U.S. Asset Returns and Fiscal Policy: New Empirical Investigation

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The persistent government budget deficit that has plagued both the United States and other countries since the 1970s invariably has attracted attention from academicians and politicians. Although the United States enjoyed a brief period of federal budget surplus from 1998 to 2001, it is once again experiencing a ballooning budget deficit, partly due to a recession in 2001, to extended bearish stock markets fueled by the dot.com bust in 2000, and to a significant boost in government spending to fight the war against terrorism. This study investigates whether the U.S. stock and corporate bond markets are informationally efficient with respect to fiscal policy (proxied by government budget deficit). The Granger causality test procedure is utilized in the context of a vector autoregression (VAR) model to investigate a dynamic relationship between excess U.S. asset returns and fiscal policy. This study finds that both stock (small-capitalization as well as large-capitalization) and bond excess returns fully captured information on fiscal and monetary policies in the United States during the period 1969-1998, a period marked by persistent budget deficit.

Introduction

The United States and other industrialized countries have been plagued by persistent, ballooning government budget deficit since the early 1970s. Although the United States briefly enjoyed a period of federal budget surplus between 1998 and 2001, it is currently experiencing a swelling budget deficit, partly due to a recession in 2001, to extended bearish stock markets triggered by the dot.com bust in 2000, and to deep income tax cuts and a significant boost in government spending to finance a war against terrorism.[1]

For the United States, the ratio of public debt (i.e., total value of government indebtedness due to sustained budget deficit) to annual GDP (gross domestic prod-

[1] Other industrialized countries such as Belgium, France, Germany and the United Kingdom had briefly experienced government budget surplus in the late 1990s and early 2000s. Like the United States, these countries are also currently mired in escalating budget deficits.

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uct) declined from over 100 percent immediately after World War II to 25 percent in 1974. The ratio rose steadily to 44.7 percent in 1998. The trend for the public debt/GDP ratio for the major OECD (Organization for Economic Cooperation and Development) countries such as Belgium, France, Germany and Japan exhibited a similar pattern during the same period—that is, falling from the late 1940s to a low point in 1973 and then steadily rising thereafter.[2]

If financial markets are informationally efficient with respect to fiscal policy,[3] then asset prices should capture any changes in fiscal policy quickly and fully once this information becomes publicly available.[4] Past information on fiscal policy would be of no use in explaining current fluctuations in security prices in an efficient market because this information has been impounded in past prices. On the other hand, in an informationally inefficient market, past information on fiscal policy should be useful in explaining current movements in security prices because there is a lag in the adjustment of security prices to new information.

Several studies (Barnhart and Darrat, 1989; Darrat, 1988 and 1990; and Darrat and Brocato, 1994) provide evidence that stock markets of the United States and Canada are not informationally efficient with respect to fiscal policy.[5] Ewing (1998) also finds that the stock markets of Australia and France are not efficient with respect to publicly available information on fiscal policy, while Lee (1997) demonstrates that stock markets of four Asian countries—Hong Kong, Singapore, South Korea and Taiwan—are not informationally efficient with respect to both monetary and fiscal policies.

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[2] Source: International Financial Statistics.

[3] The terms *budget deficit* and *fiscal policy* are used interchangeably in this study because the United States had experienced persistent federal budget deficits during the study's sample period (1969-1998).

[4] The term *informational efficiency* used in this study differs conceptually from the term *economic efficiency* used widely in the finance literature. The security market is deemed economically efficient if investors are not expected to earn an abnormal or excess rate of return from securities by acting upon the available information. It is altogether possible that informational inefficiency may lead to economic efficiency if either the relationship between variables is weak or the trading costs are not negligible. That is, even if lagged values of a variable can predict stock prices, the marginal benefit of acting on this information may not entirely offset the marginal trading cost such that excess return cannot be generated.

[5] The channel through which government budget deficit affects stock prices has been investigated extensively in the past (e.g., Blanchard, 1981). One of the fundamental tenets of financial theory is that the intrinsic value of common stock is equal to the present value of expected future dividends. Because firms pay dividends from their earnings, which in turn depend on real economic activity, stock prices should reflect current and expected future real economic activities. As macroeconomic theory posits a negative relationship between federal budget deficit and expected future economic activity (Geske and Roll, 1983), a significant, negative relationship between aggregate stock prices and fiscal policy should exist.

These findings are particularly puzzling, given that there is overwhelming empirical evidence indicating that stock markets of the United States and other industrialized countries are informationally efficient with respect to salient macro-economic variables, such as money supply, inflation rate, and real economic activity (e.g., Darrat, 1988 and 1990; Darrat and Brocato, 1994; Fama, 1981; Geske and Roll, 1983; Wasserfallen, 1989; and Jensen, Mercer, and Johnson, 1996).

