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REGIME CHANGES IN STOCK RETURNS

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ABSTRACT

This paper discriminates between three potential sources of instability in parameter estimates of stock return models. First, mean expected returns may vary with time. Second, return volatility may change. Third, observed returns may be affected by institutional factors as the trading mechanism evolves. To study this, we model stock returns as a stochastic function of a constant expected return and the financing costs resulting from an institutional feature, delayed delivery. We then use Goldfeld and Quandt's (1976) D-method of switching regression, deterministic switching based on time, to study the structural change in our model. We examine two eight-year sample periods and find that both contain a regime shift driven by an abrupt change in volatility. In addition, the switches occur during critical events affecting the economic environment: the first switch occurs during the turmoil of an international monetary crisis amid important Watergate developments, and the second is on the first trading day after the reappointment of Paul Volcker as chairman of the Federal Reserve Board. Although parameters estimating the impact of time-varying expected returns and the delivery system are in some cases qualitatively different between the regimes, the differences are not statistically significant and do not produce changes in our model of stock returns.

Regime Changes in Stock Returns

Changes in stock returns and in the parameters of stock return models have long been of interest to financial economists. Mehta and Beranek (1982), for example, use switching regressions to study changes in a stock's volatility (as measured by the Capital Asset Pricing Model's β coefficient) across different regimes. They find that the parameter esrimates of their model change considerably through time. More recently, Keim and Stambaugh (1986), Fama and French (1988), and Chan (1989) examine long-run expected stock returns, concluding that expected returns are cyclical and predictable. French, Schwert, and Stambaugh (1987) study the link between expected stock returns and volatility, reporting that the conditional variance of stock returns is a significant determinant of expected stock returns. Bollerslev (1987) and Baillie and Bollerslev (1989) show that conditional heteroscedasticity characterizes much financial data; this suggests that if investors are not risk-neutral and if shocks to the volatility-generating function are permanent or decay only very slowly, the process generating mean returns might also change.

This paper studies three potential sources of instability in parameter estimates of stock return models. First, expected returns may vary with time. This is consistent with Keim and Stambaugh (1986), Fama and French (1988), and Chan (1989). Second, return volatility may change, which is consistent with the conditional heteroscedasticity model of Engle (1982). Third, observed returns may be affected by changes in the institutional features of the market. Baillie and DeGennaro (1989) provide one example, demonstrating that the opportunity costs associated with delayed delivery have important effects on observed stock returns.

Perhaps surprisingly, none of these studies uses switching regressions. Goldfeld and Quandt's (1976) D-method of switching regression, deterministic switching based on time, seems especially promising. This method not only identifies switch points and estimates the model coefficients, but also provides a parameter σ_{ω}^{2}, which measures the abruptness of the change. If this parameter is statistically different from zero, the switch is interpreted as gradual. Otherwise, the switch is characterized as abrupt.

This paper applies the Goldfeld-Quandt method to a model of stock returns with three potential sources of parameter instability. We model stock returns as a stochastic function of three components: a constant expected return; the financing costs associated with delayed delivery (an institutional feature identified by Lakonishok and Levi [1982] and Flannery and Protopapadakis [1988] and tested by DeGennaro [1990] and Baillie and DeGennaro [1989]); and a rational expectations error.

Each of the three sources of parameter instability makes different predictions for the nature of parameter changes in this model. If expected returns are the only source of parameter variation, the estimate of the expected return must be the source of the regime change. The switch should be gradual, since changes in expected returns are probably slow. If financing costs or the delivery and payments mechanism is responsible for a regime change, the switch might be abrupt or gradual: regulatory change should lead to abrupt switches, while technological evolution should lead to gradual changes. Finally, it is unclear whether switches due to volatility should be abrupt or gradual. For example, one interpretation of shifts driven by changes in volatility is that they proxy for changes in omitted variables. The switching regressions method does not permit distinguishing between this and other less-ambitious interpretations, which treat volatility as exogenous. However, other researchers such as French, Schwert, and Stambaugh (1987) use a GARCH-in-mean model to show that volatility shocks are permanent. This suggests that regime shifts resulting from changes in volatility are likely to be abrupt.

