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Real Estate Returns, Money and Fiscal Deficits: Is the Real Estate Market Efficient?

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Abstract

This research examines the causal relationship between several financial variables and a portfolio of real estate returns using monthly data from January 1965 to December 1986. The empirical analysis is based on multivariate Granger-causality tests in conjunction with Akaike's final prediction error criterion. The results indicate that measures approximating monetary policy and market returns play an important role in causing changes in real estate returns. In particular, our findings suggest that base money and market returns have had significant lagged effects on current real estate returns.

Numerous empirical studies have examined the validity of the stock market efficiency (SME) hypothesis, particularly in regard to the role of monetary policy. In its semi-strong form, the SME hypothesis contends that stock prices reflect rapidly all publicly available (lagged) information including monetary policy moves. Our main purpose is to assess empirically the relationship between monetary policy (and other financial variables) and real estate returns. We hypothesize that while the stock market as a whole might not exhibit a lagged relation with money growth and other key variables, particular industries may. As different industries react to and against each other, the overall lagged relationship could thus be masked.

The empirical evidence so far has not been inconsistent with both the SME hypothesis and the portfolio theory views of the stock market. [For example, see Tanner and Trapani (1977), Fama (1981), Davidson and Froyen (1982), Sorensen (1982), Wong (1986), and Pearce (1987).] However, there is considerable evidence that the market may not be as efficient as the SME hypothesis posits.

Shiller (1979,1981) found that bond price changes cannot be explained by rational expected changes in future interest rates. De Bondt and Thaler (1985) found a reversal anamoly. Securities that experienced high yields in one period had low yields in subsequent periods. Waud (1970) found that Federal Reserve discount rate changes were correlated with stock price changes, and Homa and Jaffee (1971) used money supply to predict Standard and Poor's 500 index levels.

However, neither of these latter studies has been confirmed with trading rule tests.[1]

Real estate presents a unique anomoly—it has earned high returns while having low risk.[2] The general expectation is that real estate is a good inflation hedge (see Sirmans and Sirmans, 1987). Fogler, Granito, and Smith (1985) attempt to explain this occurrence. They examine two explanations: first, real estate may have had high returns because of unexpectedly high rates of inflation; second, there may have been a structural shift in the demand for real estate assets because of relative (perceived) shifts in the inflation beta sensitivities of real estate assets and stocks. While they were unable to distinguish firmly between their hypotheses, they raise the issue of investor perceptions as being important in analyzing anomoly behavior. To the extent that monetary and fiscal policy impact perceptions about real variables and their interaction, real estate prices would be influenced.

At least two interesting questions are addressed in this article. First, do real estate returns react with a significant lag to historical data on monetary policy and other financial variables? Second, do these returns react to movements in federal budget deficits in addition to money growth and market portfolio? The empirical evidence on the economic effects of federal deficits has been mixed. For example, Hoelscher (1983), Evans (1985, 1987), and Plosser (1987) have reported results suggesting that deficits are irrelevant for determining interest rates and perhaps for almost everything else in the economy. On the other hand, studies by Feldstein (1982), DeLeeuw and Holloway (1985), Darrat (1986), Hoelscher (1986), and Zahid (1988) have indicated the importance of federal deficits in influencing economic activity. Curiously, none of the previous studies examined the effects of federal deficits on the real estate market. This seems surprising since the mortgage market is so closely tied to the government bond market. If government bonds represent a component of the asset holders' total portfolios, changes in bond-financed federal deficits could disturb the equilibrium position of government bonds relative to the other assets in the portfolios. Attempts on the part of the asset holders to return to portfolio equilibrium would result in asset substitutions and consequent price changes. Therefore, movements in federal deficits could have important wealth effects on real estate returns. Moreover, concerns about the implications of high budget deficits and the associated increases in overall economic uncertainty may cause the equity risk premium to rise, adversely affecting real estate returns. Finally, if higher budget deficits raise interest rates, other things remaining unchanged, bonds become more attractive to hold than alternative stocks (including real estate assets), leading in turn to a fall in stock prices. Of course, whether budget deficits do exert the theoretically hypothesized effect on real estate stocks can only be answered empirically.