The current literature is replete with studies (e.g., Fama and French, 1989; Campbell and Ammer, 1993; Jensen, Mercer, and Johnson, 1996; and Booth and Booth, 1995) asserting U.S. stock and bond excess returns move together over time because they are both affected by a common set of variables that closely track macroeconomic conditions—namely dividend yield, default spread, term spread, and money supply.[6] Accordingly, if fiscal policy proxied by federal budget deficit significantly and unambiguously affects U.S. stock returns as documented by previous studies, then one can surmise that U.S. corporate bond returns also could be significantly affected by fiscal policy.

The objective of this study is to determine whether the U.S. corporate bond market is informationally efficient with respect to fiscal policy proxied by the federal budget deficit. If fiscal policy is found to significantly affect excess U.S. corporate bond returns, then such evidence would strengthen the claim that financial assets are not informationally efficient with respect to fiscal policy. On the other hand, if excess bond returns are not affected by fiscal policy, then such a finding would diminish the credibility of the widespread assertion that there is a strong link between asset returns and fiscal policy.

This study further investigates (i) the relationship between U.S. stock (both small- and large-capitalization) market returns and fiscal policy and (ii) the relationship between both U.S. corporate bond and stock market returns and monetary policy (proxied by monetary base) in an attempt to shed more light on the dynamic relationship between U.S. financial assets returns and macroeconomic policies.

The Granger causality test procedure is utilized in the context of a vector autoregression (VAR) model to test the relationship between the U.S. asset returns and

[6] If stock prices rise due to improvements in economic condition, then dividend yield (dividend/price) and expected stock returns should immediately decline. Expected bond returns should also decline because long-term interests rise in anticipation of more robust economic activities ahead. So there should be a positive relationship between dividend yield and asset returns. The default spread is high when economic conditions are dismal and is low when economic activity remains robust. Accordingly, default spread should also move directly with expected security returns. The term spread declines near the peak of business cycle and increases near the trough of business cycle. Therefore, expected asset returns are also positively related to the term spread. Please refer to Fama and French (1988, 1989) and Fama (1990) for empirical evidence supporting these assertions.

macroeconomic policies during the 1969-1998 period, a period during which the United States had experienced a persistent, ballooning budget deficit.[7]

It is found that U.S. long-term corporate bond and stock returns fully captured the information on monetary and fiscal policies during the period 1969-1998. The evidence that U.S. stock market is informationally efficient with respect to fiscal policy is clearly at odds with the findings of Barnhart and Darrat (1989), Darrat (1988, 1990), Darrat and Brocato (1994), and Ewing (1998) that claim the stock markets of the United States and other industrialized countries are informationally inefficient with respect to fiscal policy.

Empirical Methodology

In their seminal work on macroeconometric modeling, both Hall (1978) and Nelson and Plosser (1982) detected the prevalence of non-stationary patterns in many macroeconomic variables.[8] A non-stationary process requires differencing to induce stationarity.[9] A variable that needs to be differenced "d" times to reach stationarity is referred to as an I(d) or an integrated of order d process. I(d) is also said to contain "d" unit roots.

A long-run equilibrium relationship among a set of non-stationary variables can be investigated in a rigorous manner by applying a cointegration technique introduced by Engle and Granger (1987). According to Engle and Granger, all the variables in an $n \times 1$ vector Z_t are cointegrated of order (d,b) only if all the variables in Z_t are integrated of order d and there exists an $n \times 1$ cointegrating vector, β, such that a linear relationship $\beta' Z_t$ is integrated of order (d-b).

The augmented Dickey-Fuller (ADF) test can be utilized to determine whether each of the variables in Z_t is a non-stationary process (Dickey and Fuller, 1979 and 1981). The ADF test is performed upon estimating the following regression equation for each of the variables in Z_t:

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

where Y_t is the variable under consideration. The order of k should be set large enough to ensure that the residual series, ξ_t, is a white noise process. The null hypothesis that Y_t contains a unit root is rejected when the estimated coefficient of

[7] Although the annual United States federal government budget was in surplus in 1998, the budget was in deficit for both the first and fourth quarters in 1998. Therefore, the year 1998 is included in this study's sample period.

[8] Hall (1978) has shown that aggregate consumption pattern follows a random walk process, whereas Nelson and Plosser (1982) have documented that 14 other macroeconomic variables also exhibit non-stationary behavior over time.

[9] A stationary process has time-invariant joint and conditional distributions. Consequently, the mean and variance of this process remain constant over time.

the lagged variable, ρ, in equation (1) is statistically less than zero. The null hypothesis that Y_t is I(d), where $d > 1$, can be tested in the same manner after differencing Y_t by d times.

DeJong and Whiteman (1991), Diebold and Rudebusch (1991), and Kwiatkowski, Phillips, Schmidt, and Shin (1992), however, have argued that traditional unit root tests such as the ADF procedure are neither statistically reliable nor robust in testing the null hypothesis of a unit root against several specific alternative hypotheses (such as roots near unity). In an effort to remedy such shortcomings, Kwiatkowski, Phillips, Schmidt, and Shin (1992) have devised a test of the null hypothesis of stationarity that may provide further insight into the dynamic behavior of the time-series data when used in conjunction with a standard test of the null hypothesis of unit root such as the ADF.

