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Different Market Reactions to Discretionary and Nondiscretionary Accounting Changes

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The recent accounting literature has reported the results of a number of studies that have examined the market effects of accounting changes (ACs). None of the published studies has dealt directly with nondiscretionary changes. Because the accounting change signals to investors may differ between discretionary and nondiscretionary ACs, the potential market effect of nondiscretionary ACs provides an interesting empirical question. Moreover, the published studies to date have used research designs that share a common weakness in the control group to which the returns on the experimental firms were compared. Using a research design that mitigates the effect of that weakness, this study provides some evidence that the two categories of accounting changes (dichotomized on discretion characteristics) convey uniquely different signals to investors. The research design used here also incorporates a number of additional refinements that have not heretofore been applied to AC research.

The theoretical link between economic events and stock returns in an efficient capital market is summarized in Beaver [1972]. The empirical link between accounting events and stock returns is well documented in

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Ball and Brown [1968] and Beaver [1968], among others. Relative to other accounting events (if one assumes that net income is either directly or potentially affected by a change), ACs possess the desirable characteristic that they allow for the positive assessment of the market's comparison of the difference between two income constructs. This evidence is believed to provide insights into investors' decision processes.

Discretionary and Nondiscretionary Accounting Changes

Discretionary ACs can serve as signals regarding firms' production-investment decisions or expectations about future cash flows (e.g., tax payments). Alternatively, they may be made in order to conform to legal requirements, such as debt or other covenants, or they may be induced by the desire to smooth income in unusually profitable periods or to keep an earnings trend from tapering off. Given that management has decided to make a change, the list of potential sources of motivation is almost endless. Therefore, while the AC, per se, may appear to have no substantive effect on the entity, the joint implication of the change and other events in the firm may have some substance. For this reason, predicting the market effects of most discretionary ACs is a very tenuous activity. It would be convenient if the accounting profession's authoritative pronouncement on ACs, *APB Opinion No. 20*, detailed the conditions under which ACs could be made, but this is not the case. Consequently, it is difficult to isolate the aspect(s) of a discretionary AC which is (are) most closely related to concurrent market behavior.

Nondiscretionary ACs, on the other hand, are induced by some agent exogenous to the firm, so that management has less control over the content and the timing of the change announcement. Accordingly, the signal emitting from the nondiscretionary change is more likely to be interpreted as a signal about the *change*, rather than about other events in the firm. Consider, for example, a discretionary AC that increases net income. This change may be interpreted as a signal of management's expectations about events which will have an unfavorable bearing on the firm. The same AC, if induced by an FASB statement, may be interpreted as a signal about the exact magnitude by which management has conservatively biased reported income. Thus the likelihood of a confounding informational effect is reduced for the nondiscretionary AC. Knowledge of the market effects of nondiscretionary ACs, relative to the evidence on the market effects of discretionary ACs, should benefit the FASB in understanding the use made by investors of accounting data.

The above discussion suggests that the discretion characteristic is one determinant of the market effect of an accounting change. It is possible that the sign of the earnings effect is also important to investors. To see this, consider what would happen if the income effects of the hypothetical ACs considered above were reversed. A discretionary AC that re-

duces net income may impart the signal that management expects future events more than to offset the negative earnings impact of the change. As a corollary, it may reflect favorably on the "quality" of the earnings number and, by implication, on the firm in general. The nondiscretionary AC which reduces net income may be interpreted as a measure of management's liberal bias in previously reported income and hence may produce an unfavorable signal about the firm.

