test the null hypothesis of no cointegration between stock prices and money supply.[8]

Consider a general VAR model specified as follows:

$$(8)\ \Delta X_t = \Gamma_1 \Delta X_{t-1} + \ldots + \Gamma_{k-1} \Delta X_{t-k-1} + \Gamma_k \Delta X_{t-k} + \mu + \varepsilon_t$$

where:

- X_t = An $n \times 1$ vector of the variables;
- Γ = An $n \times n$ coefficient matrix;
- μ = An $n \times 1$ constant vector; and
- ε_t = An $n \times 1$ vector of white noises with mean zero and finite variance.

The rank of the coefficient matrix Γ indicates the number of cointegrating relationships existing among the variables subsumed in the vector X (Johansen, 1988). Because vector X in this study consists of stock prices and either M1 or M2, n is set at 2. Therefore, the null hypothesis of no cointegration between the stock prices and money supply (either M1 or M2) is equivalent to the hypothesis that the rank of $\Gamma = 0$. On the other hand, if the rank of the coefficient matrix is one, then stock prices and the money supply are cointegrated.

---

[7] I would like to thank one of the referees for providing these critical values.

[8] A step-by-step instruction for Johansen's multivariate cointegration test is presented in Dickey, Jansen, and Thornton (1991, pp. 66). I would like to thank one of the referees for suggesting this test.

The Johansen test procedure starts with the estimation of the following regression equations:

(9) $\Delta X_t = \Gamma_1 \Delta X_{t-1} + \ldots + \Gamma_{k-1} \Delta X_{t-k-1} + \varepsilon_{1t}$

(10) $X_{t-k} = \Gamma_1 \Delta X_{t-1} + \ldots + \Gamma_{k-1} \Delta X_{t-k-1} + \varepsilon_{2t}$

Define the product moment matrices of the residuals, ε_{1t} and ε_{2t}, as $S_{ij} = T^{-1}\Sigma^{T}_{t=1}\, \varepsilon_{it}\varepsilon'_{jt}$ for i, j = 1, 2, where T is the sample size. Next, the squares of the canonical correlation between ε_{1t} and ε_{2t} are ranked as $\phi_1 > \phi_2 > \ldots > \phi_n$, where ϕ_i, i = 1, 2, ..., n, also are referred to as the *eigenvalues* of $S_{21}S^{-1}_{11}S_{12}$ with respect to S_{22}. Johansen derives the likelihood ratio test statistic for the null hypothesis of "at most r cointegrating relationships" as

$$(11)\ -2\ln Q_r = -T \sum_{j=r+1}^{n} \ln(1 - \phi_j).$$

where the $-2\ln Q_r$ statistic also is referred to as the *trace test statistic*. The limiting distribution of the $-2\ln Q_r$ statistic is specified as an n-r dimensional Brownian motion process. The quintiles of the distribution of the statistic are provided in Johansen and Juselius (1990, Table A3). Also, the maximum likelihood estimate of the cointegrating vector can be computed as the first eigenvector of $S_{21}S^{-1}_{11}S_{12}$ with respect to S_{22}, i.e., α^* for which $\phi_1\, S_{22}\, \alpha^* = S_{21}S^{-1}_{11}S_{12}\, \alpha^*$.

The results of the Johansen multivariate cointegration test (with n = 2) are presented in Part C of Table 2. The hypothesis of no cointegration between stock prices and either M1 or M2 (i.e., H_0: r = 0) is rejected at the 1 percent significance level.

The findings of both the Engle-Granger test and the Johansen's multivariate test indicate a long-run equilibrium relationship between stock prices and both measures of money supply from January 1978 to September 1990.

Because the cointegration relationship between stock prices and money supply has been identified statistically, the causal relationship between these variables can be examined next in the context of the following bivariate ECM:

$$(12)\ \Delta SP500_t = \alpha_1 + \gamma_1(SP500_{t-1} - \beta_1 M_{t-1}) + \sum_{i=1}^{k1} \mu_{1i}\Delta SP500_{t-i} + \sum_{i=1}^{k2} \theta_{1i}\Delta M_{t-i} + \varepsilon_t$$

$$(13)\ \Delta M_t = \alpha_2 + \gamma_2(M_{t-1} - \beta_2\, SP500_{t-1}) + \sum_{i=1}^{c1} \mu_{2i}\Delta M_{t-i} + \sum_{i=1}^{c2} \theta_{2i}\Delta SP500_{t-i} + \varepsilon_t$$

The length of the lag structures—k1, k2, c1 and c2—for equations (12) and (13) are specified by using Hsiao's (1981) two step search method over the range of the lag length from 1 through 12. The lag lengths are chosen invoking the AIC procedure. The lag lengths that minimize AIC are determined as 2, 1, 3, and 1, respectively, for k1, k2, c1, and c2, with M1 as the measure of money supply. With M2 used as a proxy for money supply, the optimal values are 2, 1, 1, and 1, respectively.