We study two eight-year sample periods from 1971 to 1986 and find that both contain a regime shift driven by an abrupt change in volatility. In addition, the switches occur during important events affecting the economic environment: the first switch, driven by an sudden <u>increase</u> in volatility, occurs during the turmoil of an international monetary crisis in 1973, and the second, marked by an abrupt <u>decrease</u> in volatility, is on the first trading day after the reappointment of Paul Volcker as chairman of the Federal Reserve Board in 1983. Although the other parameters are in some cases qualitatively different, they do not vary enough to produce changes in our model of stock returns.

Our conclusion that changes in volatility are of primary importance gains force when one considers that our model never identifies switches near periods of important changes in the economy that might well have been expected to affect the other coefficients in the model. For example, stock returns were large from late 1971 through early 1973 and again after August 1982, while investors experienced low returns and even losses in portions of the early 1970s, particularly in 1974, and from the late 1970s through the middle of 1980. These large ex post changes in returns might be expected to affect the coefficients of our model; in fact, they do not. Similarly, our market proxy for the daily interest rate is the federal funds rate. The changes in Federal Reserve operating procedures on October 6, 1979 and October 9, 1982 might be expected to change the efficacy of this proxy; in fact, this parameter is never identified as the source of the change in regimes.

The paper is organized as follows. Section I develops the model of stock returns, section II outlines our method, section III describes the data and presents the results, and section IV contains our conclusions.

I. The Model

Three elements compose the model of stock returns. The first is the expected return. If the expected return is constant, the switching-regressions method will not identify a regime change driven by variation in expected returns. If the expected return does change, the method should position a switch at the appropriate point.

The second component captures a source of volatility due to an institutional feature of the stock market. Stock exchange procedures require the purchaser to deliver a bank check to the seller five business days from the date of the trade. Lakonishok and Levi (1982) note that such checks require another business day to clear, making the total payment delay six business days. Until final payment is made, the stock trade remains conditional and official title stays with the seller, who in turn cannot use the proceeds of the sale. Baillie and DeGennaro (1989) show that, although often ignored, this aspect of securities trading is an important determinant of stock returns. Since Kane and Unal (1988) interpret switches as evidence of movements in omitted variables, we control for this potential source of structural instability.

The third component of the model is a rational expectations error. Although the model's first two components control for two possible sources of structural shifts, changes in the variance of this error itself are possible. If such a change occurs, the switching regression method should identify it by positioning a switch at the point where volatility increases or decreases.

To derive the model, we first write the stock price (or the level of a stock index) at t as a function of the price at time t-1, the stock return from t-1 to t, and the dividend yield,

$$P_t = P_{t-1} * \exp(R_t - d_t), \tag{1}$$

where P_t and P_{t-1} are the observed prices at t and t-1, R_t is the return at t, and d_t is the dividend yield at t.

If buyers compensate sellers for payment delays, the prices in equation (1) diverge from prices that would be observed if delays did not exist. We call this price P'_t. The observed price at t, P_t, equals P'_t plus compensation for delayed payment. The observed price is

$$P_t = P'_t * \exp(\sum_{i=1}^{D_t} c_{i,t}), \tag{2}$$

where D_t is the number of calendar days from the trade date, t, until a check presented at settlement clears, and $c_{i,t}$ is the rate of compensation for each day i during the delay for trades made at t. If the rate of compensation is positive, P_t is greater than P'_t.

Since equation (2) holds at any t, we can also write

$$P_{t-1} = P'_{t-1} * \exp(\sum_{i=1}^{D_{t-1}} c_{i,t-1}). \tag{3}$$

Substituting (2) and (3) into (1) yields

$$P'_t * \exp(\sum_{i=1}^{D_t} c_{i,t}) = P'_{t-1} * \exp(\sum_{i=1}^{D_{t-1}} c_{i,t-1}) * \exp(R_t - d_t). \tag{4}$$

Taking natural logs and rearranging obtains

$$R_t = [\log(P_t') - \log(P_{t-1}') + d_t] + (\sum_{i=1}^{D_t} c_{i,t}) \cdot \sum_{i=1}^{D_{t-1}} c_{i,t-1}. \tag{5}$$

