Stock Price and Risk Related to Accounting Changes in Inventory Valuation

Shyam Sunder

During an inflationary period, changes to the Last In, First Out (LIFO) method of inventory valuation generally result in reduction of reported earnings and in deferment of tax payments. If the investors rely on the reported earnings, the stock price of the firms which change to the LIFO method will decrease; if they rely on the economic value of the firms, the stock price will increase. Several studies (Kaplan and Roll, 1972; Archibald, 1968) of the relationship between accounting changes and stock price behavior have been conducted by using a research design proposed by Fama, Fisher, Jensen and Roll (1969). The design involves the use of the market model to isolate the stock price changes associated with specific events from the market-wide price changes. It has been shown by Sunder (1973) that: (a) a possibility exists that the changes in accounting for inventory valuation may be associated with changes in the relative risk of stocks, and (b) in the presence of risk changes, application of Fama et al.'s research design which assumes that the relative risk of the firms involved is constant may yield misleading results.

The present study is an attempt to measure the association between the accounting and price changes by abstracting the effect of risk changes. This is accomplished by estimating the time path of the relative risk of stocks during the months surrounding the date of accounting change. The problem of estimating the relative risk of stocks when it is not constant is considered in the next section. The use of Cooley and Prescott's (1972, 1973a, 1973b) adaptive regression model is proposed for the estimation of risk, and this estimation procedure is applied to stock price data of the firms which made accounting changes to or from the LIFO method. Conclusions of the study about the relationship between stock price behavior and accounting changes are presented in the last section.

Estimation of Relative Risk in a Changing Environment

Random Walk as a Description of the Behavior of Risk

Relative risk of a firm is determined by the nature of its assets, business environ-

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ment, and future prospects. Since these factors do not change very frequently, the risk of a firm measured at regular intervals is likely to be highly autocorrelated.

Bogue (1972) proposed a theory for the behavior of relative risk based on continual depletion of old assets and the addition of new assets to the firm. He theorized a first order autoregressive process for the risk parameter β of which random walk is a special case. If β_t is relative risk in period t, β_{t-1} is relative risk in period $t-1$, ρ is the autocorrelation coefficient, and u_t is an identically and independently distributed random variable, the first order autoregressive process is characterized by (1):

$$\beta_t = \rho\beta_{t-1} + u_t. \qquad (1)$$

If the autoregressive parameter, ρ, is unity, this process is reduced to a random walk:

$$\beta_t = \beta_{t-1} + u_t \qquad (2)$$

Sunder (1973b) provided some evidence that if the relative risk of firms is assumed to follow the process defined by (1), then the autocorrelation coefficient is, indeed, quite close to unity. Fisher (1970) and Bogue (1972) provided further empirical support to the view that the behavior of risk can be closely approximated by a random walk. Therefore the random walk model (2) of relative risk is used in the remainder of this study.

Specification

This subsection presents a brief review of the work in the engineering and econometrics literature which is directly applicable to the problem of estimating relative risk of firms in a changing environment. The market model is written in the following form:

$$y_t = \beta_{1t} + \beta_{2t}x_t \quad (t = 1, \cdots, T) \quad (3)$$

where y_t, the log-return on a stock in period t is the dependent variable; β_{1t} is the random intercept in period t; β_{2t} is the relative risk in period t; and x_t, the log-return on the market index, is the independent variable. There are T observations indexed by $(t=1, \cdots, T)$. Conventionally, an estimate of the coefficient β_{2t} is called a smoothed estimate, a current or filtered estimate, or a forecast, depending on whether t is less than, equal to, or greater than T. For the purpose of tracing the time path of relative risk in this study, the primary concern is with the smoothed estimates. For example, to make an estimate of the risk of a firm in January 1950, there is no reason why data from the months before as well as after this date (which is now available) should not be used to gain efficiency of estimation.

The problem of tracing the path of relative risk of a firm is analogous to the engineering problem of tracing the path of a rocket with the help of radar signals. Since no single signal is completely accurate and the rocket may change its direction at any time because of a large number of factors, its position is continually recalculated from the past and current radar signals. In such a system, later signals are given greater weight than the earlier signals. The ordinary least square procedure is comparable to a system in which all signals are given equal weight irrespective of the time they are received. In order to obtain optimal estimates of relative risk, a weighting scheme for the observations must be devised. A frequently used criterion of optimality is minimizing the variance of estimate.

