Financial Deregulation, Monetary Policy, and Stock Market (In)Efficiency

A.F. Darrat*
Louisiana Tech University

R.N. Dickens
Louisiana Tech University

Lee (1994) finds evidence of market inefficiency using a bivariate model of monetary aggregates and stock prices; changes in monetary stock unidirectionally cause stock prices. We use the same test period and data as the Lee study, but estimate a multivariate model using variables suggested by the literature. We find (slight) evidence that stock prices cause changes in monetary stocks, but no evidence suggesting market inefficiency. Our results are robust for different testing procedures and provide further evidence of efficient financial markets.

INTRODUCTION

Researchers have expended enormous effort assessing whether money growth provides a leading indicator of stock prices. The bulk of empirical evidence supports the stock market efficiency (SME) hypothesis that common stock prices reflect all available information on monetary policy moves. Examples of such compelling evidence include Fama (1970), Rozeff (1974), Kraft and Kraft (1977), Rogalski and Vinso (1977), Davidson and Froyen (1982), Sorensen (1982), Pearce and Roley (1985), Pearce (1987), and Serletis (1993), to name just a few.

Lee has studied the subject more recently (1994). Lee uses Granger causality tests and error-correction modeling to examine the relationship between monetary aggregates (M1 and M2) and stock prices (SP500) from January 1978 to September 1990. His empirical results reveal the presence of a significant long-run cointegration relationship between the monetary aggregates (both M1 and M2) and stock prices during the estimation period and attribute this cointegration to the money stock unidirectionally causing stock prices particularly from October 1982 to September 1990.

Lee concludes, "Stock prices did not fully digest and reflect all available information on monetary policy changes during this period," (p. 48). Lee ascribes this apparent inefficiency finding to financial deregulation during the estimation period. Starting in the early 1980s, regulatory authorities allowed depository institutions to issue various interest-bearing checkable accounts (e.g., NOW accounts, Super-NOWs and money market demand accounts). The period's deregulatory changes blurred the traditional distinction among the monetary aggregates. [See Mehra (1989) and Gauger and Schroeter (1990).] In particular, M1 no longer can be regarded as a transactions only aggregate because it contains interest-bearing balances. Such developments may have increased uncertainty in the U.S. financial environment, thereby causing alleged market inefficiency during the period.

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Lee's empirical findings raise the possibility that a diligent investor could reap abnormal profits by exploiting a trading rule based on the observed behavior of monetary policy, implying a rejection of the SME hypothesis, at least from 1978 to 1990. Lee's empirical evidence appears shaky due to possible misspecification biases.

A striking feature of Lee's model is that it is bivariate, containing only two variables: M1 (or M2) and stock prices. Yet Lutkepohl (1982) and others long have demonstrated that Granger causality results from a bivariate model are misleading because of the omission of important variables. One variable (x) may appear to cause another variable (y) in a bivariate model, but this outcome may be the result of other relevant (but omitted) variables propagating the observed causality between x and y. Conversely, if in the bivariate system, x is found not to cause y, this does not necessarily imply that such an inference would hold in the context of a larger economic system that includes other relevant variables. Lutkepohl (1982, p. 367) notes, "this conclusion is a consequence of the well-known problem that a low dimensional subprocess contains little information about the structure of a higher dimensional system." Moreover, Granger (1988) also has shown that cointegration and Granger causality are closely related concepts. Cointegration results from Lee's bivariate model also could be susceptible to a similar omission-of-variables bias. This conjecture receives empirical support from Miller (1991) who reports that omitting relevant variables from his model significantly distorts cointegration inferences.

This paper demonstrates that Lee's bivariate model is too restrictive, as it overlooks several variables relevant for modeling stock prices. Results from an expanded model significantly alter Lee's main conclusion and suggest that the stock market was efficient even during the heavy deregulatory period.