The Kwiatkowski, Phillips, Schmidt, and Shin test procedure is fairly straightforward. Suppose Y_t is a time-series variable to be tested for stationarity. The null hypothesis of level stationarity of Y_t is tested with the η_μ statistic where

$$\eta_\mu = N^{-2}\left(\sum_{t=1}^{N} S_t^2 / S^2(1)\right), \tag{2}$$

$S_t = \sum_{i=1}^{t} e_i$, $S^2(l) = N^{-1}\sum_{t=1}^{N} e_t^2 + 2N^{-1}\sum_{s=1}^{l} w(s,l) \sum_{t=s+1}^{N} e_t e_{t-s}$, and $w(s,l) = 1 - (s/l+1)$.

e_t, $t = 1, 2, \ldots, N$, is the residual from the regression of Y_t on only the intercept, where N is the sample size. Similarly, the null hypothesis of trend stationarity of Y_t, is tested with the η_t statistic, where η_t is defined as

$$N^{-2}\left(\sum_{t=1}^{N} S_t^2 / S^2(1)\right)$$

and e_t is the residual from the regression of Y_t on both the intercept and time trend. The critical values for these two statistics are provided in Table 1 of Kwiatkowski, Phillips, Schmidt, and Shin (1992).

If all the variables within the $n \times 1$ vector, Z_t, are found to exhibit the same order of integration, then a multivariate cointegration method such as that proposed by Johansen (1988) can be employed to uncover possible cointegrating relationships among these variables. Johansen's method involves the estimation of the following k^{th} order vector autoregressive (VAR) equation:

$$\Delta Z_t = \Gamma_1 \Delta Z_{t-1} + \ldots + \Gamma_{k-1}\Delta Z_{t-k-1} + \varepsilon_{1t} \tag{3}$$

$$Z_{t-k} = \Gamma_1 \Delta Z_{t-1} + \ldots + \Gamma_{k-1}\Delta Z_{t-k-1} + \varepsilon_{2t} \tag{4}$$

where Z_t is an $n \times 1$ vector of the variables that are integrated of same order, Γ is an $n \times n$ coefficient matrix, and ε_t is an $n \times 1$ vector of residuals. The product moment matrices of the residuals, ε_{1t} and ε_{2t}, are $S_{ij} = N^{-1}\Sigma^{N}_{t=1}\varepsilon_{it}\varepsilon'_{jt}$ for $i,j = 1,2$, where N is the sample size. The squares of the canonical correlation between ε_{1t} and ε_{2t} are ranked as $\phi_1 > \phi_2 > ... > \phi_n$, where ϕ_i, i = 1,2, ...,n, are also referred to as the eigenvalues of $S_{21}S^{-1}_{11}S_{12}$ with respect to S_{22}. Johansen introduced a likelihood ratio test statistic that can be used to test the null hypothesis of at most r cointegrating relationships:

$$\lambda_{trace}(r) = -N \sum_{j=r+1}^{n} \ln(1 - \phi_j). \quad (5)$$

where the λ_{trace} statistic is designed to test the null hypothesis against the alternative hypothesis of more than r cointegrating relationships. The limiting distribution of the λ_{trace} statistic is specified as an (n-r)-dimensional Brownian motion process. The quantiles of the distribution of the statistic are provided in Johansen and Juselius (1990, Table A3).

If there exists at least one cointegrating relationship among the variables in vector Z_t, then a causal relationship among these variables can be investigated in the context of the following vector error correction model (VECM) (Sims, Stock, and Watson, 1988)**:**

$$\Delta Z_t = \sum_{i=1}^{k-1} \Gamma_i \Delta Z_{t-i} + \alpha\beta' Z_{t-k} + \delta + \varepsilon_t \quad (6)$$

where δ is a n × 1 constant vector representing a linear trend in a system, k is a lag structure, ε_t is a n × 1 white noise residual vector, Γ is a n × n matrix describing short-term adjustments among the n variables, and α denotes the speed of adjustment.

On the other hand, if variables in the vector Z_t do not exhibit the same order of integration, then these variables are not deemed as cointegrated over time and the causal relationship among these variables should be investigated in the context of the following standard vector autoregression (VAR) model of n equations, i.e. the system of regression equation (6) without the error correction term:

$$Z_t = \alpha + A_1Z_{t-1} + A_2Z_{t-2} + ... + A_pZ_{t-p} + \varepsilon_t \quad (7)$$

where α is a n × 1 vector of constant term, A_i is a n × n matrix of coefficients for i = 1, 2, … , p, and ε_t is a n × 1 vector of serially uncorrelated white noise error terms that have a zero mean and variance-covariance matrix Σ_ε.