Exponential smoothing is a scheme in which weights of observations decrease exponentially with the distance from the point of estimation. This procedure has been shown to be optimal for very large (theoretically infinite) number of observations (see Muth, 1960). For a finite number of observations, exponential smoothing is only a good approximation. Ad hoc procedures for applying exponential smoothing to finite samples are given by Holt, Modigliani, Muth and Simon (1963) and Wade (1967).

In considering the current estimation problem for a general, discrete, dynamic system, Kalman (1960) provided an optimal weighting scheme (filtering scheme) for a finite number of observations.[1] Rosenberg (1968) extended Kalman's results to smoothing and forecasting. Fisher (1970, 1971a) applied these results to the market model. Fisher assumed that: (a) the intercept term β_{1t}, in the market model (3) is always zero and (b) the relative risk parameter, β_{2t}, follows a random walk,

$$\beta_{2t} = \beta_{2,t-1} + u_t \qquad (4)$$

where u_t is an identically and independently distributed random variable with zero mean. Under these two assumptions, he obtained the minimum variance linear unbiased estimate, b_t, of β_{2t}:

$$b_t = \sum_{j=1}^{t} \phi_j x_j y_j \Big/ \sum_{j=1}^{t} \phi_j x_j^2 \qquad (5)$$

$$\phi_{t+1} = \phi_t + \frac{\omega^2}{\sigma_u^2} \sum_{h=1}^{t} x_h^2 \phi_h \qquad (6)$$

where ω^2 is the step variance of changes in relative risk and σ_u^2 is the variance of the intercept term, β_{1t}, in the market model (3). ϕ_t's are the weights assigned to individual observations.[2] The procedure for applying this scheme is to put ϕ_1 (or ϕ_T) equal to any positive number, say 1, and work forward (or backward) with the observations x_t and the variance ratio (ω^2/σ_u^2). Fisher does not give a procedure for estimating the variance ratio. He assumes that this ratio is equal for all stocks and conducts estimation with several assumed values of this ratio.

In contrast to Kalman, Rosenberg and Fisher's work, Cooley and Prescott (1972, 1973a, 1973b) view this problem in the framework of a regression model.[3] Their model is more general than those considered above in the sense that it allows for both permanent and transitory changes in parameters of the linear model. Cooley and Prescott's formulation is more suitable for the problem at hand and is discussed next.

Market model (3) can be rewritten as

$$y_t = x_t'\beta_t \quad (t = 1, \cdots, T) \qquad (7)$$

where x_t is a two-component column vector of explanatory variables, the first element being 1 and the second being the return on the market factor. β_t is a two-component column vector of coefficients β_{1t} and β_{2t} in period t. The first component represents the random intercept, and the second is the risk coefficient (relative risk) of the market model. Vector β_t can change from one period to another. Changes either can be transitory, which last only for a single period, or can be permanent, which persist into the future periods once they have occurred. Denote the two-component vector of transient parameter change in period t by u_t. Then the realized value of the parameter in period t can be written as the sum of a permanent component, w_t, and the transient change, u_t.

[1] The problem considered by Kalman can be described as follows: Observable variables y_t and $(x_{0t}, x_{1t}, \cdots, x_{nt})$ are related by:

$$y_t = \beta_{0t}x_{0t} + \beta_{1t}x_{1t} + \cdots + \beta_{nt}x_{nt}$$

or, in vector notation,

$$\underset{(p \times 1)}{y_t} = \underset{(p \times n)}{x_t} \; \underset{(n \times 1)}{\beta}$$

Parameter vector β follows the linear dynamic model

$$\beta(t+1) = \phi(t+1; t)\beta(t) + u(t)$$

where $u(t)$ is an independent Gaussian n-vector with zero mean, $\beta(t)$ is an n-vector, $y(t)$ is a p-vector, and $\phi(t+1; t)$ is the $(n \times n)$ transition matrix.

[2] It might be interesting to compare (5) with the corresponding OLS estimator under a similar set of assumptions.

$$b_{t(OLS)} = \sum_{j=1}^{t} x_j y_j \Big/ \sum_{j=1}^{t} x_j^2$$

which implies $\phi_1 = \phi_2 = \cdots = \phi_T$.