AN EXPANDED MODEL AND EMPIRICAL RESULTS

AN EXPANDED MODEL

Established literature on modeling stock prices provides a logical extension to Lee's bivariate model. Ross's (1976) arbitrage pricing theory (APT) and various developments in subsequent research reveal that more than money supply is required to model stock prices. Fama (1981), Chen, Roll, and Ross (1986), and Darrat and Brocato (1994) suggest the need to incorporate the effects of three more variables: inflation, industrial production, and interest rates.[1]

For maximum comparability with Lee's study, we reestimate our five variable model of stock prices over the same time period as his,[2] use identical definitions for the monetary aggregates and stock prices as Lee,[3] and employ similar testing procedures. Hence, any differences in the results should be primarily due to our broader (multivariate) frame-

[1] These variables are not exhaustive; other variables also may be potential candidates. Our main purpose is not to specify the most comprehensive model for stock prices. Rather, our task is much simpler and illustrative, aimed at showing that a slightly better model is consistent with the efficient market assumption.

[2] Extending the estimation period through December 1994 yields similar results, which are available from the authors upon request.

[3] We compile our data on the two variables examined by Lee (SP500 and M1) from the same source; the *Survey of Current Business* for SP500, and from monthly issues of the *Federal Reserve Bulletin* published between April 1978 and December 1990 for M1.

work. Because Lee's inefficiency results are similar for M1 and M2 (and perhaps somewhat stronger when M1 is used[4]), we focus below on results for the narrow M1 stock in the proposed multivariate setting.

UNIT ROOTS TEST RESULTS

Following Lee, we employ the augmented Dickey-Fuller (ADF) test to examine unit roots (nonstationarity) in each of the five variables. These variables are the narrow money stock (M1), stock prices (SP500), industrial production (IP), interest rates (TB) as measured by the yield on three month Treasury bills, and the inflation rate (NF) as measured by the percentage change in the CPI. We obtain the monthly data on three additional variables (NF, IP, and TB) from the Citibank data tape. We determine the proper lag structure for the ADF test by means of the Akaike final prediction error (FPE) criterion. Thornton and Batten (1985) report evidence supporting the use of the FPE procedure over other alternative criteria (including the AIC procedure).[5]

Table 1 (Panel A) reports the results from applying the ADF test on the five variables in levels and in first differences over the estimation period (January 1978 to September 1990). We also perform the ADF test on each variable from October 1982 to September 1990, a period identified by Lee as the most turbulent financial deregulation period allegedly responsible for his inefficiency evidence. Table 1 suggests that all variables, save one, are nonstationary in levels, but stationary if expressed in first differences.[6] The only exception is the inflation variable which appears stationary in levels from October 1982 to September 1990. These results seem robust to Schwert's (1987) argument that the above ADF tests may be biased if the underlying time series are generated by moving as well as autoregressive elements. Following Mehra (1991), we check for Schwert's criticism using longer lags than those suggested by the FPE criterion and repeat the tests. The longer lags do not alter the unit root results we report in Table 1.

COINTEGRATION TEST RESULTS

The next step in the empirical analysis is to test for cointegration among the five variables because they are integrated of the same (first) order. We perform, as Lee (1994) does, the Engle-Granger (1987) two stage testing procedure. In the first stage, we estimate appropriate cointegrating regressions for the variables in their nonstationary (level) forms and retrieve the resultant residuals. The second stage involves applying a similar ADF test on the levels of the estimated residuals and checking for the presence of unit roots (nonstationarity). The variables are cointegrated if these residuals are stationary.

When more than two variables are being tested for cointegration, as is the case here, a problem arises with the Engle-Granger procedure because the cointegration vector may not be unique. [See Banerjee *et al.* (1986).] Thus, a choice must be made regarding which of

[4] For example, compared to M2, the results for M1 exhibit stronger Granger causality effects in terms of the size of the associated coefficients and the calculated F-statistics. Such a difference between M1 and M2 should not be surprising because the deregulation impacts are more pronounced for M1.

[5] This FPE procedure is asymptotically equivalent to the Akaike information criterion (AIC) used by Lee.