In equation (5), $\sum_{i=1}^{D_t} c_{i,t}$ and $\sum_{i=1}^{D_{t-1}} c_{i,t-1}$ control for differences in financing costs due to payment delays in the return R_t. At $t-1$, the observed price P_{t-1} is the value of the asset if delays do not occur, plus compensation for delayed payment. Similarly, P_t reflects the unobservable value of the asset, plus compensation for delayed payment. Unless $\sum_{i=1}^{D_t} c_{i,t}$ equals $\sum_{i=1}^{D_{t-1}} c_{i,t-1}$, the observed return misstates the actual return on the asset since it includes this change.

Some proxy for c must be used for empirical work. We use federal funds rates because they are responsive to economic conditions and are readily available. Substituting federal funds rates f for c in $\sum_{i=1}^{D_t} c_{i,t}$ and $\sum_{i=1}^{D_{t-1}} c_{i,t-1}$ and letting A indicate changes, equation (5) becomes

$$R_t = [\Delta\log(P_t') + d_t] + \Delta(\sum_{i=1}^{D_t} f_{i,t}). \tag{6}$$

The term in brackets represents the realized return in the absence of delays, and the second term controls for the delays. Writing the realized return as the expected return plus error, we obtain

$$R_t = E[\Delta\log(P_t') + d_t] + \Delta(\sum_{i=1}^{D_t} f_{i,t}) + e_t. \tag{7}$$

For notational convenience, we write $\Delta(\sum_{i=1}^{D_t} f_{i,t})$ as ΔF_t. Because these AF_t may be jointly determined with R_t, we use predicted values of ΔF_t, labeled $\hat{\Delta F}_t$, in our empirical work to avoid problems with simultaneity. Substituting this into (7) and assuming a constant expected return, the test equation is:

$$R_t = \beta_0 + \beta_1 \Delta F_t + e_t. \tag{8}$$

Since investors are risk averse and hold stocks in the expectation of earning a positive return, β_0 should be positive and a one-tailed test is appropriate. Similarly, if buyers compensate sellers for payment delays, β_1 is positive and a one-tailed test is again appropriate.

II. The Switching Regression Method

The switching regression technique we use was introduced by Goldfeld and Quandt (1973, 1976). This technique allows the data to identify the switch point and provides information about the type of switch (abrupt or gradual). Lin and Oh (1984) use this technique to test the stability of the U.S. short-run money demand function, while Kane and Unal (1988) use it to study changes in the market's perception of risk in the stock of banks and savings and loans. The two-regime stock return is described as follows:

$$R_t = \beta_0^a + \beta_1^a \Delta\hat{F}_t + e_t^a \qquad t = 1, 2, \ldots, t^* \tag{9}$$

$$R_t = \beta_0^b + \beta_1^b \Delta\hat{F}_t + e_t^b \qquad t = (t^*+1) \ldots, N \tag{10}$$

where

$$e_t^a = \rho^a e_{t-1}^a + u_t^a$$
$$e_t^b = \rho^b e_{t-1}^b + u_t^b,$$

and the other variables are as defined previously, with the superscripts a and b denoting the regime index. The autoregressive structure is necessary because we use a portfolio return as the dependent variable. Although Scholes and Williams (1977) show that a moving-average parameterization is strictly correct in this case, higher-order autocorrelations approach zero very rapidly for small values of ρ. In addition, the autoregressive structure is convenient for computation.

The regime change is assumed to be time-dependent. The two regimes may be combined by introducing a dummy variable D as follows:

$$R_t = (1 - D_t)(\beta_0^a + \beta_1^a \Delta\hat{F}_t + \rho^a e_{t-1}^a + u_t^a)$$
$$+ D_t(\beta_0^b + \beta_1^b \Delta\hat{F}_t + \rho^b e_{t-1}^b + u_t^b), \qquad (11)$$

where D_t indicates the probability of a specific regime for each observation t. If the regime change occurs abruptly, then

$$D_t \begin{cases} = 0, \text{ if } t \leq t^* \\ = 1, \text{ otherwise.} \end{cases}$$