[3] A good reference on the relationship of recursive and regression approaches to the problem is Duncan and Horn (1972).

$$\beta_t = w_t + u_t \quad (t = 1, \cdots, T) \tag{8}$$

Permanent component, w_t, of the parameter vector differs from the permanent component, w_{t-1}, in period $t-1$ by the permanent change vector v_t:

$$w_t = w_{t-1} + v_t \quad (t = 2, \cdots, T) \tag{9}$$

u_t and v_t are assumed to have a bivariate normal distribution. It is also assumed that each vector has the same distribution in each period, and is intertemporally independent of itself and of each other. Let Σ^*u and Σ^*v be the covariance matrices of u_t and v_t respectively. Then,

$$u_t \sim N_2(\emptyset, \Sigma^*u) \tag{10}$$

$$v_t \sim N_2(\emptyset, \Sigma^*v) \tag{11}$$

$$\begin{aligned} E(u_t u_s') &= E(v_t v_s') = 0 \quad \text{for } t \neq s \\ E(u_t v_t') &= 0 \quad \text{for all } t \end{aligned} \tag{12}$$

Since the covariances matrices Σ^*u and Σ^*v, are not independently identifiable, Cooley and Prescott suggest the following reparametrization:

$$\begin{aligned} \Sigma^*u &\equiv (1-\gamma)\sigma^2\Sigma u \\ \Sigma^*v &\equiv \gamma\sigma^2\Sigma v \end{aligned} \tag{13}$$

where covariance matrices, Σu and Σv, are known up to a scale factor. In other words, the degree of instability in each parameter relative to the other has to be prespecified. For the market model, an appropriate structure for these matrices is suggested in the next paragraph. σ^2 is the scaling parameter to be estimated. $0 \leq \gamma \leq 1$ defines the partition of the total variance between permanent and transient changes.

The reparameterization procedure will be clear from the following discussion of its application to the market model. First, Σu and Σv are written as

$$\begin{aligned} \Sigma u &= \begin{bmatrix} s_{11} & s_{12} \\ s_{21} & s_{22} \end{bmatrix} \\ \Sigma v &= \begin{bmatrix} r_{11} & r_{12} \\ r_{21} & r_{22} \end{bmatrix} \end{aligned} \tag{14}$$

Since these matrices are defined only up to a scale factor, it does not matter if all their elements are divided or multiplied by the same nonzero number. Corresponding estimates of σ^2 and γ from the procedure defined later will be adjusted accordingly. The transient variance of the intercept term is represented by s_{11} which also corresponds to the error term in the ordinary regression model. Term s_{11} can be arbitrarily set equal to one and the value of other elements of the matrix can be defined in relation to it. The variance of transient changes in the slope parameter is represented by s_{11}. Only permanent changes in this parameter shall be considered since this assumption obviates the necessity of having to specify the comparative magnitude of transitory variance between the intercept and the slope parameter. Off diagonal terms s_{12} and s_{21} also are zero when s_{22} is zero. Thus, up to a scale factor,

$$\Sigma u = \begin{bmatrix} 1 & 0 \\ 0 & 0 \end{bmatrix} \tag{15}$$

The variance of permanent changes in the intercept term of the market model is represented by r_{11}. Since the intercept of the model is theoretically zero, there is no need to allow for permanent changes in the intercept, and r_{11} can be set equal to zero, which implies that the covariance terms r_{12} and r_{21} are also zero. This matrix can be normalized by setting r_{22}, the variance of permanent changes in the slope parameter, equal to one.

Thus up to a scale factor

$$\Sigma v = \begin{bmatrix} 0 & 0 \\ 0 & 1 \end{bmatrix} \tag{16}$$

and covariance matrices of u_t and v_t from (13) can be written as

$$\text{Cov}(u_t) = (1-\gamma)\sigma^2\Sigma u = (1-\gamma)\sigma^2 \begin{bmatrix} 1 & 0 \\ 0 & 0 \end{bmatrix} \tag{17}$$

$$\text{Cov}(v_t) = \gamma\sigma^2\Sigma v = \gamma\sigma^2 \begin{bmatrix} 0 & 0 \\ 0 & 1 \end{bmatrix} \tag{18}$$

In other words, the variance of transient changes in the intercept is $(1-\gamma)\sigma^2$, and the variance of permanent step changes in the slope[4] is $\gamma\sigma^2$. Since both parameters σ^2 and γ as well as the linear coefficients β_{1t} and β_{2t} are estimated by the procedure described next, there is no need to prespecify how much variance arises from the slope changes.