The estimated equations (12) and (13) with the two different measures of money supply are presented in Table 3. The estimates for the adjustment coefficient γ are statistically significant at the 5 percent level regardless of whether M1 or M2 is used as a measure of money. Also, as expected, the signs of the estimated adjustment coefficient are all negative. This implies that, given a fixed money supply, an increase in stock prices at time t due to a random shock will induce the stock return for the subsequent period to be negative. An equilibrium path for stock prices and the money supply will be sustained over a long period of time.

The Breusch-Godfrey procedure tests whether significant serial correlation is present in the residuals of the estimated equations (12) and (13).[9] Results indicate that the null hypothesis of no serial correlation cannot be rejected at the 5 percent significance level with either M1 or M2 as the measure of money.[10]

Because the error terms of the estimated equations are not autocorrelated, I test the following null hypotheses to determine the causal relationship between stock prices and money supply:

H_0: $\gamma_1 = 0$ and $\theta_{1i} = 0$, for all i,(i.e., no causality running from M to SP500); and

H_0: $\gamma_2 = 0$ and $\theta_{2i} = 0$, for all i,(i.e., no causality running from SP500 to M).

If the stock market is informationally efficient, then current prices should reflect all available information relevant to stocks. That is, including lagged values of money supply into an information set already consisting of lagged stock prices should not improve the forecast of stock prices. Also, movements in stock prices should lead movements in money supply in an efficient market. Therefore, SP500 unidirectionally should cause money supply in an efficient market.

Table 4 contains the results of the causality tests. Part A of the Table 4 for the January 1978 through September 1990 period shows that both M1 and M2 unidirection-

---

[9] According to the Breusch-Godfrey procedure, R^2 is obtained after the estimated residuals are regressed on both d lags of past residuals and the right side variables of the bivariate ECM. Under the null hypothesis of no serial correlation, TR^2 (where T is the number of observation) should be distributed as a χ^2 with d degrees of freedom. In view of the sample size, a fourth order lag (i.e., d = 4) is chosen for this test. See Godfrey (1978) and Breusch (1978) for more detail.

[10] For the equation (12), χ^2 statistic with four degrees of freedom has a value of 2.352 and 1.462 with M1 and M2, respectively, as the measure of money supply. For equation (13), the χ^2 statistic with four degrees of freedom has a value of 3.640 and 3.225 with M1 and M2, respectively, as the measure of money supply. The 5 percent critical value for the χ^2 statistic with four degrees of freedom is 9.488.

Table 3—Estimated Bivariate Error Correction Models (12) and (13): January 1978 to September 1990

(i) SP500 and M1

$$\Delta SP500_t = \underset{(-2.90)}{-7.6865} - \underset{(-3.51)}{.1040}(SP500_{t-1} - .4914M1_{t-1}) + \underset{(6.52)}{.5030}\,\Delta SP500_{t-1} - \underset{(-1.64)}{.1415}\,\Delta SP500_{t-2} - \underset{(-1.03)}{.1497}\,\Delta M1_{t-1}$$

Adjusted $R^2 = .2434$, $Q(36) = 40.8008$

$$\Delta M1_t = \underset{(1.15)}{1.5668} - \underset{(-3.33)}{.0016}(M1_{t-1} - 2.035\,SP_{t-1}) + \underset{(.21)}{.2693}\,\Delta M1_{t-1} + \underset{(.97)}{.0082}\,\Delta M1_{t-2} + \underset{(2.47)}{.2013}\,\Delta M1_{t-3} + \underset{(1.39)}{.0547}\,\Delta SP500_{t-1}$$

Adjusted $R^2 = .1197$, $Q(36) = 37.0830$

(ii) SP500 and M2

$$\Delta SP500_t = \underset{(-2.45)}{-7.2412} - \underset{(-3.28)}{.0800}(SP500_{t-1} - .1370\,M2t_{-1}) + \underset{6.24)}{.4841}\,\Delta SP500_{t-1} - \underset{(-1.91)}{.1623}\,\Delta SP500_{t-2} - \underset{(-1.53)}{.1104}\,\Delta M2_{t-1}$$