The rest of the article is organized as follows. Section 1 describes the data and the empirical methodology used. Section 2 reports the results and analyzes their implications. Section 3 offers some concluding remarks.

1. Data and methodology

Monthly data spanning the period January 1965 through December 1986 are used. The real estate stock returns are for a portfolio of real estate firms listed on the CRSP (Center for Research in Security Prices, the University of Chicago) tapes. These represent firms of the American and New York Stock Exchanges. There are 31 firms in the portfolio consisting of three primary types of firms: real estate investment trusts (REITS), builders and investment, and management firms.[3]

Although the bulk of real estate assets are not traded on the American or New York Exchanges, the use of exchange data allows us to examine the relationship between real estate firm returns and macro variables. Of course, research that uses index or appraisal values is subject to estimation error. By using actual trading prices, we have the market's estimate of return changes without the leveling and other effects of appraisal approaches. In this direction, we follow the lead of Hite, Owners, and Rogers (1984), Titman and Warga (1986), and Allen and Sirmans (1987).

This article estimates whether changes in monetary and fiscal policy and other financial variables have significant causal effects on real estate stock returns. The empirical analysis is performed using multivariate Granger-causality tests in conjunction with Akaike's final prediction error (FPE) criterion.

Three comments about the procedure are appropriate. First, in addition to monetary and fiscal policy variables, other variables are included in our analysis to avoid the "omission of variables" problem (see Lutkepohl, 1982). We include: industrial production index (as a proxy for real gross national product, or GNP); a proxy for risk premium measured, following Chen, Roll, and Ross (1986), by the difference between the long-term Treasury bond yield and the BBB Moody's index yield; the return on the value-weighted market index from CRSP; the unemployment rate; and the rate of inflation. These additional variables are those frequently hypothesized to be important determinants of stock returns. [See, for example, Tanner and Trapani (1977), Sorensen (1982), and Chen, Roll, and Ross (1986).]

Second, several previous applications of multivariate Granger-causality tests regress the dependent variable on its own lagged values and the lagged values of other candidate variable using a common lag length on all variables. [See, for example, Dwyer (1982), Mishkin (1982), Plosser (1987), and Koray and Hill (1988).] Causality inferences are then generated on the basis of the joint significance of coefficients. However, as discussed by Ahking and Miller (1985), the use of a common lag length for all variables has no basis in theory and can produce biased results if some variables exhibit different lag specifications. To avoid this problem, we use Granger-causality tests in conjunction with Akaike's final prediction error (FPE) criterion to determine the appropriate lag length for each explanatory variable in the model. Such a procedure was originally proposed by Hsiao (1979, 1981) for the bivariate case and has been applied to multivariate models by several other authors [e.g., McMillin (1986) and Darrat (1988)].

Finally it should be emphasized that the tests are performed here using only lagged values of the explanatory variables to ensure the availability of data. As such, our model is predictive in the sense of Rozeff (1974).

Using Granger's (1969) popular definition of causality, a variable (X) is said to Granger-cause another variable (Y) if the prediction error of current Y declines by using past values of X in addition to past values of Y. Based on the Monte Carlo evidence of Geweke, Meese, and Dent (1983), the one-sided distributed lag test of Granger is used here. The empirical technique can be briefly summarized as follows:

1. The FPE/multivariate Granger-causality tests presume the use of stationarity data. Following previous work, stationarity is achieved for every variable by expressing it in a logarithmic form and then applying the appropriate degree of differencing. Thus, each variable in the logarithmic first-difference form is regressed against time and a constant. If the coefficient on time proves insignificant, the variable in that form is considered stationary. Otherwise, a second (third, etc.) difference is used until the coefficient on time losses significance. The Granger-causality tests further require that the stationary series have zero mean. We accomplish this by adding a constant to all regressions.