Estimation

Parameter levels in a random walk process are nonstationary, and likelihood functions for them cannot be specified. But a likelihood function conditional on some specific realization at some point in time can be defined and the value of the parameter at that point can be estimated. Suppose a total of T observations ($t=1, \cdots, T$) are available for estimation, and the realized values of the parameter vector for period $t=\tau$ are to be estimated. Parameter vector β_t at any point in time can be written in terms of the parameter vector at the time τ and the error terms from (8) and (9):

$$\beta_t = \begin{cases} w_\tau - \sum_{s=t+1}^{\tau} v_s + u_t & \text{for } 1 \le t < \tau \\ w_\tau + \sum_{s=\tau+1}^{t} v_s + u_t & \text{for } \tau < t \le T \end{cases} \tag{19}$$

Since w_τ is the parameter of interest for estimation, let

$$\beta = w_\tau \tag{20}$$

Model (7) can now be rewritten in terms of β and the error term μ_t

$$y_t = x_t'\beta + \mu_t \tag{21}$$

where

$$\mu_t = \begin{cases} x_t'u_t - x_t' \sum_{s=t+1}^{\tau} v_s & \text{for } 1 \le t < \tau \\ x_t'u_t + x_t' \sum_{s=\tau+1}^{t} v_s & \text{for } \tau < t \le T \end{cases}$$

Since u_t and v_t are normally distributed with mean zero and covariance $\Sigma^* u$ and $\Sigma^* v$, it can be shown that μ_t is also normally distributed with mean zero and the covariance matrix,

$$\text{Cov}(\mu) = \sigma^2(1-\gamma)R + \gamma Q \equiv \sigma^2\Omega_\gamma \tag{22}$$

where R is a $(T \times T)$ diagonal matrix with

$$r_{ii} = x_i'\Sigma u x_i \tag{23}$$

since

$$\Sigma u = \begin{bmatrix} 1 & 0 \\ 0 & 0 \end{bmatrix}$$

and the first element of x_i is 1,

$$R = I_T$$

and ij^{th} element of matrix Q is given by,

$$q_{ij} = \begin{cases} \min(|\tau - i|, |\tau - j|) x_i'\Sigma v x_j & \text{for } \begin{cases} i \text{ and } j < \tau \\ i \text{ and } j > \tau \end{cases} \\ 0 & \text{for } \begin{cases} i \le \tau \text{ and } j \ge \tau \\ i \ge \tau \text{ and } j \le \tau \end{cases} \end{cases} \tag{24}$$

then the full model can be written as

$$y = X\beta + \mu \tag{25}$$

where

- y is a $(T \times 1)$ vector of dependent variable
- X is a $(T \times 2)$ matrix of independent variables
- β is a (2×1) vector of parameter values in period τ as defined in (20)
- μ is a $(T \times 1)$ vector of error terms as defined above

[4] Fisher (1970) parameterizes the problem somewhat differently. In his formulation K is the ratio of the step variance of the slope coefficient to the intercept variance:

$$K = \frac{\gamma\sigma^2}{(1-\gamma)\sigma^2} = \frac{\gamma}{1-\gamma}$$

or

$$\gamma = K/(1+K)$$

Thus γ and K have a one to one relationship. Fisher, however, does not propose a procedure for the estimation of K.

y is distributed normally with mean $X\beta$ and covariance $\sigma^2\Omega_\gamma$. When γ is known, the estimate of β can be obtained by applying generalized least squares to (25) since the covariance matrix is a function of X and γ alone. In the market model for stocks, the value of γ for each stock is not known and has to be estimated. Cooley and Prescott propose both maximum likelihood and Bayesian estimators for γ, σ^2, and β. The Bayesian procedure is used here for the estimation of the expected value of γ for each stock because this procedure allows inclusion of prior information about the distribution of γ (see equation (29)). The estimated expected value of γ is then used in Aitken's generalized least square procedure to estimate β.

Generalized least square estimate, B, of β conditional on γ is

$$B_\gamma = (X'\Omega_\gamma^{-1}X)^{-1}X'\Omega_\gamma^{-1}y \qquad (26)$$

The estimate of σ^2 conditional on B_γ and γ is

$$S_\gamma^2 = \frac{1}{T}(y - XB_\gamma)'\Omega_\gamma^{-1}(y - XB_\gamma) \qquad (27)$$

If the prior density of γ is $f(\cdot)$, the posterior density for γ is given by

$$p(\gamma:y) \propto f(\cdot)\,|\Omega_\gamma|^{-1/2}\,|(X'\Omega_\gamma^{-1}X)^{-1}|^{1/2} \cdot (TS_\gamma^2)^{-(T-2)/2} \qquad (28)$$

The expected value of the posterior distribution of γ is obtained by numerical integration. This value is then used in (26) and (27) to estimate the value β of linear coefficients at time τ and σ^2 by B_γ and S_γ^2 respectively.