Adjusted $R^2 = .2380$, $Q(36) = 34.8722$

$$\Delta M2_t = \underset{(1.57)}{4.9316} - \underset{(-3.10)}{.0039}(M2_{t-1} - 7.2993\,SP500_{t-1}) + \underset{(4.58)}{.3511}\,\Delta M2_{t-1} + \underset{(1.89)}{.1482}\,\Delta SP500_{t-1}$$

Adjusted $R^2 = .1409$, $Q(36) = 31.0047$

t-statistics are in parentheses

SP500 represents the Standard & Poor's Composite Stock Index for 500 stocks at the end of the month (source: *Survey of Current Business*)

The data on M1 and M2 are obtained from various issues of the *Federal Reserve Bulletin*

ally cause SP500, and not vice versa, at the 5 percent significance level. This finding contradicts those of Kraft and Kraft (1977) and Rogalski and Vinso (1977), who document no causality running from M1 to SP500 from 1955 to 1974 and 1963 to 1974, respectively. Therefore, between January 1978 and September 1990, the U.S. stock market did not fully digest and reflect changes in both M1 and M2. That is, the stock market was not informationally efficient during this period.

In October 1979 the Federal Reserve System implemented a drastic change in its operating procedure; it moved from targeting the federal funds rate to focusing on nonborrowed reserves. In October 1982 the Fed changed its operating procedure to the borrowed reserve-based approach in face of an increasingly volatile relationship between M1 and various measures of economic performance. (See Bradley and Jansen, 1986; and Cosimano and Jansen, 1988 for empirical evidence supporting the shifts in the Fed's operating procedure.).[11]

In view of these changes in the Fed's operating procedure during the sample period, I next test the structural stability of equations (12) and (13) over this period. I test the null hypothesis of interperiod stability of the stock price-money supply relationship over three

[11] Kane (1982) attributes the failure of the Fed to control M1 between October 1979 and October 1982 effectively to financial deregulation; government imposition of credit controls from March to July of 1980, which restricted the growth of consumer and business loans; and the recession of 1981-1982.

Table 4—Granger Causality Tests for SP500, M1, and M2: January 1978 to September 1990 and for Three Subperiods

Causal Relationship	F-test	Degrees of Freedom
Panel A January 1978 to September 1990		
M1 $\not\Rightarrow$ SP500	6.2417*	(2,148)
SP500 $\not\Rightarrow$ M1	0.9893	(2,147)
M2 $\not\Rightarrow$ SP500	5.6738*	(2,148)
SP500 $\not\Rightarrow$ M2	2.1846	(2,149)
Panel B January 1978 to September 1979		
M1 $\not\Rightarrow$ SP500	2.1978	(2,16)
SP500 $\not\Rightarrow$ M1	0.2560	(2,15)
M2 $\not\Rightarrow$ SP500	1.3850	(2,16)
SP500 $\not\Rightarrow$ M2	0.2122	(2,17)
Panel C October 1979 to September 1982		
M1 $\not\Rightarrow$ SP500	0.0747	(2,31)
SP500 $\not\Rightarrow$ M1	1.5046	(2,30)
M2 $\not\Rightarrow$ SP500	0.3800	(2,31)
SP500 $\not\Rightarrow$ M2	2.4826	(2,32)
Panel D October 1982 to September 1990		
M1 $\not\Rightarrow$ SP500	4.0169*	(2,91)
SP500 $\not\Rightarrow$ M1	0.3960	(2,90)
M2 $\not\Rightarrow$ SP500	4.4414*	(2,91)
SP500 $\not\Rightarrow$ M2	1.0262	(2,92)

The null hypotheses tested are

H_0: $\gamma_1 = 0$ and $\theta_{1i} = 0$, for all i, (i.e., no causality running from M to SP500) and
H_0: $\gamma_2 = 0$ and $\theta_{2i} = 0$, for all i, (i.e., no causality running from SP500 to M) from the equations (12) and (13)

* The null hypothesis of no causality is rejected at the 5 percent level

SP500 represents the Standard & Poor's Composite Stock Index for 500 stocks at the end of the month (source: *Survey of Current Business*)

The data on M1 and M2 are obtained from various issues of the *Federal Reserve Bulletin*

subperiods: January 1978 to September 1979, October 1979 to September 1982, and October 1982 to September 1990. The Chow test rejects the null hypothesis at the 5 percent significance level,[12] indicating that changes in the Fed's operating targets during the sample period altered the relationship between stock prices and money supply.