2. After transforming the variables to stationary series, the next step is to determine the appropriate own lag of real estate security returns. This is done by searching for the lag that minimizes FPE within the following univariate model:

$$DRER_t = \beta_0 + \beta_1^k(L)DRER_t + e_t \tag{1}$$

where $DRER$ is the real estate portfolio return (in percent changes), L is the lag operator, $\beta_1(L)$ is a distributed lag polynomial in L, k is the order of the lag ($k = 1, 2, \ldots 12$),[4] and e is the associated error term. The FPE is defined as:

$$FPE(k) = [(n + k + 1)/(n - k - 1)] \cdot RSS(k)/n \tag{2}$$

where n is the number of observations and RSS is the residual sum-of-squares. The lag length that minimizes the FPE is considered appropriate (k^*). The FPE criterion for lag selection is appealing, as Hsiao (1981) noted, because it balances the risk of selecting a higher lag against the risk of selecting a lower lag. When an additional lag is included, the first term in the FPE formula is increased, but simultaneously the second term is decreased. When their product (i.e., the FPE) reaches a minimum, the two opposing forces are balanced.

3. Once the appropriate own lag is determined, the following bivariate regressions are then estimated:

$$DRER_t = \beta_0 + \beta_1^{k*}(L)DRER_t + \beta_2^h(L)Z_t + e_t \tag{3}$$

where Z is the vector of the other candidate variables, considered one at a time, and h is the lag order, again varying from 1 to 12 lags. The appropriate lag order of h (h^*) is that which minimizes the following modified *FPE* (observe that the own lag is fixed at its previously obtained order, k^*):

$$FPE(k^*,h) = [(n + k^* + h + 1)/(n - k^* - h - 1) \cdot RSS(k^*,h)/n. \tag{4}$$

4. Next is the estimation of trivariate regressions. At this stage, however, we need to determine the rank by which the explanatory variables are included in the model. Following Caines, Keng, and Sethi (1981), this ranking is based on the "specific gravity" criterion. Consequently, to the appropriate own lag (k^*), we add first the variable with the highest specific gravity (or lowest minimum *FPE*) among all non-real estate return variables with its appropriate lag length determined in the previous step. Thus the trivariate equations take the form:

$$DRER_t = \beta_0 + \beta_1^{k*}(L)DRER_t + \beta_2^{h*}(L)Z_{1,t} + \beta_3^{m}(L)Z_t + e_t \tag{5}$$

where $Z_{1,t}$ is the variable with the highest specific gravity among all bivariate regressions, and Z is the vector of remaining variables, considered one at a time. As before, the lag m is varied from 1 to 12, and the appropriate lag length is that which minimizes an analogous *FPE*.

5. Continue the process with the trivariate (quadrivariate, etc.) models until all variables are included in the final real estate return model, each with its appropriate lag specification.

6. Subject the final model to diagnostic checks to assess its structural stability and serial correlation properties.

7. The final step is the tests of coefficient significance for each variable in the real estate pricing model. These tests are based on a series of likelihood ratio (LR) statistics that are asymptotically distributed as $\chi^2(q)$, where q is the number of restrictions. Note that previous studies that used the *FPE* criterion have often directly compared the values of *FPE* obtained, say, from the bivariate and trivariate systems to generate causality inferences. Thus, formal significance tests have typically been ignored. To avoid the obvious shortcomings of such an approach, formal likelihood ratio tests are used instead to obtain our causality inferences.

2. Empirical results

2.1. Model specification

Applying the previous econometric procedure over the monthly period from January 1985 to December 1986 (before adjustment for lags), the following real estate return model is obtained and estimated using ordinary least squares (OLS):

$$DRER_t = \beta_0 + \beta_1^{10}(L)DRER_t + \beta_2^6(L)DDB_t + \beta_3^{11}(L)DMR_t + \beta_4^1(L)DUE_t + \beta_5^1(L)DIP_t + \beta_6^1(L)DIN_t + \beta_7^1(L)DFP_t + \beta_8^1(L)DRP_t + e_t \tag{6}$$

where $DRER_t = (1 - L)\log RER_t$, RER is the real estate portfolio return for 31 real estate securities; $DDB_t = (1 - L)^2 \log B_t$, B is the monetary base (monthly averages); $DMR_t = (1 - L)\log MR_t$, MR is the return on the value-weighted market portfolio from CRSP; $DUE_t = (1 - L)\log UE_t$, UE is the unemployment rate for all workers; $DIP_t = (1 - L)\log IP_t$, IP is the industrial production index; $DIN_t = (1 - L)\log IN_t$, IN is the inflation rate measured by the consumer price index; $DFP_t = (1 - L)FD_t$, FD is the federal budget deficit; $DRP_t = (1 - L)RP_t$, RP is the risk premium proxy; L is the lag operator, and e is a white noise error term. The coefficient estimates are presented in table 1.