In the next section, results from the estimation of risk of stocks around the date of accounting changes from the procedure defined above are presented.

Empirical Evidence on the Relationship of Accounting Changes and Stock Price Behavior

The results from the empirical analysis of the behavior of stock prices in relation to the accounting changes to and from the LIFO method of inventory valuation are presented next. Relative risk of each stock was estimated for each of the 24 months surrounding the time of accounting change by using the adaptive regression model described in the previous section. The estimates of relative risk were used to estimate the residuals which provide a measure of association between the accounting and price changes. Estimates of risk were averaged cross-sectionally to examine if the accounting changes are associated with shifts in relative risk. During the 21-year period from 1946 to 1966, 126 firms changed to the LIFO method (Group A) and 29 firms abandoned this method of accounting (Group B). The data used here are described in detail in Sunder (1973b). The data analysis was conducted in the following steps:

1. A prior probability density function for γ was assumed. Since no information was available to distinguish one stock from another in this respect, the assumed density was identical for all stocks. $\gamma=0$ implies that all the variance arises from the intercept or the additive disturbance term in the market model and the slope parameter representing the relative risk of the firm is a constant over time. $\gamma=1$, on the other hand, implies that all the variance arises from changes in the slope, and the additive disturbance term for all observations is identically equal to zero. A prior probability density function should represent the researcher's preconception about the chances of the occurrence of the values of γ in the zero to one interval. From the results given in Sunder (1973b, Chapter VI), γ is much more likely to be close to zero than to one because the null hypothesis that the risk of the firms during the months surrounding the accounting change was stable was rejected for only 20% of the stocks tested. Therefore a declining-

ramp prior density function for γ, defined by (29), was used in the present study.

$$f(\gamma) = 2(1-\gamma) \quad \text{for } 0 \leq \gamma \leq 1 \quad (29)$$

This density function has a value of two at $\gamma=0$ and declines uniformly to zero at $\gamma=1$. For each stock j, steps 2 through 4 were repeated.

2. The posterior density of γ was estimated from equation (28). Up to a total of 75 months of data from month -37 to 37 were used for this estimation. The month of accounting change was designated as month 0 and the data for this month were included for estimation. The first moment of the posterior distribution of γ was calculated. This statistic, $\hat{\gamma}_j$, was used as an estimate of γ for stock j. The stocks for which a minimum of 25 observations in the range specified above were not available were excluded from this analysis. Out of 155, 133 stocks were analyzed.

3. The intercept and the risk coefficients of the market model, as well as parameter σ^2, were estimated for each of the 24 months from month -11 to 12 by using estimate $\hat{\gamma}_j$ of γ in equations (26) and (27). A maximum of 41 months and a minimum of 25 months of data were used to estimate the relative risk for each month. For example, to estimate the risk in month -11, τ was put equal to -11 and the data from month -31 to month 9 were used. Similarly for the estimation of risk in month 12, the data from the month -8 to 32 were used. The purpose of using this moving series scheme was to maximize the precision of estimates from a given number (41) of observations since most precise estimates are obtained at the midpoint of the time series. Twenty-four estimates of relative risk for each stock (b_{jt}, $t=-11, \cdots, 12$) were stored.

4. For each of the 24 months surrounding the accounting change, the regression residual was calculated using the estimated coefficients of the market model for the respective months. Twenty-four residuals that were thus obtained ($\hat{u}_{jt}$, $t=-11, \cdots, 12$) were also stored in the memory.

5. After steps 2 through 4 had been complated for all stocks, the cross-sectional mean of the relative risk of a group of stocks was calculated for each of the 24 months. If there are N stocks in the group being analyzed, the average relative risk of the stocks in this group in month t is denoted by $\bar{b}_t$:

$$\bar{b}_t = \frac{1}{N}\sum_{j=1}^{N} b_{jt} \quad (t = -11, \cdots, 12) \quad (30)$$

6. The cross-sectional mean of residuals for the group in period t is $\bar{u}_t$:

$$\bar{u}_t = \frac{1}{N}\sum_{j=1}^{N} \hat{u}_{jt} \quad (t = -11, \cdots, 12). \quad (31)$$