Models Used

The usual assumptions that (*a*) equation (1) below reflects the return-generating function common to all firms and (*b*) capital markets are efficient in the semistrong form (see Fama [1970]) are made here. Assumption (*a*) provides the common denominator that serves as a basis for comparing sample firms. Assumption (*b*) provides the link between the denominating unit, security returns, and the information content of events. If one makes the capital market efficiency assumption, one avoids the untenability of implicitly assuming market efficiency and then stating conclusions in terms of whether the evidence proves or disproves the assumption.[1]

In this study, the modification of Sharpe's [1964] capital asset pricing model proposed by Black [1972] is assumed to generate the single-period expected return from holding security i, as in (1) below. A tilde (~) denotes a random variable throughout.

$$E(\tilde{R}_{it}) = E(\tilde{R}_{zt}) + \beta_{it}[E(\tilde{R}_{mt}) - E(\tilde{R}_{zt})] \tag{1}$$

where: E = expectation operator
R = return, i.e., price change plus dividends, all for period t
$\tilde{R}_{it}$ = return on security i for period t
$\tilde{R}_{zt}$ = return on a portfolio of securities, z, with $\beta_{zt} = 0$, for period t
$\tilde{R}_{mt}$ = return on a portfolio of securities, m, representing the market, for period t
β_{it} = $\mathrm{Cov}(\tilde{R}_{it}, \tilde{R}_{mt})/\mathrm{Var}(\tilde{R}_{mt})$, i.e., the relative risk of security i, for period t.

Literature Review

The market effects of accounting changes have been examined by Archibald [1972], Ball [1972], Baskin [1972],[2] Kaplan and Roll [1972], and Sunder [1973; 1975]. All of these studies used some variant of the Abnormal Performance Index (API) as the measure of the market reaction to the ACs under investigation. The (implicit) control group in

[1] See Sunder [1973, p. 6] on this point.

[2] Baskin's [1972] study ostensibly tested the effect of consistency exceptions. Yet, his design does not distinguish whether the event of interest was the consistency exception or the accounting change; accordingly, his study is considered to be an AC study.

the Archibald [1972] and Ball [1972] studies was the equilibrium expected return of the experimental group of change firms, where the expectation is given by some variant of (1), or of the so-called market model, given by:

$$E(\tilde{R}_{it}) = \alpha_i + \beta_i R_{mt} \tag{2}$$

where R_{it} and R_{mt} are as defined in (1); α_i and β_i are regression constants, analogous to R_{zt} (1-β_{it}) and β_{it}, respectively, in (1), but here are assumed to be constant across time.

In all six of the studies cited above, the API (i.e., e_{it}, the component of realized R_{it} that is peculiar to security i) during period t was computed as:

$$e_{it} = R_{it} - E(\tilde{R}_{it}). \tag{3}$$

Then, in all the studies except Baskin [1972], inferences about the effect of the ACs were made on the basis of the behavior of the means of the e_{it}, averaged across securities i for each time period t, that is, $\bar{e}_t$, during the hypothesized impact period.

It should be noted that the designs used by Baskin [1972], by Kaplan and Roll [1972], and by Sunder [1973; 1975] did use control groups of (other) firms. However, none of the control groups was specifically matched in order to hold constant the effects of production-investment and financing decisions. As a consequence, it is not known whether, in the absence of the experimental groups' having made ACs, one should expect a priori the realized returns on the control groups to equal the realized returns on the experimental groups.

Research Design

The research design used here is a modification of the one introduced by Gonedes [1975] and used by him to examine the effects of special income items. The main feature of the design is that it compares realized total returns of two groups of firms whose expected returns are equivalent because the estimated relative risks, $\hat{\beta}_{it}$, of experimental and control groups are equated at the beginning of the test period. Thus, any realized return differences are more likely to be associated with the independent variable of interest than in studies which use randomly selected control groups. Moreover, a contemporaneous comparison of returns on an experimental group and a separate control group will also reduce the measurement error that results from misspecification of the return-generating model or from estimation of its parameter(s).

The modification introduced in the present study involves the possibility that relative risk does not capture all the firm characteristics needed to ensure that $E(\tilde{R}_1) = E(\tilde{R}_2)$, where the subscripts refer to accounting change and nonchange, respectively. Fama and Miller [1972] develop the theory that the expected return of a firm is a function of its (unobservable) production-investment and financing decisions. Relative

risk, β_{it}, is merely an observable surrogate for those decisions. Two firms (groups of firms) may have the same relative risk and engage in very different production-investment and financing decisions. Thus, it is possible that the different production-investment and financing decisions of the two firms (groups of firms) with equal β_{it} could produce very different signals to investors, quite apart from the information content of the independent variable under examination.