However, if the regime change is gradual, the dummy variable may be approximated by a continuous function that increases gradually from zero to one for observations two through N. One approximation suggested by Goldfeld and Quandt (1976) is the normal distribution with two parameters, t^* and σ_w^2:

$$D_t = \int_{-\infty}^{t} [(2\pi)^{1/2}\sigma_\omega]^{-1} * \exp\{- 1/2[\frac{\delta - t^*}{\sigma_w}]^2\}d\delta, \qquad (12)$$

where t^* indicates the central point of the switch and σ_ω^2 characterizes the length of the switching period. The switch is gradual if σ_ω^2 is significantly different from zero.

When a regime change occurs at t^*, it is likely that the first-order autoregressive error parameters will also change. Therefore, this change must be built into the log-likelihood function. Assuming that u_t^a and u_t^b are independently and normally distributed with zero mean and variance $\sigma^{2(a)}$ and $\sigma^{2(b)}$, the variance of the combined error is

$$\sigma_{rt}^2 = (1 - D_t)^2[(1 - D_{t-1}) + D_{t-1} / (1 - (\rho^a)^2)^{1/2}]^2\sigma^{2(a)}$$
$$+ D_t^2[D_{t-1} + (1 - D_{t-1}) / (1 - (\rho^b)^2)^{1/2}]^2\sigma^{2(b)} \qquad t = 2, \ldots N. \qquad (13)$$

Therefore, the log-likelihood function is

$$\log(1) = - (\frac{N-1}{2})(\log 2\pi) - \frac{1}{2}\sum_{i=1}^{N}\log[\sigma_{rt}^2 - \frac{1}{2}\sum_{i=2}^{N}[\frac{(R_t - \hat{R}_t)^2}{\sigma_{rt}^2}], \qquad (14)$$

where N equals the number of observations and

$$\hat{R}_t = (1-D_t)[\beta_0^a + \beta_1^a\Delta\hat{F}_t + (1 - D_{t-1})\rho^a(R_{t-1} - \beta_0^a - \beta_1^a\Delta\hat{F}_{t-1})] + D_t[\beta_0^b + \beta_1^b\Delta\hat{F}_t + D_{t-1}\rho^b(R_{t-1} - \beta_0^b - \beta_1^b\Delta\hat{F}_{t-1})]. \quad (15)$$

We maximize equation (14) with respect to the model coefficients, $\sigma^{2(a)}$, $\sigma^{2(b)}$, ρ^a, ρ^b, t^*, and σ^2_ω using the GQOPT software package written by Goldfeld, Quandt, and Ertel. We use the GRADX routine to search for the maximum. To avoid mistaking a local maximum for the global maximum, we use several different sets of initial estimates for each sample period with the convergence criterion set at 10^{-10}.

III. Data and Results

A. Data

The proxy for the continuously compounded stock return is the natural logarithm of one plus the return on the value-weighted portfolio, including dividends, provided by the Center for Research in Security Prices (CRSP) at the University of Chicago. We use 16 years of daily data, from 1971 through 1986, a total of 4,042 observations. Federal funds rates are from the Federal Reserve Board. Predicted values of ΔF_t are obtained by regressing ΔF_t on the five most recently observed values of ΔF available at t. We divide our sample into two eight-year samples (containing 2,019 and 2,023 observations, respectively) for two reasons. First, the computational demands of our method are heavy. Second, we find switches in both eight-year samples. Applying our method to the full sample must, therefore, misspecify the dimension of our model: we would need at least two breakpoints to adequately describe the data for the full 16 years.

B. Results

First Eight-Year Sample. This sample extends from January 1, 1970 through December 31, 1978. We first determine if a switch exists in this sample. To do this, we use two methods. The first follows Quandt (1958): we estimate the model with no switches and again with one switch. We then conduct a likelihood ratio test. Twice the difference in the log-likelihoods is distributed chi-square with degrees of freedom equal to the number of restrictions implied by the null hypothesis, which is six in this case.' The second method is due to Schwarz (1978). Schwarz defines N as the total number of observations and k_r as the number of parameters that must be estimated in the r regimes. He proves that subtracting $.5k_r\log(N)$ from the maximum of the log-likelihood provides asymptotically optimal estimates.