7. Cumulative abnormal residual U_t for the group in period t is defined as

$$U_t = \sum_{s=-11}^{t} \bar{u}_s \quad (t = -11, \cdots, 12) \quad (32)$$

Results

The average relative risk of several groups and subgroups of stocks for the 24-month period around the date of accounting change is shown in Table 1. The first column of the table gives the average risk of 118 firms of Group A which switched over to LIFO in month 0. By themselves, these results show only small changes in the relative risk. But when considered with the results obtained in Sunder (1973a), they look quite significant because they confirm the earlier results which indicated that the average risk of these firms during the pre-change months was lower than in the post-change months. During the two years, the average risk of these firms increased by 5.4% from 1.058 to 1.115. The increase in risk was the general trend for this group for the entire 24-month period, though 78% of the total change occurred

TABLE 1

AVERAGE RELATIVE RISK OF STOCKS IN 24 MONTHS AROUND THE DATE OF ACCOUNTING CHANGE

Sample Size	Group A 118		Group A Without Steel 96		Group A Steel Only 22		Group B 21	
Month	Relative Risk (1)	Change (2)	Relative Risk (3)	Change (4)	Relative Risk (5)	Change (6)	Relative Risk (7)	Change (8)
−11	1.058		0.974		1.418		1.090	
−10	1.064	+0.006	0.974	+0.000	1.445	+0.027	1.106	+0.016
− 9	1.073	+0.009	0.979	+0.005	1.473	+0.028	1.098	−0.008
− 8	1.076	+0.003	0.979	+0.000	1.487	+0.014	1.107	+0.009
− 7	1.080	+0.004	0.986	+0.007	1.483	−0.004	1.106	−0.001
− 6	1.085	+0.005	0.988	+0.002	1.503	+0.020	1.105	−0.001
− 5	1.086	+0.001	0.987	−0.001	1.511	+0.008	1.111	+0.006
− 4	1.082	−0.004	0.980	−0.007	1.518	+0.007	1.102	−0.009
− 3	1.089	+0.007	0.985	+0.005	1.534	+0.016	1.087	−0.015
− 2	1.095	+0.006	0.997	+0.012	1.514	−0.020	1.076	−0.011
− 1	1.094	−0.001	0.993	−0.004	1.527	+0.013	1.070	−0.006
0	1.102	+0.008	1.000	+0.007	1.536	+0.009	1.065	−0.005
1	1.097	−0.005	0.992	−0.008	1.551	+0.015	1.054	−0.011
2	1.092	−0.005	0.988	−0.004	1.545	−0.006	1.066	+0.012
3	1.094	+0.002	0.990	+0.002	1.548	+0.003	1.061	−0.005
4	1.101	+0.003	0.997	+0.007	1.550	+0.002	1.058	−0.003
5	1.103	+0.002	1.001	+0.004	1.544	−0.006	1.028	−0.030
6	1.100	−0.003	0.998	−0.003	1.547	+0.003	1.095	+0.067
7	1.112	+0.012	1.011	+0.013	1.554	+0.007	1.007	−0.088
8	1.110	−0.002	1.009	−0.002	1.550	−0.004	1.021	+0.014
9	1.115	+0.005	1.016	+0.007	1.544	−0.006	1.040	+0.019
10	1.113	−0.002	1.023	+0.007	1.504	−0.040	1.012	−0.028
11	1.107	−0.006	1.015	−0.008	1.507	+0.003	1.023	+0.011
12	1.115	+0.008	1.026	+0.011	1.503	−0.004	1.032	+0.009

in the 12 months preceding the accounting change.

Since the steel firms represented a disproportionately large part of this sample (22 out of 118), steel and nonsteel firms were analyzed separately. The results are given in Table 1 and they generally support the remarks made above for the entire sample. The average relative risk of nonsteel firms increased by 5.3%, whereas the average relative risk of the steel firms increased by 6% during the 24-month period surrounding the date of accounting change.

For the 21 firms in Group B which made an accounting change from LIFO to FIFO, the average relative risk decreased by 5.3% during the 24-month period surrounding the date of accounting change.

The average abnormal price changes for each group of stocks, adjusted for relative risk and changes in relative risk, are shown in Figure 1 through Figure 4. Figure 1 gives the cumulative average abnormal residuals starting month −11 through month 12 for 118 firms which changed to the LIFO method. During the first 12 months, the average abnormal price change was 4.7% resulting in a 24-month increase of 2.9%. The increase of 4.7% during the first 12 months is only a little smaller than the 5.3% increase measured with the constant risk assumption in Sunder (1973a).