Table 1. Coefficient estimates of the base model (6)[a]

Variables	Coefficient Estimates	Absolute Values of t-Statistics	Variables	Coefficient Estimates	Absolute Values of t-Statistics
Constant	0.082	0.37			
DRER (t-1)	−0.623	8.15	DMR (t-1)	−0.184	1.01
DRER (t-2)	−0.690	7.98	DMR (t-2)	0.073	0.29
DRER (t-3)	−0.642	6.38	DMR (t-3)	−0.129	0.41
DRER (t-4)	−0.540	4.88	DMR (t-4)	−0.106	0.32
DRER (t-5)	−0.539	4.73	DMR (t-5)	0.282	0.84
DRER (t-6)	−0.413	3.64	DMR (t-6)	−0.052	0.15
DRER (t-7)	−0.324	2.92	DMR (t-7)	−0.143	0.43
DRER (t-8)	−0.139	1.36	DMR (t-8)	−0.623	1.93
DRER (t-9)	−0.268	3.11	DMR (t-9)	−0.441	1.48
DRER (t-10)	−0.036	0.49	DMR (t-10)	−0.685	2.78
			DMR (t-11)	−0.262	1.58
DDB (t-1)	0.446	1.04			
DDB (t-2)	0.562	0.93	DUE (t-1)	0.065	1.35
DDB (t-3)	0.233	0.34	DIP (t-1)	−0.084	0.52
DDB (t-4)	−0.233	0.34	DIN (t-1)	−0.151	0.42
DDB (t-5)	−1.477	2.45	DFP (t-1)	−0.00003	0.30
DDB (t-6)	−0.959	2.20	DRP (t-1)	0.001	0.20

Model Summary Statistics: $R^2 = 0.49$, standard-error of the regression = 1.892, F = 6.46, DW = 1.97
Breusch-Godfrey chi-square statistic = 14.28

[a]Dependent variable: $DRER_{(t)} = [\log RER_{(t)} - \log RER_{(t-1)}]$; monthly data from January 1965 to December 1986.

Notes: F is an F-statistic to test the null hypothesis that all the independent variables except for the constant term are jointly zero (d.f. = 32,217), and DW is the Durbin-Watson statistic.

As the table indicates, model (6) fits the data quite well as about half of the total variation in real estate prices is explained by the model. Considering that the dependent variable is cast in percentage changes, an $R^2 = 0.49$ is indeed sufficiently high. As to the serial correlation feature of the regression, the Durbin-Watson test, although reported, is known to be biased in the presence of a lagged dependent variable among the regressors. However, the Breusch-Godfrey test for up to the tenth order is used instead and evidences no significant serial correlation. As Johnston (1984) argues, the virtue of the Breusch-Godfrey procedure is that it is a robust test against general autoregressive and moving-average processes of the error term. In view of its critical importance to tests of market efficiency, the Geary nonparametric test was also performed to assess the serial correlation of the residuals. The Geary test, too, indicated no significant serial correlation of any type (the tau statistic = 133 in 250 observations).

Another desirable feature of the estimated model is its temporal stability. To assess the structural stability of our model, two tests were employed. Using the OPEC price hike of October 1973 as the breaking date, the Chow (1960) test was applied. Following Farley, Hinich, and McGuire (1975), the mid-point was also used as a breaking date to maximize the empirical power of the Chow test. None of the two breaking dates indicated any structural shift in the estimated equation. (The calculated F-values are, respectively, 0.74 and 1.35 with 33,184 degrees of freedom.) In addition to the Chow test, the Farley and Hinich (1970) test was also applied. While the Chow technique tests for a single point shift, the Farley and Hinich procedure tests for a continuous drift in the estimated relationship. The results from the Farley and Hinich test confirm the absence of any significant shift in the estimated equation reported above. To conserve on space, the table reporting the Farley and Hinich test results is not provided here, but is available from the authors upon request.