Data

The monthly data for corporate bond returns, large stock returns, small stock returns, dividend yield, term structures, default spread, and T-bill yield are obtained from the *Stocks, Bonds, Bills, and Inflation 1999 Yearbook* published by Ibbotson Associates. The monthly data for the U.S. budget deficit (DEFICIT) are obtained from the *International Financial Statistics* published by the International Monetary Fund (IMF). The monthly data for monetary base (MB)—proxy for monetary policy—are obtained from the FRED (Federal Reserve Economic Database), an economic time-series database maintained by the Federal Reserve Bank of St. Louis.[10]

The corporate bond returns are defined as the total return (the sum of capital appreciation return and income return) on the Salomon Brothers Long-Term High-Grade Corporate Bond Index. This index includes nearly all Aaa- and Aa-rated corporate bonds traded over the counter. The large stock returns are calculated as the total return on the Standard & Poor's 500 Composite Index. The small stock returns are defined as the total return generated by the Dimensional Fund Advisors (DFA) Small Company 9/10 Fund. This fund is a value-weighted index of the stocks in the ninth and tenth deciles of the New York Stock Exchange (NYSE) and stocks listed on the American Stock Exchange (AMEX) and over-the-counter (OTC) with the same or less capitalization as the upper bound of the NYSE ninth decile.[11]

The excess returns on corporate bonds (EXBOND), large stocks (EXLGSR), and small stocks (EXSMSR) are obtained by subtracting the total return on T-bill from the respective base returns. T-bill total return is computed as the return on a one-bill portfolio containing the shortest-term Treasury bills with not less than one month to maturity. The dividend yield series (DIVYLD) is constructed as the large stock total return less the large stock capital appreciation return. The term structures variable (TERM) is constructed as the (1 + long-term government bond total return)/(1 + treasury bill total return) – 1. The default premium variable (DEFAULT) is defined as the (1 + long–term corporate bond total return)/(1 + long-term government bond total return) – 1. Long-term government bond return is the total return on a Treasury bond portfolio with approximately 20 years to maturity.

Empirical Results

The ADF test is performed first to determine whether the aforementioned variables—i.e., EXBOND, EXLGSR, EXSMSR, DIVYLD, TERM, DEFAULT, DEFICIT, and MB—are non-stationary processes and integrated of same order. Because the results of the ADF test can be unduly sensitive to the choice of the lag length k in equation (1) (Thornton and Batten, 1985), the Akaike (1974) information criterion (AIC) is used to determine the optimal lag-structure specification of equation (1). AIC searches for the lag length k that would minimize

$$\text{AIC}(k) = N \ln(\text{SSR}(k)/N) + 2p, \tag{8}$$

where N is the number of observations, SSR(k) is the sum of squared residuals of the regression with k as the lag length, and p is the number of regressors equal to k+1.

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[10] Monetary base (currency plus reserves), rather than M1 or M2, is used in this study as a proxy for monetary policy because it accurately captures the motive of the Fed's monetary policy. The trend for either M1 or M2 depends not only on monetary base, but also on money multiplier that can be rather volatile over time.

[11] Source: *Stocks, Bonds, Bills, and Inflation*, 1999 Yearbook, Ibbotson Associates.

The appropriate order of the lag length for each variable in equation (1) is obtained by computing AIC(k) over the range of values for k from 1 to 12. Maximum lag of 12 is chosen as monthly data are used in this study. The k^{th} lag that minimizes AIC, k*, is provided in Table 1 for each variable. The ADF test statistic, t_t, on ρ in equation (1) is then obtained for all variables. Because the ADF statistic does not necessarily follow the usual student t-distribution under the null hypothesis of a unit root, the tabulated values provided in Fuller (1976) are used.

Table 1—Augmented Dickey-Fuller (ADF) Test Results 1969:1—1998:12

Y_t	k*	$t_t(k^*)$
Panel A: H_0: I(1) $\quad \Delta Y_t = \alpha + \gamma t + \rho Y_{t-1} + \sum_{i=1}^{k} \beta_i \Delta Y_{t-i} + \xi_t$		
EXBOND	9	-5.6866**
EXLGSR	12	-5.5461**
EXSMSR	1	-12.9495**
DIVYLD	1	-22.0595**
TERM	9	-5.7460**
DEFAULT	5	-8.4666**
DEFICIT	11	0.4517
MB	8	0.0627
Panel B: H_0: I(2) $\quad \Delta^2 Y_t = \alpha + \gamma t + \rho \Delta Y_{t-1} + \sum_{i=1}^{k} \beta_i \Delta^2 Y_{t-i} + \xi_t$		
DEFICIT	10	-17.0635**
MB	7	-3.5459*

EXBOND, EXLGSR, EXSMSR, DIVYLD, TERM, DEFAULT, DEFICIT, and MB, respectively, denote excess corporate bond return, excess large stock return, excess small stock return, dividend yield, term spread (long-term government bond rate – T-bill rate), default spread (long-term corporate bond return - long-term government bond return), federal budget deficit, and monetary base. Only two variables (DEFICIT and MB) were tested for second-order nonstationarities processes because the null of I(1) is not rejected for both. * and ** denote the significance at the 5 percent and 1 percent level, respectively. The critical values for t_t statistic are -3.13 (10 percent), -3.43 (5 percent), and -3.99 (1 percent) (Fuller, 1976). k*, the optimal lag-length, is determined by the Akaike information criterion

As can be seen from Panel A of Table 1, the null hypothesis of I(1) is rejected at both the 1 percent and 5 percent significance level for EXBOND, EXLGSR, EXSMSR, DIVYLD, TERM, and DEFAULT. The null hypothesis of I(1) is not rejected for DEFICIT and MB at either the 1 percent or 5 percent significance level.