The results are contained in table 1. The value of the log-likelihood for the no-switch case is 6858.6, while for the one-switch case, the value is 6938.6. Twice the difference in these values easily exceeds the 1 percent critical value of 16.81, so the likelihood-ratio test rejects the model with no switches in favor of at least one switch in the first eight-year sample. The Schwarz criterion agrees.

The model positions the switch on Wednesday, March 14, 1973. Consistent with the model, β_0 and β_1 are positive in both regimes. The stock returns implied by the intercepts are close to the actual stock returns during the sample. The estimated value of β_0^a (before the switch) implies a return of 12.51 percent, while the realized value was 12.14 percent. The rate implied by β_0^b (after the switch) is 2.11 percent versus the realized return of 2.13 percent. The coefficients on the payment delay,

β_1^a and β_1^b, are positive and statistically significant. Consistent with Lakonishok and Levi (1982) and DeGennaro (1990), both exceed unity, suggesting that the rate of compensation for delays is greater than the federal funds rate.

The parameter σ_ω^2 is of special interest. Goldfeld and Quandt (1976) interpret it as the degree of discrimination between regimes, or the "mushiness" of the switch. Here, its estimated value is 4.90 days, and its t-ratio is only 0.56. We cannot reject the hypothesis that σ_ω^2 is zero, so we conclude that the switch was abrupt.

To determine which parameter is responsible for the regime change, we conduct t-tests on each of the four pairs of parameters. These t-tests reveal that the regime change is driven by $a^{2(a)}$ and $\sigma,^{2(b)}$ the residual variances. The t-value is -14.05, which is easily significant at the 1 percent level. In contrast, despite the 31 percent decline in the intercept, the t-value for β_0^a and β_0^b is only 0.88, which is not significant. The standard errors of the parameters are too large to permit the model to attribute a change to these parameters. The t-values for the other parameters are also insignificant: for β_1 the t-value is only 1.28, and for ρ it is 1.37.

Although Kane and Unal (1988) caution against attributing regime changes to a specific event, we believe it is worthwhile to make such an attempt. Kane and Unal use monthly data and obtain σ_ω as large as nine months. As they note, several events typically occur during such extended periods. In contrast, our estimated a_ω is less than three days, greatly reducing the number of events that can occur. Nevertheless, we offer the evidence below as suggestive rather than conclusive.

To determine whether the economy was subject to any economic shocks during the week surrounding the Wednesday, March 14, 1973 switch point, we examined the Wall Street Journal for evidence of unusual events. We found several. The Thursday, March 8 edition reported that Arthur Burns, then chairman of the Federal Reserve Board, told Congress that the "...task of overhauling the international monetary system must be done in a matter of months rather than years." The report concluded that his statements "...indicated a new level of urgency." The following Monday, the penultimate day of the first regime, the Journal reported that six Common Market countries agreed jointly to float their currencies against the U.S. dollar. Further, the Gaullists retained their majority in the French National Assembly, which was widely viewed as making French participation in the joint float possible. The Treasury devalued the U.S. dollar that evening. Consistent with the empirical result that the completion of the regime change took about a week, the United States did not agree to participate in the plan at this time, promising only to meet that Friday.

Two events immediately after the selected break point may also have added uncertainty to the markets and contributed to the increase in volatility that motivated the switch. First, on March 14, the morning of the first day of the second structural regime, the Journal carried a front-page story regarding the Watergate proceedings. While such stories were common at the time, this article reported the Senate Judiciary Committee's "direct challenge" to then-President Richard M. Nixon. The president had blocked an aide from appearing before the committee, but capitulated on March 13, the last day of the first structural regime. This cleared the way for new and likely damaging testimony.