[6] Lee finds essentially similar unit root results for M1 and SP500 from January 1978 to September 1990.

Table 1—Augmented Dickey-Fuller Unit Root Test Results

Variable (y)	Test Statistics	Appropriate Lag Lengths	BG Statistics
Panel A: Sample Period January 1978 to September 1990			
In Levels			
Narrow Money Stock (M1)	-2.24	5	7.05
Standard & Poor's 500 (SP500)	-2.72	2	7.20
Industrial Production (IP)	-2.69	3	8.58
Three Month Treasury Bill Rate (TB)	-1.27	6	10.92
The Inflation Rate (NF)	-1.58	7	3.95
In First Differences			
ΔM1	-3.65**	4	7.19
ΔSP500	-7.78**	1	7.49
ΔIP	-9.25**	1	7.20
ΔTB	-7.15**	5	10.90
ΔNF	-7.58**	7	9.30
Panel B: Sample Period October 1982 to September 1990			
In Levels			
Narrow Money Stock (M1)	-1.28	3	8.26
Standard & Poor's 500 (SP500)	-2.55	2	6.35
Industrial Production (IP)	-0.94	1	7.57
Three Month Treasury Bill Rate (TB)	-1.48	2	3.13
The Inflation Rate (NF)	-5.03**	3	10.44
In First Differences			
ΔM1	-4.54**	1	8.18
ΔSP500	-6.00**	1	6.61
ΔIP	-4.35**	2	7.14
ΔTB	-5.88**	1	1.91
ΔNF	-6.43**	6	5.95

The testing equations of the ADF procedure are of the form:

$$\Delta y_t = \alpha + \sum_{s=1}^{n} ds\Delta y_{t-s} + \rho y_{t-1} + \beta T + \text{residuals}$$

where T is a time trend included whenever proved statistically significant, and n is the lag length determined by Akaike's final prediction error criterion. An ** indicates rejection of the null hypothesis of a unit root (nonstationarity) at the 5 percent level of significance. The BG is the Breusch-Godfrey chi-squared statistic to test the null hypothesis of no autoregressive and moving average processes in the error term of up to the sixth order

the variables we use as the left-side conditioning variable. We follow Hall (1986) and Miller (1991) and resolve this normalization problem by examining all possible cointegrating regressions and selecting the regression that maximizes the adjusted R^2.[7] The equation of the monetary aggregate (M1) provides the highest adjusted R^2; thus, we choose it as the proper regressand in the cointegrating equation.

We report results from the Engle-Granger test in Table 2 for the entire estimation period (Panel A). In contrast to Lee, we can not reject the null hypothesis of no cointegration among the variables (test statistic = -1.81, 5 percent critical value = -4.72). This finding of no cointegration is corroborated further by the cointegration regression Durbin-Watson (CRDW) test score (= 0.12, compared to the 5 percent critical value of 0.39).[8]

[7] The Monte Carlo study of Banerjee *et al.* (1986) recommends a similar practice.

[8] See Charemza and Deadman (1992) for a discussion on the merits of the CRDW test.

Table 2—Multivariate Cointegration Test Results

Engle-Granger's Test Statistics	Appropriate Lag Lengths	CRDW's Statistics	Johansen's Statistics
Panel A: Sample Period January 1978 to September 1990			
-1.81	2	0.12	80.90
Panel B: Sample Period October 1982 to September 1990			
-2.49	1	0.15	57.68

The Engle-Granger procedure involves two stages. First, we estimate a proper (M1) cointegrating equation and retrieve the resulting residuals (e). In the second stage, we apply the following ADF test on these residuals:

$$\Delta e_t = \lambda e_{t-1} + \sum_{s=1}^{k} g_s \Delta e_{t-s}$$

The CRDW is the cointegration-regression Durbin-Watson test. The test statistics for the Engle-Granger and the CRDW procedures are below the corresponding critical values. [See Maddala (1992, p. 607).] This implies nonrejection of the null hypothesis of no cointegration. The Johansen procedure gives a similar conclusion, whereby the 5 percent critical value is 87.31 for Panel A and 62.99 for Panel B