Accordingly, the null hypothesis of I(2) is subsequently tested for DEFICIT and MB, the results of which are presented in Panel B of Table 1. The ADF statistic indicates that the null hypothesis of I(2) can be rejected at the 5 percent significance level for both variables. Therefore, excess bond returns, excess large stock returns, excess small stock return, dividend yield, term structures variable, and default premium for the United States can be appropriately characterized as stationary processes, whereas deficit and monetary base can be characterized as first-order, nonstationary processes.

Table 2—Kwiatkowski, Phillips, Schmidt, and Shin Test Results 1969:1—1998:12

	Lag Truncation Parameter (l)	
Y_t	l = 4	l = 8
Panel A: η_μ Test Statistic for H_0: Level Stationarity of Y_t, where $Y_t = r_0 + \varepsilon_t$		
EXBOND	.2377	.2292
EXLGSR	.4074	.4534
EXSMSR	.0995	.1283
DIVYLD	1.780**	1.1134**
TERM	.2882	.2794
DEFAULT	.1063	.1137
DEFICIT	1.1911**	.9257**
MB	6.5527**	3.6653**
Panel B: η_t Test Statistic for H_0: Stationarity of Y_t around a Deterministic Linear Trend, where $Y_t = r_0 + r_1 t + \varepsilon_t$		
EXBOND	.0799	.07794
EXLGSR	.0235	.02543
EXSMSR	.0968	.1226
DIVYLD	.5927**	.3791**
TERM	.0603	.0608
DEFAULT	.0128	.0139
DEFICIT	.2794**	.2289**
MB	1.7288**	.9707**

Note:

$$\eta_\mu = N^{-2}\left(\sum_{t=1}^{N} S_t^2 / S^2(l)\right),$$

where $S_t = \sum_{i=1}^{t} e_i$, $S^2(l) = N^{-1}\sum_{t=1}^{l} e_t^2 + 2N^{-1}\sum_{s=1}^{l} w(s,l)\sum_{t=s+1}^{N} e_t e_{t-s}$, $w(s,l) = 1 - (s/l+1)$, and $Y_t = r_0 + e_t$.

$\eta_t = N^{-2}\left(\sum_{t=1}^{N} S_t^2 / S^2(l)\right)$, when $Y_t = r_0 + r_1 t + e_t$.

N is the sample size. The critical values for η_μ are .347, .463, and .739 at the 10 percent, 5 percent, and 1 percent significance levels, respectively. The critical values for η_t are .119, .146, and .216, at the 10 percent, 5 percent, and 1 percent significance levels, respectively [Table 1, Kwiatkowski, Phillips, Schmidt, and Shin (1992)]. * and ** respectively denote the significance at the 5 percent and 1 percent level

The results of the Kwiatkowski, Phillips, Schmidt, and Shin test are presented in Table 2. Panel A of Table 2 provides estimated values of the test statistic, η_μ, a statistic specifically designed to test the null hypothesis of level stationarity against the alternative of a unit root, for all variables. Kwiatkowski, Phillips, Schmidt, and Shin warn that the value of this statistic maybe unduly sensitive to the choice of the lag truncation parameter, l, in equation (2). They suggest that l=8 should be chosen as the appropriate lag length for the test as it circumvents both large size distortions that may result under the null hypothesis when l=4 and a very low power that may result under the alternative hypothesis when l=12. The results of the Kwiatkowski, Phillips, Schmidt, and Shin test are therefore presented with both l=4 and l=8 in this study.

It can be seen from Panel A of Table 2 that the null hypothesis of level stationarity can be rejected for DIVYLD, DEFICIT and MB at the 1 percent significance level regardless of the lag length imposed on the equation (2).

Panel B of Table 2 provides estimated values of the test statistic, η_t, a statistic for testing the null hypothesis of trend stationarity against the alternative of a unit root. The empirical results are again provided with l=4 and l=8. It is found that the null hypothesis of trend stationarity can be rejected at the 1 percent significance level for DIVYLD, DEFICIT and MB regardless of the lag length.

Therefore, the Kwiatkowski, Phillips, Schmidt, and Shin test results are for most part consistent with those derived with ADF test procedure. Both tests conclude EXBOND, EXLGSR, EXSMSR, TERM, and DEFAULT are most appropriately characterized as stationary processes. DIVYLD, however, is characterized as a first-order nonstationary process according to the Kwiatkowski, Phillips, Schmidt, and Shin procedure, but as a stationary process according to the ADF test procedure. DEFICIT and MB are characterized as first-order nonstationary processes according to both the ADF and Kwiatkowski, Phillips, Schmidt, and Shin test procedures.