The second event was reported **March** 16. In a lengthy article on page three, the Journal reported that the Securities and Exchange Commission planned major changes in the central market structure, aimed primarily at "...eliminating many of the competitive differences between the nation's stock exchanges and non-exchange markets." Certain anticipated changes were substantial. For example, the Third Market, a network of securities dealers who trade away from the floor of the exchanges, would be required to meet the same obligations as the specialists on the exchanges. Brokers for large trades, called "block positioners," would be forced to break the block to fill limit orders with the specialist. The proposed regulations would also impose uniform and more stringent capital requirements for specialists operating on regional exchanges. Although the SEC did not expect the new requirements to become effective for at least two years, the scope of the changes may well have added to the uncertainty surrounding the events of the week. In contrast, we find no evidence of events that might have caused changes in expected stock returns, delivery terms, or our federal funds proxy for the opportunity cost those terms entail.

Without making expected stock returns endogenous, it is hard to say what event might cause a change in expected returns. However, Chan (1989) conjectures that output shocks might be one factor; we find no evidence of a substantial shock in that area. This does not imply that output shocks have no effect on expected returns. In fact, because the change in our model is driven by a change in a factor other than expected returns, Chan predicts we would **not** find news of any obvious output shock during that period. However, we do uncover events supporting the result that changes in market volatility drive the structural shift. We conclude that the evidence provides strong support for the hypothesis that the source of structural change in our model is due to market volatility. We find little or no evidence, either in the empirical results or in the print media, to support the hypotheses that expected returns or rates of compensation during payment delays cause regime changes.

Second Eight-Year Sample. Table 2 contains the results from the second eight-year sample, which extends from January 1, 1979 to December 31, 1986. In general, the results are similar to the first eight-year sample. Maximizing equation (14) for the no-switch case obtains a log-likelihood of 6831.1. Under the one-switch model, the value is 6861.8. Twice the difference of the log-likelihood is again distributed chi-square with six degrees of freedom, and again it easily exceeds the 1 percent critical value of 16.81. As is true in the first eight-year sample, the Schwarz criterion also rejects the model with no switch, and we conclude that at least one switch occurs in the second eight-year sample.

The data suggest the switch occurs on Monday, June 20, 1983. All parameter estimates are consistent with the model: β_0 and β_1 are positive and significant both before and after the switch. As would be expected given the larger ex post returns on stocks during the second eight-year sample, both β_0 coefficients are larger than in table 1. The coefficients imply rates of return of 19.24 percent and 13.24 percent before and after the change, versus the CRSP realized returns of 19.10 and 13.11 percent, respectively. The coefficients on the variables that control for payment delays, β_1^a and β_1^b, are also correctly signed and significantly different from zero. Since neither differs from unity, we conclude that the rate of compensation for delays is approximately equal to the federal funds rate.

The degree of abruptness parameter, σ_{ω}^{2}, is 6.72, and the t-ratio is 0.62: we cannot reject the hypothesis that the change was abrupt.

T-tests clearly show that the change was driven by the volatility parameters, $a^{2(a)}$ and $\sigma_{,}^{2(b)}$ which decline from about 8.2×10^{-5} to about. 5.3×10^{-5}. The t-value is 7.48, which is significant at the 1 percent level. For the intercept, slope, and autoregressive parameters, the comparable statistics are only 0.56, 1.08, and -0.18, respectively, none of which approaches significance. We again conclude that volatility is of critical importance to the model.

Economic events at the time of the switch provide more evidence in support of our conclusion that volatility is the source of the regime change. On Saturday, June 18, 1983, President Ronald Reagan reappointed Paul Volcker as Chairman of the Federal Reserve Board. In its Monday edition, the Wall Street Journal reported this "... ended months of guessing about whether the president would retain Volcker ...." Further, the consensus opinion was that this would help stabilize the economy. For example, the chairman of a $10 billion trust company called Volcker's reappointment "... an incredibly positive move ...." We believe the model's ability to determine a switch on the first day that economic agents could respond to the reappointment is compelling evidence in favor of the model. It also highlights the influence such major political figures wield over the volatility of financial markets.