The exclusion of the steel firms from Group A (see Figure 2) reduces the abnormal price change during the first 12 months to 2.3% and during the 24 months to 0.8%. Although these numbers can hardly be considered significant by themselves, they are consistent with the price changes observed for Group A as a whole—abnormal price rise during the pre-change months

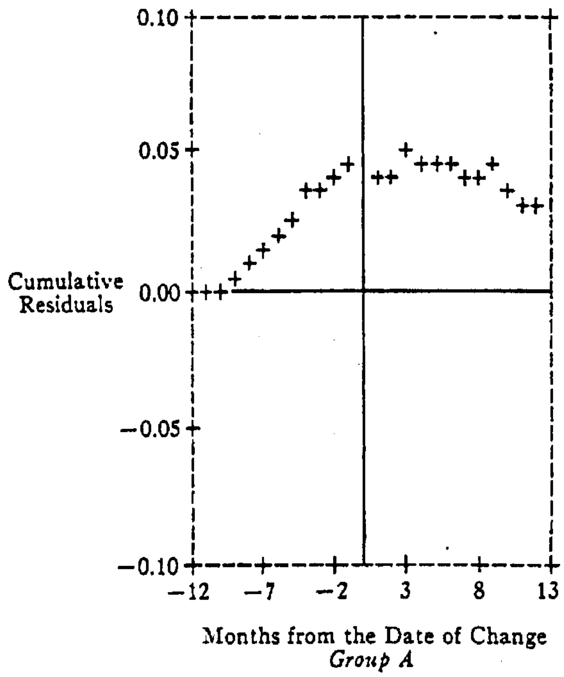

FIGURE 1
CUMULATIVE RESIDUALS FOR 24 MONTHS AROUND THE DATE OF ACCOUNTING CHANGE

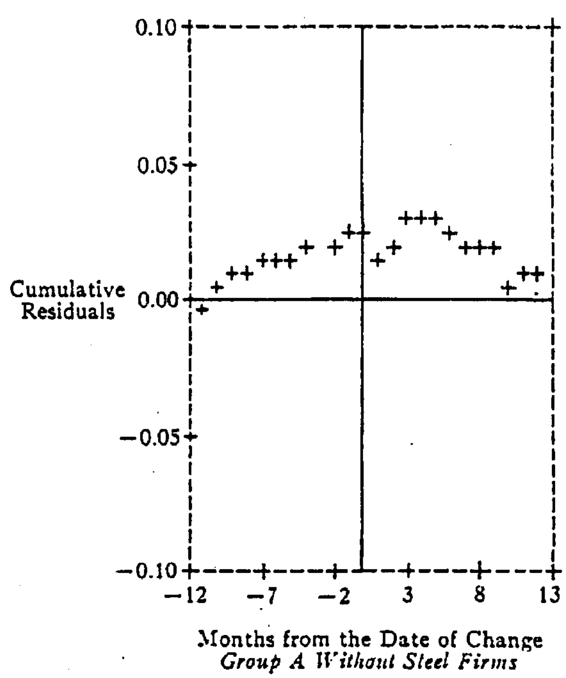

FIGURE 2
CUMULATIVE RESIDUALS FOR 24 MONTHS AROUND THE DATE OF ACCOUNTING CHANGE

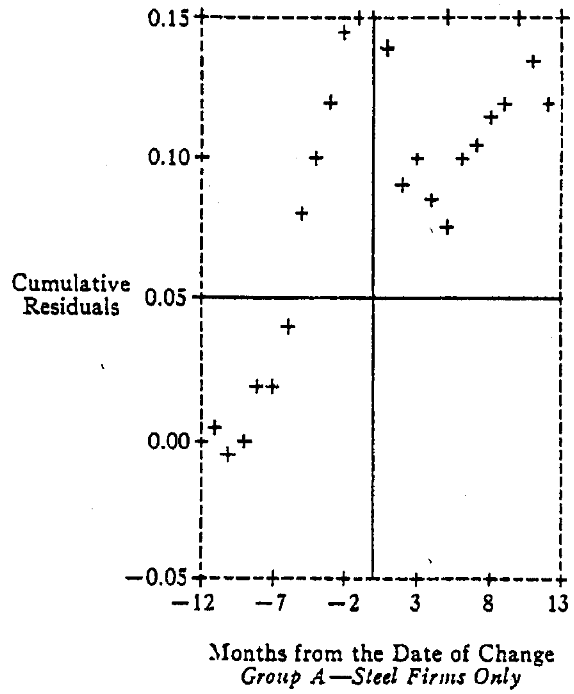

FIGURE 3
CUMULATIVE RESIDUALS FOR 24 MONTHS AROUND THE DATE OF ACCOUNTING CHANGE

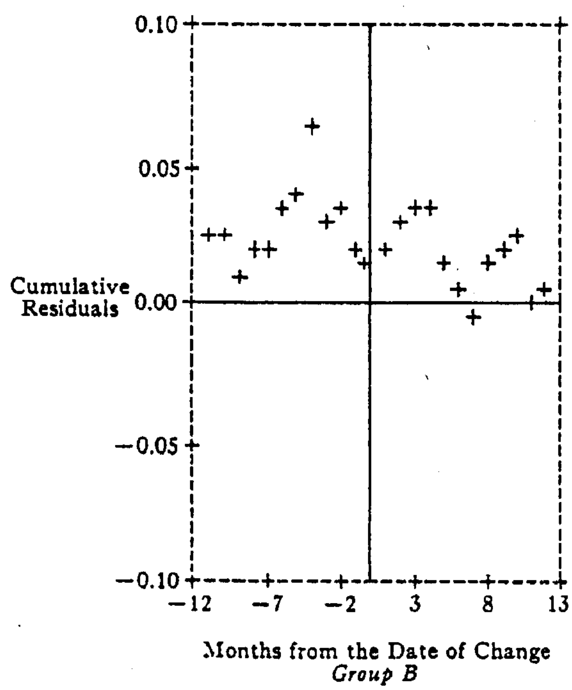

FIGURE 4
CUMULATIVE RESIDUALS FOR 24 MONTHS AROUND THE DATE OF ACCOUNTING CHANGE

followed by a smaller decrease in the post-change months.

The behavior of the price of steel stocks (see Figure 3) is also similar in nature, but the price changes are much larger. The abnormal price rise for these stocks was 14.9% during the 12 pre-change months followed by a 2.9% decrease during the 12 postchange months. These results correspond closely to those obtained in Sunder (1973a) under the constant risk assumption.

Figure 4 for Group B indicates that the average cumulative abnormal price change for these stocks is very close to zero. Because of the smallness of the sample, this figure has more noise than Figure 1. The price of these stocks does not seem to have increased as a result of the accounting change from LIFO to FIFO and the associated increase in the reported earnings. Industry-wide analysis of Group B was not conducted because further subdivision of this small sample was unliely to provide any further insights.

Conclusions

Accounting changes to the LIFO method of inventory valuation are found to be associated with an abnormal increase in the market price of stocks of the firms during the twelve months preceding the accounting changes. During the twelve months following the accounting changes, no significant abnormal price changes are observable for these firms. These results obtained after making adjustment