---

[12] In regard to the relationship between SP500 and M1, F(17,109) = 7.528 with M1 as the dependent variable and F(5,145) = 1.432 with SP500 as the dependent variable. For the relationship between SP500 and M2, F(15,115) = 1.867 with M2 as the dependent variable and F(5,145) = 2.345 with SP500 as the dependent variable.

I also perform causality tests for each of the three subperiods.[13] As can be seen in Parts B through D in Table 4, both M1 and M2 unidirectionally cause stock prices during the third subperiod (October 1982 to September 1990), although the causal relationships are not detected for the earlier subperiods. The causal relationship between money supply and stock prices that appears to exist from January 1978 through September 1990 is due primarily to the October 1982 to September 1990 subperiod.[14]

The absence of a causal relationship between the money supply and stock prices from January 1978 through September 1982 implies that any observed relationship inferred from a regression equation between these two variables may be spurious. This finding is consistent with Fama's (1981) contention that a spurious correlation between these two variables existed during the 1954 through 1976 period because both variables are related to expected inflation.

Unidirectional causality running from both measures of money supply to stock prices (i.e., market inefficiency) during the third subperiod may be attributed to the gradual deregulation of the financial system, most of which was completed by the end of 1982. The elimination of the ceiling on yields paid by the financial institutions on most of their deposit liabilities subsequently made the distinction between various monetary aggregates fuzzy, thereby inducing a more volatile relationship between money supply and overall economic activity (Marlow, 1991).[15] Therefore, investors may have begun to question the role of money supply as a reliable leading economic indicator in the early 1980s. Such doubts may have led investors to overlook the relevant informational content subsumed in either M1 or M2 in predicting movements of stock prices from October 1982 to September 1990.

CONCLUSION

The relationship between the money supply and aggregate U.S. stock prices has been investigated extensively in the past. This study has attempted to provide additional insight into this issue by examining the money supply-stock prices relationship from January 1978 to September 1990, an era of increased uncertainty surrounding the U.S. financial market stemming from several changes in Federal Reserve operating procedures and the gradual elimination of various regulations on the deposits issued by financial institutions.

Empirical findings indicate that a long-run equilibrium relationship exists between stock prices and money supply during this period. Stock prices did not fully digest and

---

[13] Johansen's multivariate procedure is applied to test whether money supply and stock prices are cointegrated over each subperiod. The $-2\ln Q_r$ statistics for H_0: $r = 0$ (i.e., no cointegration between money supply and stock prices) are 11.02, 19.68, and 20.56 for three subperiods, respectively, with M1 as the measure of money. With M2 as the measure of money, the $-2\ln Q_r$ statistics are 16.26, 20.95, and 21.46 for the three subperiods, respectively. The critical values for $-2\ln Q_r$ are 17.96 (10 percent), 20.17 (5 percent), and 24.99 (1 percent) from Johansen and Juselius (1990, Table A3). Therefore, the cointegrating relationship between money supply and stock prices is found to exist for the second and third subperiods only. Accordingly, the causality test is performed for the first subperiod with the error correction term dropped from the equations (12) and (13).

[14] The estimated adjustment coefficient γ is found to be statistically significant at the 5 percent level and negative for the third subperiod. Moreover, the results are not sensitive to the choice of M1 or M2 as the measure of money supply.

[15] Judd and Motley (1984) also document that the relationship between money and real GNP started to go off track beginning in 1982.

reflect all available information on monetary policy changes during the period, however. This market inefficiency is due partly to the financial deregulation that began in 1978.

REFERENCES

1. Akaike, H., "A New Look at Statistical Model Identification," *IEEE Trans. Automatic Control AC-19* (1974), pp. 716-722.

2. Bradley, M.D., and D.W. Jansen, "Federal Reserve Operating Procedure in the Eighties: A Dynamic Analysis," *Journal of Money, Credit and Banking,* 18 (1986), pp. 323-335.

3. Breusch, T.S., "Testing for Autocorrelation in Dynamic Linear Models," *Australian Economic Papers,* 17 (1978), pp. 334-355.