Because all pertinent variables are not integrated of same order, a long-run (or cointegrating) relationship among these variables may not exist. Consequently, a causal relationship between excess asset returns and macroeconomic policies should be investigated in the context of the following standard VAR model:

$$Z_t = \alpha + A_1 Z_{t-1} + A_2 Z_{t-2} + \ldots + A_p Z_{t-p} + \varepsilon_t, \tag{7}$$

where Z_t = [EXBOND, EXLGSR, EXSMSR, DIVYLD, TERM, DEFAULT, DEFICIT, MB]', α is a 8×1 vector of constant term, A_i is a 8×8 matrix of coefficients for i = 1, 2, … , p, where p is the optimal lag length of the VAR model, and ε_t is a 8×1 vector of serially uncorrelated white noise error terms that have a zero mean and variance-covariance matrix Σ_ε.

As has previously been noted, DEFICIT and MB are proxies for fiscal and monetary policy, respectively. TERM, DIVYLD, and DEFAULT variables also are included in this model because Fama and French (1989), Fama (1990), and Schwert (1990) have shown that a significant portion of variations in expected stock and bond returns can be explained by these three variables. Because DEFICIT and MB are first-order nonstationary processes, these series are differenced before the model is estimated.

The following null hypotheses are tested using the standard F-test statistic in the context of the VAR model to determine the causal relationship between excess security returns and fiscal policy:

H_0: $a_{17}(1) = a_{17}(2) = \ldots a_{17}(p) = 0$, (i.e., budget deficit does not Granger cause excess bond returns),

H_0: $a_{27}(1) = a_{27}(2) = \ldots a_{27}(p) = 0$, (i.e., budget deficit does not Granger cause excess large stock returns), and

H_0: $a_{37}(1) = a_{37}(2) = \ldots. a_{37}(p) = 0$ (i.e., budget deficit does not Granger cause excess small stock returns).

$a_{ij}(n)$ in the above null hypotheses represents the i, j^{th} coefficient of the A_n matrix in equation (7) which gauges the impact of the n^{th} lagged value of j^{th} variable on variable i. It can be seen from the ordering of variables in the Z_t vector that the first three variables in this vector are excess bond returns (EXBOND), excess large stock returns (EXLGSR), and excess small stock returns (EXSMSR), respectively. Budget deficit (DEFICIT) is the seventh variable in the Z_t vector.

If estimated coefficients of the j^{th} variable as a group are found to be statistically significant (i.e., variable j Granger causes variable i), then the past values of the variable j can explain variations in variable i. If, on the other hand, estimated coefficients of the j^{th} variable as a group is statistically insignificant (i.e., variable j does not Granger cause variable i), then the past values of the variable j are not useful in explaining current fluctuations in variable i. These results imply that stock (or bond) market informational efficiency requires fiscal policy does not Granger cause excess stock (or bond) returns, i.e., null hypothesis is not rejected.

In the same manner, the following null hypotheses can be tested to determine the causal relationship between excess security returns and money supply, which is the eighth variable in the vector Z_t:

H_0: $a_{18}(1) = a_{18}(2) = \ldots. a_{18}(p) = 0$, (i.e., money supply does not Granger cause excess bond returns),

H_0: $a_{28}(1) = a_{28}(2) = \ldots. a_{28}(p) = 0$, (i.e., money supply does not Granger cause excess large stock returns), and

H_0: $a_{38}(1) = a_{38}(2) = \ldots. a_{38}(p) = 0$ (i.e., money supply does not Granger cause excess small stock returns.

Table 3 summarizes the results of the Granger causality test. The appropriate lag length, p, of the VAR model is specified using the likelihood ratio test of the form

$$(N\text{-}p)(\log |\Sigma_r| - \log |\Sigma_u|), \quad (8)$$

where N is the number of observation, p is the number of parameters estimated in each equation of the unrestricted system, and Σ_r and Σ_u, respectively, represent the variance/covariance matrix of the restricted and unrestricted systems. This statistic has a χ^2 distribution with degrees of freedom equal to the number of restrictions in the system. The likelihood ratio statistic is computed with lag lengths ranging from 4 to 12 at the increments of 4. The optimal lag length, i.e., lag length that maximizes the likelihood ratio test statistic, is determined as 12 for the VAR model.[12]

[12] The VAR model is also estimated without restricting the lag length across the equations to (12). Hsiao's (1981) two-step search method is used to specify lag lengths. The results obtained with Hsiao's method are virtually identical to those reported in this paper. As such, empirical results seem to be robust with respect to the choice of the lag length of the VAR

Table 3—Granger Causality Test and Block Exogeneity Test Results 1969:1—1998:12

$Z_t = \alpha + A_1 Z_{t-1} + A_2 Z_{t-2} + \ldots + A_p Z_{t-p} + \varepsilon_t$ (7)

where Z_t = [EXBOND, EXLGSR, EXSMSR, DIVYLD, TERM, DEFAULT, DEFICIT, MB]'