Given the intuitively pleasing result that a major economic event coincides with the selected switch date, it is perhaps surprising that σ_{w}^{2} is as large as 6.72, even though it is statistically zero. Certain other events during the period, however, suggest that another long-term disruption in the economy at this time might have been feared, thereby increasing volatility and counteracting the calming influence of the Volcker reappointment. On Thursday, June 16, 1983, the Journal reported that, "... in a major and unexpected decision ...," the Washington State Court had freed Washington utilities from their obligation to pay $2.25 billion of debt on canceled power plants, dramatically increasing the likelihood of a default by the largest municipal bond issuer in the nation. James Durham, vice president and senior counsel for another utility, was quoted as saying this implies "... commitments made in good faith can be dishonored by government bodies. Apparently nobody's word is good for anything anymore--even if it's in writing." The next day's edition carried a major story reporting that the probable default was raising interest costs for all large borrowers in the Northwest, not just for utilities, and that the North Carolina Municipal Power Agency had postponed a $350 million offering indefinitely as a result of the ruling. This court ruling may have increased uncertainty, which contaminated the economic climate around the Volcker reappointment. This might explain why σ_{ω}^{2} is not closer to zero

Other Dates. Also of interest are dates that might plausibly be expected to cause a change in the structure of our model, but which are not selected. For example, Rogalski (1984) selects October 1, 1974--the day the New York Stock Exchange extended trading an additional 30 minutes--as the beginning of his sample period. He may suspect that this altered the daily pattern of returns. Our results, however, suggest researchers need not be concerned with this non-event. Also not selected is February 8, 1980, when the Federal Reserve moved its money supply announcements from Thursday afternoon to Friday afternoon. Cornell (1985) tests whether the intraweek pricing pattern changed after the introduction of stock-index futures trading on April 21, 1982. He concludes it did not, and our results support his. Although all of these changes might have affected β_0 or σ^2, apparently none did so to the extent that other events affected $a^{2(a)}$ and $\sigma^{2(b)}$. We add that although the introduction of futures trading is sometimes blamed for increasing stock return volatility, we find nothing to support this claim. Indeed, political activity appears to be far more important.

Other events could conceivably have affected β_1. For example, in 1977 brokerage houses began offering cash management accounts. These accounts became widespread by about 1979, and might have altered the time between stock transactions and the crediting of accounts. Congress passed the Depository Institutions Deregulation and Monetary Control Act of 1980, creating incentives to invest in faster check-processing technologies. This would reduce the delay and possibly cause the structural change. Since we use the federal funds rate as our proxy for the financing costs during the payment delays encountered in stock transactions, one might have expected October 6, 1979 to have been selected. On that date, the Federal Reserve began targeting the level of nonborrowed reserves rather than the federal funds rate. After this decision, the federal funds rate is known to have become more volatile. This could conceivably have affected β_1^a and β_1^b. On October 9, 1982, the Fed began attempting to stabilize rates. This, too, might be expected to have caused a change. Yet, none of these dates are selected. The data continually indicate that market volatility is the factor determining structural change in our model of stock returns. We find no evidence to support the hypotheses that changes in expected returns, rates of compensation for delays, our proxy for interest costs, or the autoregressive parameters contribute to structural changes in our model.

A Sensitivity Test

As reported by Baillie and DeGennaro (1989) and DeGennaro (1990), the variable that controls for the opportunity costs associated with delivery procedures is always significant. As a sensitivity check, we estimate equations (9) and (10) without including this variable. The results are not substantially affected. A regime change is identified in each eight-year subperiod, although not at the same point. For the first eight-year period, the break occurs on February 21, 1973, three weeks earlier than in table 1 For the second eight years, the switch is placed on March 15, 1983, three months sooner than in table 2. No other differences are apparent. For example, as is true for equations (9) and (10), the regime change is driven by shifts in σ^2 in both subperiods. Because likelihood-ratio tests reveal that the models in tables 1 and 2 are preferred to this simple model, we do not report the simpler model in tabular form, but the results are available on request.

IV. Conclusions

This paper models stock returns as a function of three components: a constant expected return, the impact of the mechanism for executing trades, and a rational expectations error. We examine changes in these parameters using Goldfeld and Quandt's (1976) deterministic switching based on time. This method not only allows us to learn if and when the regression structure changes, but also provides a measure of the speed of transition from one regime to the other. We find that, regardless of the sample period, all regime shifts are due to changes in the estimated variance of the error.