4. Cosimano, T.F., and D.W. Jansen, "Federal Reserve Policy, 1975-1985: An Empirical Analysis," *Journal of Macroeconomics,* 10 (1988), pp. 27-47.

5. Dickey, D.A., and W.A. Fuller, "Distribution of the Estimators for Autoregressive Time Series with a Unit Root," *Journal of the American Statistical Association,* 74 (1979), pp. 427-431.

6. Dickey, D.A., and W.A. Fuller, "Likelihood Ratio Statistics for Autoregressive Time Series with a Unit Root," *Econometrica,* 49 (1981), pp. 1057-1072.

7. Dickey, D.A., D.W. Jansen, D.L. and D.L. Thornton, "A Primer on Cointegration with an Application to Money and Income," *Federal Reserve Bank of St. Louis Review* (March/April 1991), pp. 57-78.

8. Engle, R.F., and C.W.J. Granger, "Cointegration and Error Correction: Representation, Estimation and Testing," *Econometrica,* 55 (1987) pp. 251-276.

9. Engle, R.F., and B.S. Yoo. "Forecasting and Testing in Cointegrated Systems," *Journal of Econometrics,* 35 (1987), pp. 119-142.

10. Engle, R.F., and B.S. Yoo, "Cointegrated Economic Time Series: A Survey with New Results)," UCSD Working Paper #8-89013 (August 1989).

11. Fama, E.F., "Stock Returns, Real Activity, Inflation, and Money," *American Economic Review.* 71 (1981), pp. 545-565.

12. Fuller, W.A., *Introduction to Statistical Time Series* (New York: Wiley, 1976).

13. Gilbert, R.A., "Requiem for Regulation Q: What It Did and Why It Passed Away," *Federal Reserve Bank of St. Louis Review* (1986), pp. 22-37.

14. Godfrey, L.G., "Testing Against General Autoregressive and Moving Average Error Models When the Regressors Include Lagged Dependent Variables," *Econometrica,* 46 (1978), pp. 1293-1302.

15. Granger, C.W.J., "Investigating Causal Relations by Econometric Models and Cross Spectral Methods," *Econometrica,* 37 (1969), pp. 424-438.

16. Granger, C.W.J., "Co-integrated Variables and Error-Correcting Models," Department of Economics working paper (University of California, San Diego, CA, 1983)

17. Hall, R.E., "A Stochastic Life Cycle Model of Aggregate Consumption," *Journal of Political Economy,* 86 (1978), pp. 971-987.

18. Hsiao, C., "Autoregressive Modeling and Money-Income Causality Detection," *Journal of Monetary Economics,* 7 (1981), pp. 85-106.

19. Johansen, S., "Statistical Analysis of Cointegrating Factors," *Journal of Economic Dynamics and Control,* 12 (1988), pp. 231-254.

20. Johansen, S., and K. Juselius, "Maximum Likelihood Estimation and Inference on Cointegration—With Applications to the Demand for Money," *Oxford Bulletin of Economics and Statistics,* 52 (1990), pp. 169-210.

21. Judd, J.P., and B. Motley, "The 'Great Velocity Decline' of 1982-83: A Comparative Analysis of M1 and M2," *Federal Reserve Bank of San Francisco Economic Review* (1984).

22. Kane, E.J., "Selecting Monetary Targets in a Changing Financial Environment," *Monetary Policy Issues in the 1980s*, Federal Reserve Bank of Kansas City (1982).

23. Kraft, J., and A. Kraft, "Determinants of Common Stock Prices: A Time Series Analysis," *Journal of Finance*, 32 (1977), pp. 417-425.

24. Marlow, M.L., "Money and Velocity in the 1980s," *Quarterly Review of Economics and Business*, 31 (1991), pp. 36-46.

25. Nelson, C.R., and C. Plosser, "Trends and Random Walks in Macroeconomic Time Series," *Journal of Monetary Economics*, 10 (1982), pp. 139-162.

26. Rogalski, R., and J. Vinso, "Stock Returns, Money Supply and the Direction of Causality," *Journal of Finance*, 32 (1977), pp. 1017-1030.

27. Sims, C.A., J.H. Stock, and M.W. Watson, "Inferences in Linear Time Series Models with Some Unit Roots," Hoover Institutions working paper E87.1, revised version (1988).

28. Thornton, D.L., and D.S. Batten, "Lag-Length Selection and Tests of Granger Causality Between Money and Income," *Journal of Money, Credit and Banking*, 17 (1985), pp. 164-178.