2.2. Model implications

The focus of this research is on the efficient market hypothesis that contends that past (lagged) public information on monetary, fiscal, and other financial variables should already be impounded in current real estate security returns. That is, the lagged coefficients on the various explanatory variables in equation (6) should be insignificantly different from zero. With respect to the unemployment rate, industrial production, inflation, fiscal deficit, and risk premium, this implication of the market efficiency could not be rejected. Specifically, based on the FPE/multivariate Granger-causality tests, the appropriate lag for each of these variables is only one month. More important, the one-month lagged coefficients are all statistically insignificant. Of course, real estate returns may respond contemporaneously to changes in these variables. As discussed above, such contemporaneous values of the variables were omitted from the estimated equation in order to generate credible tests of market efficiency.

However, in regard to the base money and the market returns, the empirical results are at odds with the efficient market hypothesis. Specifically, there appears to be a significant *lagged* relationship between base money growth and real estate returns. Six monthly lags were found empirically appropriate for the base money variable. While the first through fourth lagged coefficients on the base money growth are not statistically significant, the fifth and sixth lagged coefficients are highly significant at better than the 5 percent level. Table 2 reports the likelihood ratio tests of the various hypotheses in model (6). The likelihood ratio tests corroborate the finding that the base money has a significant lagged relationship with current real estate returns (χ^2 = 16.93, $\chi^2(.05)$ = 12.59, with 6 d.f.). As to the impact of market returns, the results also show a significant lagged relationship with current reale state returns. The appropriate lag profile of the market returns contains 11 monthly lags whereby the eighth and the tenth lagged coefficients are statistically significant. According to the likelihood ratio test, the coefficients on market returns are jointly significant at the 5 percent level (χ^2 = 26.29, $\chi^2(.05)$ = 19.68, with 11 d.f.).

Summing up, the regression results in table 1 and the likelihood ratio tests in table 2 provide evidence that the base money (monetary policy) and market returns have significant lagged effects on current real estate returns. Clearly, these findings seem contrary to what one might expect in an efficient real estate market since available information on monetary policy moves and market returns seem underutilized by market participants.[5]

The results raise the possibility that a diligent investor in the real estate market could reap abnormal profits by using a trading rule based on the observed behavior of the monetary base. However, we must caution the reader that such profitable opportunities may not be exploitable due to inhibiting transaction costs. Furthermore, any existing abnormal profits in excess of the transaction costs may also gradually disappear as an increasing number of investors begin to utilize available information on, for example, monetary policy moves, leading in turn to a more efficient market. Although unlikely—given the speed with which data are disseminated—it could be charged that a publication lag of more than one month is required in order for market participants to possess full knowledge of the base money and market returns figures. To check that, we reestimated the model imposing at least a two-month lag on both variables. The empirical results, not shown here becuase of space limitations, are not sensitive to the possibility of a longer publication lag and continue to indicate significant lagged impacts of both the base money and market returns upon current real estate returns.

In addition, one might further object to the simultaneous inclusion of both the monetary base and the budget deficit variables in model (6) on the grounds that the two policy variables are related in practice through the monetization process (the accommodation hypothesis). Under this situation, the two policy variables become highly collinear, thus hindering the statistical precision of the estimated coefficients. However, a growing body of empirical literature has rejected this ac-

Table 2. Likelihood ratio (LR) tests for the basic model (6)

Hypotheses[a]	Chi-Squared Statistics[b]	Degrees of Freedom (Number of Constraints)
1. $\beta_1(L) = 0$	101.96*	10
2. $\beta_2(L) = 0$	16.93*	6
3. $\beta_3(L) = 0$	26.29*	11
4. $\beta_4(L) = 0$	2.08	1
5. $\beta_5(L) = 0$	0.32	1
6. $\beta_6(L) = 0$	0.20	1
7. $\beta_7(L) = 0$	0.11	1
8. $\beta_8(L) = 0$	0.04	1