for risk changes are substantially similar to the results obtained by the researcher in a previous study (Sunder, 1973a) in which adjustment for risk changes was not made. Therefore, the conclusions of this study about the relationship between accounting changes to the LIFO method and stock price changes are also similar. The results support the hypothesis that the changes in market price of stocks are associated with the changes in the economic value of the firms rather than with the changes in reported earnings. However, all the price increases observed during the twelve months preceding the accounting change cannot be interpreted as the *effect* of the accounting change since the decision to make the accounting changes is often taken towards the end of the fiscal year. An alternative explanation of the observed price behavior is that the firms which change their accounting method had better than average business prospects which were reflected in an abnormal rise in the price of these stocks and motivated the management to bring about the accounting change. Lack of availability of data on the dollar effect of the accounting changes in most cases did not permit identification between these two explanations.

The accounting changes to the LIFO method are found to be associated with an increase in the relative risk of the stocks.

No significant price changes are observable for the firms which abandoned the LIFO method. This is in contrast with the sharp abnormal decline in the price of these stocks measured without adjustment of risk changes (see Sunder, 1973a). The relative risk of these stocks decreased during the two-year period surrounding the date of accounting change.

This study does not support the view that the reduction in the reported income of a firm which accompanies a change to the LIFO method of inventory valuation is viewed by the stock market as a sign of adverse performance on the part of the firm. The results also do not support the view that corporate managers can, on the average, manipulate the stock price of their firms by adopting or abandoning the use of the LIFO method of inventory valuation.

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