We obtain similar conclusions when we include a time trend in the testing equations. Thus, our results are not sensitive to the possible bias discussed by Engle and Yoo (1989). Again, we perform all the tests over the heavy deregulation period (October 1982 to September 1990).[9] The results, displayed in Panel B of Table 2, suggest a similar inference that there is no cointegration among the variables. (The Engle-Granger test statistic is -2.49, and the CRDW score is 0.15, both of which are far below the corresponding 5 percent critical values.)[10]

Is our evidence against cointegration between U.S. stock prices and the remaining four variables (including M1) procedure-specific? That is, how robust is our evidence of no cointegration compared to other tests? Unlike Lee's model, ours is multivariate—and a more appropriate test of cointegration should incorporate possible dynamic interactions among the various variables. An attractive procedure in this regard is the testing procedure proposed by Johansen (1988, 1991). As Dickey *et al.* (1991) argue, the Johansen approach is promising because it is based on the well-established likelihood ratio principle. In addition, the Johansen method is not subject to biases due to possibly incorrect normalization choices because all of the variables are jointly endogenized in the testing process. Monte Carlo evidence reported by Gonzalo (1989) also supports the relative power of the Johansen methodology over other alternative techniques.[11]

Results from the Johansen test also are reported in Table 2 for the full sample period (Panel A) and for the period from October 1982 through September 1990 (Panel B). It is reassuring that the evidence from the Johansen test is the same and consistently suggests

[9] Over this period the inflation variable seems integrated of order zero, in contrast to the remaining four variables that are integrated of order one. Thus, inflation cannot be integrated with the other variables in the model and is excluded from the cointegrating equation.

[10] While we focus on the relationship between M1 and stock prices, we perform the analysis with M2 and the results are similar.

[11] Interestingly, Lee also applies the Johansen test in his bivariate model, although the test's practical usefulness is primarily for a multivariate setting.

the absence of any significant cointegration between the monetary aggregate (M1) and stock prices over the entire estimation period as well as over the heavy deregulatory period.

Regardless of the test we use or the particular estimation period we examine, the evidence from our multivariate model is decisive and indicates that there is no cointegration relationship between M1 and stock prices. Thus, contrary to Lee, but in accordance with the efficient market hypothesis, money and stock prices in the U.S. do not cointegrate, do not possess a reliable (exploitable) long-run relationship, and have no error-correction representation. Of course, the lack of a long-run relationship does not necessarily negate the presence of some short-run Granger causal linkage (Granger, 1991), and we investigate this possibility next.

GRANGER CAUSALITY TEST RESULTS

Absence of an error correction (long-run) component from the stock price dynamics reduces the process to the conventional Granger causality framework. In this case, we can interpret the lagged coefficients on the monetary variable in the stock price equation as summarizing the short-run Granger-causal impact.[12] A similar interpretation can be made for the lagged coefficients on stock prices in the money equation.

Application of the Granger causality tests requires stationary series. The variables are integrated of order one, so their first differences are stationary (integrated of order zero). Again, we use the FPE procedure to specify a multivariate equation for stock prices and another for M1. Although Lee (1994) tests for Granger causality, he does so in the restrictive bivariate framework and in the presence of a misleading error-correction component. Interestingly, Lee (p. 47) reports results showing that his Granger causality equations are structurally unstable according to a Chow test. Lutkepohl (1989) demonstrates that these equations must be stable to yield credible Granger causality inferences. Structural instability in Lee's equations seems closely related to their bivariate nature as our multivariate Granger causality equations prove highly stable.