Panel A: Granger Causality Test Results

Null Hypotheses*	F-test
H0: budget deficit does not Granger cause excess bond returns (i.e., $a_{17}(1) = a_{17}(2) = .. a_{17}(12) = 0$)	F(4,12) = .2708
H0: budget deficit does not Granger cause excess large stock returns (i.e., $a_{27}(1) = a_{27}(2) = . a_{27}(12) = 0$)	F(4,12) = .3461
H0 : budget deficit does not Granger cause excess small stock returns (i.e., $a_{37}(1) = a_{37}(2) = \ldots a_{37}(12) = 0$)	F(4,12) = 1.0389
H0: money supply does not Granger cause excess bond returns (i.e., $a_{18}(1) = a_{18}(2) = \ldots a_{18}(12) = 0$)	F(4,12) = .7886
H0: money supply does not Granger cause excess large stock returns (i.e., $a_{28}(1) = a_{28}(2) = \ldots a_{28}(12) = 0$)	F(4,12) = .3498
H0: money supply does not Granger cause excess small stock returns (i.e., $a_{38}(1) = a_{38}(2) = \ldots a_{38}(12) = 0$)	F(4,12) = .9303

* aij(p) in above null hypotheses represents the i,jth coefficient of Ape matrix in equation (7), i.e., the impact coefficient of pth lagged value of jth variable on variable i. The first three variables in the vector Z_t are excess bond returns (EXBOND), excess large stock returns (EXLGSR), and excess small stock returns (EXSMSR), respectively. Also, budget deficit (DEFICIT) and monetary base (MB) are the seventh and eighth variables in the Z_t vector, respectively. * and **, respectively, denote the significance at the 5 percent and 1 percent level

Panel B: Block Exogeneity Test Results

Null Hypotheses 8	Likelihood Ratio Test
H_o: budget deficit does not Granger cause other variables in the VAR model	$\chi^2(84) = 84.6096$
H_o: money supply does not Granger cause other variables in the VAR model	$\chi^2(84) = 62.8954$

*and ** respectively denote the significance at the 5 percent and 1 percent level

Because same lag length is imposed on all equations within the VAR model, the OLS estimation method, which yields consistent and asymptotically efficient estimates under this restriction, is used to estimate the VAR model.

As can be seen from Table 3, the null hypothesis of no Granger causality from fiscal policy to excess returns is not rejected at the 5 percent significance level for bonds, large-cap stocks, and small-cap stocks. Also the null hypothesis of no Granger causality from money supply to excess returns is not rejected at the 5 percent significance level for all three securities. These findings therefore indicate that the United States stock (large- and small-capitalization) and bond markets are informationally efficient with respect to both the fiscal and monetary policies.[13]

model. The VAR model is also estimated without differencing DEFICIT and MB series, and the results are again almost identical to those reported in this paper.

[13] Because TERM, DEFAULT, and DIVYLD track macroeconomic activities fairly accurately (Fama and French, 1988 and 1989), the VAR model has been re-estimated using the cyclically-adjusted budget deficit, i.e., the deficit adjusted for real economic activity as a proxy for

A block exogeneity test, a test procedure designed to determine whether lagged values of a given variable Granger causes other variables within the VAR model, is performed next to gain insight into the dynamic relationship between asset returns and macroeconomic policies. A block exogeneity test uses a likelihood ratio test statistic similar to that described in equation (8). This statistic has a χ^2 distribution with degrees of freedom equal to twice the number of lagged values of the variable excluded from the restricted equations.

As can be seen in Table 3, results obtained with this procedure are essentially identical to those obtained from Granger causality test. The null hypothesis that fiscal policy does not Granger cause other relevant variables in the VAR model is not rejected at the 5 percent significance level. Furthermore the null hypothesis that money supply does not Granger cause other variables in the VAR model is also not rejected at the 5 percent level.

These results therefore contradict those of Darrat (1988, 1990) and Darrat and Brocato (1994) who find that the stock markets of the United States and other countries such as Canada are informationally efficient with respect to monetary policy, but not with respect to fiscal policy.

Table 4—Variance Decomposition Analysis Results 1969:1–1998:12

Horizon	EXBOND	EXLGST	EXSMST
Panel A: Percent of Forecast Error Variance Accounted by Shocks or Innovations in Deficit			
3 Months	.802 (.023)	1.062 (.038)	.575 (.053)
6	1.496 (.024)	2.054 (.039)	1.384 (.054)
12	1.981 (.026)	2.054 (.043)	2.185 (.059)
24	2.396 (.027)	2.550 (.044)	2.995 (.061)
Panel B: Percent of Forecast Error Variance Accounted by Shocks or Innovations in Monetary Base			
3 Months	.307 (.023)	1.122 (.038)	.596 (.053)
6	.712 (.024)	1.334 (.039)	.767 (.054)
12	1.303 (.026)	2.029 (.043)	1.417(.059)
24	1.950 (.027)	3.031 (.044)	2.617(.061)

Notes: The figures in the parentheses represent the standard error of forecast of the dependent variable. EXBOND, EXLGST and EXSMST respectively represent excess corporate bond returns, excess large-cap stock returns and excess small-cap stock returns