[a]The hypothesis tested are (1) lagged values of real estate portfolio returns do not Granger-cause changes in real estate portfolio returns (RER), (2) money base does not Granger-cause changes in RER, (3) returns on the value-weighted market portfolio do not Granger-cause changes in RER, (4) the unemployment rate does not Granger-cause changes in RER, (5) industrial production does not Granger-cause changes in RER, (6) the inflation rate does not Granger-cause changes in RER, (7) the federal deficit does not Granger cause changes in RER, and (8) the risk premium does not Granger-cause changes in RER.
[b]An * indicates rejection of the associated hypothesis at the 5 percent level of significance. A likelihood ratio statistic is defined as $2T\,[\log RSS^c - \log RSS^u]$, where RSS^c and RSS^u are the sums of squared residuals of the constrained and unconstrained systems, respectively, and T is the number of observations. The test statistic is asymptotically distributed as $\chi^2(q)$, where q is the number of restrictions.

commodation hypothesis for the United States and revealed no stable relationship between monetary and fiscal policy.[6] Nevertheless, we checked the sensitivity of the results to this potential problem and deleted the budget deficit variable from the model. The empirical results from this exercise, not reported here due to space limitations, indicate no material change in the evidence regarding the strong lagged impact of the base money upon current real estate returns.

Finally, it is further arguable that model (6) is overparameterized, having a total of 32 estimated parameters. Of course, this particular lag structure of the model was determined by the FPE criterion rather than chosen arbitrarily. Yet, one might argue, our lag selection procedure should be balanced by a sense of parsimony. This is particularly so when several of the explanatory variables do not exhibit strong statistical signficiance. To assess the robustness of our results to this potential problem of overparameterization, we reestimated the model deleting all insignificant variables. The results, not shown here, are, however, essentially similar to those obtained from the full model.[7]

3. Concluding remarks

This article examines empirically the causal relationship between monetary policy (and other financial variables) and U.S. real estate returns using monthly data from January 1965 to December 1986. The empirical analysis is based on multivariate Granger-causality tests in conjunction with Akaike's final prediction error (FPE) criterion. The results show that the estimated real estate return equation fits the data quite well, with white noise residuals and structurally stable parameters.

The empirical results suggest that monetary policy (measured by base money growth) and market returns play an important role in causing changes in real estate returns. More specifically, the results show that base money and market returns have had significant *lagged* effects on current real estate returns, implying a possible refutation of market efficiency. We conducted additional tests to check the sensitivity of our results to several changes in the model structure. These tests indicate the robustness of our evidence against market efficiency. Of course, such evidence would gain more credibility if supported by trading rule tests, an inquiry worthy of future research.

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Notes

1. Homa and Jaffee construct a trading rule, but their research has been critized by others (see Malkiel and Quant, 1972; and Rozeff, 1974).
2. There is considerable evidence that real estate outperforms other asset categories on a nominal and risk-adjusted basis. See for example, Friedman (1971), Robichek, Cohn, and Pringle (1972), and Hartzell, Hekman, and Miles (1986).
3. In addition to the CRSP data list, we checked the Standard and Poor's *Stock Market Encyclopedia*, Moody's *Handbook of Common Stock*, Standard and Poor's *Stock Guide*, and Valueline's *Investment Survey*. From this material we verified that each company is substantially real estate in nature.
4. If the appropriate lag is found at 12 months, the lag length is allowed to extend by at least two additional months in the FPE procedure to check whether the appropriate lag length for that variable is indeed 12 months.
5. Of course, a clear violation of the SME hypothesis necessitates trading rules tests, which are beyond the scope of this article.
6. See, for example, Joines (1985), King and Plosser (1985), Darrat and Barnhart (1989), and Hafer and Hein (1988).
7. All empirical results not reported in the article because of space limitations are available from the authors upon request.

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