Table 3 reports the results from the multivariate Granger causality tests. (Panel A contains the results for the full period, and Panel B has the results for the heavily deregulated period.) Unlike Lee's bivariate equations, our multivariate equations prove structurally stable according to the Chow test. The short-run Granger causality results in Table 3 generally corroborate the cointegration (long-run) results in rejecting the presence of any reliable (lagged) relationship between M1 and stock prices. Such a verdict is clear in the case of the subperiod (Panel B), whereby M1 and stock prices appear causally independent and void of any causal linkage in either direction. For the entire period (Panel A) the results indicate some weak (at the 10 percent level) causal relationship between M1 and SP500, but this causality runs from the latter to the former. In other words, there seems to be some weak evidence that stock prices may be a leading indicator of monetary policy.[13] Such a relationship is consistent with the efficient market hypothesis, however,

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[12] For a similar practice, see Jones and Joulfaian (1991) and Bahmani-Oskooee and Payesteh (1993).

[13] This result should not be surprising as movements in stock prices traditionally typically precede changes in real economic activity, the latter being one of the ultimate objectives of monetary policy.

Table 3—Multivariate (Short-Run) Granger Causality Test Results

Null Hypothesis	F-Statistics	Appropriate Lag Length
Panel A: Sample Period January 1978 to September 1990		
Dependent Variable: SP500 [BG = 3.30, Chow-F (82:10) = 0.96]		
M1 does not Granger cause SP500	1.83	11
Industrial production does not Granger cause SP500	1.10	1
Inflation does not Granger cause SP500	0.02	1
Short-term interest rates do not Granger cause SP500	0.20	1
Dependent Variable: M1 [BG = 5.02, Chow-F (82:10) = 1.01]		
Short-term interest rates do not Granger cause M1	4.84**	3
Industrial production does not Granger cause M1	1.94**	12
SP500 does not Granger cause M1	2.78*	1
Inflation does not Granger cause M1	1.14	1
Panel B: Sample Period October 1982 to September 1990		
Dependent Variable: SP500 [BG = 2.37]		
M1 does not Granger cause SP500	1.78	11
Industrial production does not Granger cause SP500	1.18	1
Inflation does not Granger cause SP500	0.66	1
Short-term interest rates do not Granger cause SP500	0.35	1
Dependent Variable: M1 [BG = 12.64]		
Inflation does not Granger cause M1	3.69**	11
Short-term interest rates do not Granger cause SP500	2.71**	9
Industrial production does not Granger cause SP500	3.61*	1
SP500 does not Granger cause M1	0.86	2

See notes to Table 1. The orders of the inclusion of the variables in the various equations are determined by the specific gravity criterion. An * indicates rejection of the null hypothesis at the 10 percent level of significance and ** indicates rejection at the 5 percent level

and it merely implies that the Federal Reserve takes into account the behavior of the stock market in its policy decision-making.[14]

CONCLUDING REMARKS

This paper empirically reexamines the notion supported by Lee (1994) that U.S. financial deregulation of the early 1980s has caused some inefficiency in the stock market with respect to available information on monetary policy. Such a notion contradicts the large and growing consensus among financial analysts and economists who repeatedly demonstrate that the U.S. stock market is efficient, and that investors cannot earn abnormal profits by employing a trading strategy based on the observed behavior of monetary policy.

Although several anomalies long have been discussed in the literature,[15] the question is whether financial deregulation represents yet another anomaly. The evidence in this paper rejects this conjecture and indicates that the culprit behind the apparent contradictory conclusion in Lee's recent study is model misspecification rather than market inefficiency *per se*. Lee obtains his Granger causality and cointegration results in the context of a bivariate system. Recent research has convincingly established, however, that such a

[14] In addition to stock prices, all other macro variables have some predictable values for monetary policy moves. Mishkin (1982) reports similar evidence.

[15] For example, see Bjerring *et al.* (1983) for the Value Line phenomenon, Lustiq and Leinbach (1983) for the small firm effect, Keim (1983) for the January effect, and Keim and Stambaugh (1984) for the weekend effect.

bivariate process is overly restrictive and generally yields misleading inferences. Our results from a broader model (that is similar to Lee's in all other aspects) suggest that the U.S. stock market continued to be informationally efficient during the recent financial deregulatory period.

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