The variance decomposition analysis is performed next to measure the impact of a shock or an innovation in macroeconomic policy on the variance of the forecast error of excess security returns. As can be seen from Table 4, a shock in deficit can account for only 2.396 percent, 2.55 percent, and 2.995 percent of the forecast error variance of EXBOND, EXLGST and EXSMST, respectively, after 24 months. Also, a shock in monetary base can explain only 1.95 percent, 3.031 percent, and 2.617

fiscal policy to circumvent the multicollinearity problem. The results remain invariant to the choice of the proxy for fiscal policy.

percent of the forecast error variance of EXBOND, EXLGST, and EXSMST, respectively, after 24 months. Therefore, shocks in either deficit or monetary base do not significantly explain variations in excess security returns, even at longer forecasting horizons. These results invariably lend support to the findings of the Granger causality tests.

The plots of impulse response functions for the three excess security returns are presented in Table 5. These graphs show the behavior of the excess returns over time in response to shocks in fiscal and monetary policies. The vertical axis in Table 5 measures the percentage change in these excess returns and the horizontal axis represents the months following the shock.

As can be seen in Table 5, after an easing of monetary policy, as proxied by monetary base, all three excess security returns decline immediately, with the decline in excess small-cap stock return being the most pronounced of the three. This behavior is expected, as an increase in monetary base triggers a drop in interest rates and therefore a drop in excess security returns. Returns for both excess large-cap and small-cap stocks turn positive in the second month whereas the return for corporate bonds continues to decline until the fifth month. Thereafter, the returns for all three securities continue to oscillate between positive and negative, although the magnitude of the volatility diminishes gradually over time.

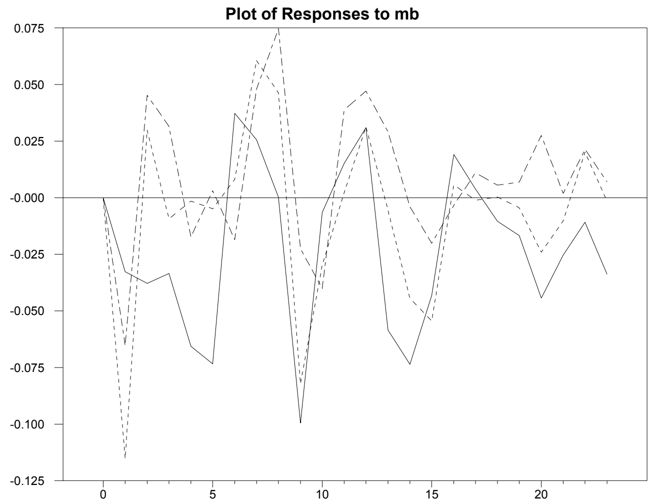
After an increase in deficit, as can be seen in Table 5, all three excess security returns increase immediately. This is also expected because a rise in the deficit induces increases in interest rates and excess security returns. But these returns quickly turn negative and then continue to oscillate between positive and negative over time. It is interesting to note that the magnitude of the volatility for excess bond returns does not seem to diminish as quickly as those of stock returns in response to deficit increase.

Conclusion

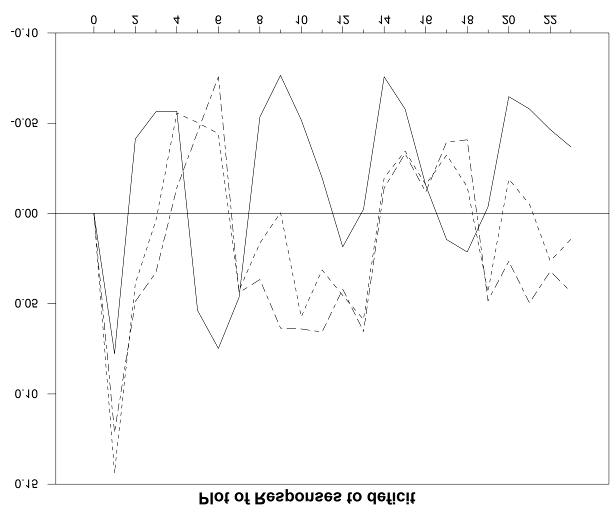
The purpose of this study was to investigate whether monetary and fiscal policies affect the movements in the U.S. corporate bond and stock (large- and small-capitalization) markets. The Granger causality test procedure is used to test the relationship between excess stock and bond returns and macroeconomic policies for the period 1969-1998, a period marked by a persistent, ballooning federal budget deficit in the United States. It is determined that U.S. stock and bond markets had fully captured and reflected publicly available information on both monetary and fiscal policies during the sample period.

Table 5—Plots of Impulse Response Function 1969:1—1998:12

Panel A: Response to a Shock in Monetary Base



Panel B: Response to a Shock in Deficit



Note: ___ EXBOND; __ _ EXLGST; - - - EXSMST.

Vertical axis represents percentage change in excess security returns and horizontal axis represents the number of month since the shock. EXBOND, EXLGST and EXSMST respectively represent excess corporate bond returns, excess large-cap stock returns and excess small-cap stock returns.

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