This is true even if the ex post return on the stock portfolio or the estimated rate of compensation for financing costs changes substantially. In addition, these changes occur during substantial changes in the business environment, driven by important political decisions. We interpret these findings as suggesting that government policy strongly affects the volatility of the stock market.

Footnotes

1. The null hypothesis of no switch restricts β_0, β_1, ρ, and σ to be equal in both regimes, as well as restricting $D = 0$ and $\sigma_\omega = 0$. These last two effectively say that we need not estimate either the location of the switch or its variance in the one-regime model.

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TABLE 1

Regressions of the rate of return on the CRSP value-weighted index, including dividends, on the predicted change in the federal funds rates during payment delays.

First Eight-Year Subperiod

$$R_t = \beta_0 + \beta_1 \Delta\hat{F}_t + e_t, \qquad (8)$$

$$e_t = \rho e_{t-1} + u_t.$$

R_t = return on CRSP value-weighted index, including dividends.

ΔF_t = predicted change in the proxy for financing costs, the total return on federal funds during payment delays at t.

Parameter	Jan. 5, 1970 - Dec. 31, 1978 Estimate (t-statistic)	Jan. 5, 1970 - March 13, 1973 March 14, 1973 - Dec. 31, 1978 Estimate (t-statistic)
β_0^a	1.98×10^{-4} (0.80)	4.97×10^{-4} $(1.42)^b$
β_1^a	1.76 $(2.77)^a$	2.87 $(3.42)^a$
ρ^a	0.27 $(12.7)^a$	.33 $(8.20)^a$
$\sigma^{2(a)}$	6.53×10^{-5} $(31.8)^a$	3.03×10^{-5} $(16.6)^a$
β_0^b	-----	8.38×10^{-5} (0.27)
β_1^b	-----	1.40 $(1.73)^c$
ρ^b		0.26 $(10.5)^a$
$\sigma^{2(b)}$	-----	7.85×10^{-5} $(27.0)^a$
σ_ω^2	-----	4.90 (0.56)
Log-likelihood	6858.6	6938.6
Schwarz Criterion	6843.4	6900.5

Note: All except ρ and σ_w^2 are one-tailed tests.

a. Significant at the 1 percent level.
b. Significant at the 10 percent level.
c. Significant at the 5 percent level.

Source: Authors' computations.

TABLE 2

Regressions of the rate of return on the CRSP value-weighted index, including dividends, on the predicted change in the federal funds rates during payment delays.
Second Eight-Year Subperiod

$$R_t = \beta_0 + \beta_1 \Delta F_t + e_t, \tag{8}$$

$$e_t = \rho e_{t-1} + u_t.$$

R_t = return on CRSP value-weighted index, including dividends.

ΔF_t = predicted change in the proxy for financing costs, the total return on federal funds during payment delays at t.

	Jan. 1, 1979 - Dec. 31, 1986	Jan. 1, 1979 - June 19, 1983 June 20, 1983 - Dec. 31, 1986
Parameter	Estimate (t-statistic)	Estimate (t-statistic)
β_0^a	$6.49\text{X}10^{-10}$ $(3.00)^a$	$7.63\text{X}10^{-4}$ $(2.41)^a$
β_1^a	1.70 $(4.09)^a$	2.06 $(3.45)^a$
ρ^a	0.15 $(6.89)^a$	.15 $(5.06)^a$
$\sigma^{2(a)}$	$6.80\text{X}10^{-5}$ $(31.8)^a$	$8.20\text{X}10^{-5}$ $(23.7)^a$
β_0^b	-----	$5.25\text{X}10^{-4}$ $(1.86)^b$
β_1^b	-----	1.19 $(2.14)^b$
ρ^b	----	0.16 $(4.75)^a$
$\sigma^{2(b)}$	-----	$5.05\text{X}10^{-5}$ $(21.1)^a$
σ_ω^2	-----	6.72 (0.62)
Log-likelihood	6831.1	6861.8
Schwarz Criterion	6815.9	6823.8

Note: All except ρ and σ_w^2 are one-tailed tests.

a. Significant at the 1 percent level.
b. Significant at the 5 percent level.

Source: Authors' computations.