This possibility has significant implications for an AC study because two earlier AC studies, those of Gosman [1973] and Sunder [1975], suggest that firms in the same industry tend to make ACs en masse. (This is also true for the AC firms examined in this study.) Sunder's [1975] results also suggest that the market effects of ACs are somewhat industry-dependent. Therefore, to the extent that industry membership affects the realized returns of AC firms, this variable needs to be controlled in this research.[3] To this end, the control group used in this study is, in general, composed of nonchange firms that have (*a*) the same estimated β_{it} and (*b*) the same industry membership as the firms that made ACs.

Another refinement that Gonedes [1975] introduced involves the possibility that market reaction to ACs varies across risk classes of firms. The reasons for believing that such a relationship may exist center on the facts that (*a*) firms' production-investment and financing decisions are determinants of their relative risks, and (*b*) events such as ACs can convey information regarding firms' production-investment decisions and/or their financing decisions. Taken together, these two facts suggest an association between ACs and relative risk. The link may be explained in terms of the signals conveyed by ACs due to the difference between the firms' risk classes. These signals should be related, in a broad sense, to the firms' production-investment and/or financing decisions, and to the extent that the two groups of firms' production-investment and financing decisions differ, the signals may differ.

Two hypothetical cases are offered as theoretical justification for supposing that investors may view the same accounting change differently due simply to the prechange relative risk of the firm that made the change. The two cases are by no means exhaustive, but they do illustrate possible reasons for expecting different market effects. In each case, it is assumed that there is some threshold that divides high-risk firms from low-risk firms and that this threshold is important to investors in pricing securities. The first case deals with discretionary ACs that increase net income. The second case concerns a nondiscretionary change that increases reported net income.

A firm may elect to change from accelerated to straight-line depreciation because of a reduction in the effect that obsolescence has on depre-

[3] Similar arguments could be raised about a variety of other factors, such as firm size, dividend policy, capital structure, etc. However, relative risk and industry membership capture, to some extent, the essence of these other variables as well.

ciation. This change may produce the signal to investors that the firm is less risky because it has relatively greater control over its asset use, that is, is less affected by factors beyond its control. It is reasonable to predict that such a change could be associated with a greater impact on the shares of high-risk firms since it produces new information relevant to valuing high-risk firms. In this scenario, the share prices of low-risk firms making this AC are not affected because the signal is consistent with investors' prior assessments of those firms as low-risk.

On the other hand, the discretionary-positive-effects AC may be viewed by investors as a signal that, without the change, the firm is unable to earn its target net income, which reflects a constant rate of growth over time. This AC may therefore signal an increase in the variability of the firm's income in the absence of the accounting change. Since Beaver, Kettler, and Scholes [1970] have shown that high earnings-per-share (EPS) variability is positively associated with high relative risk, this perceived increase in the variability of EPS may be incompatible with investors' tolerance for earnings variability of low-risk firms. This may in turn lead to a more pronounced effect on the returns of low-risk securities.

Next, consider a nondiscretionary change, from cost to equity for long-term investments. If this change increases reported income, it may be taken as a signal that the change firms had to be forced by an outside body to adopt an accounting method that reports more income than the companies had willingly chosen to report. That is, the firms had taken a more conservative option with respect to reporting income. This conservatism may logically be imputed to low-risk firms whose investment attractiveness lies more in their ability to generate dividends than to produce superior earnings (growth). If this is true, then it is reasonable that the returns of low-risk firms should show very little impact from this type of AC, because the nature of the change is consistent with other signals emitted by the firm. Investors in high-risk firms that made this same AC could be expected to interpret this change as a signal that their firms are not so risky as they had formerly believed, and this could lead in turn to an effect on their security return.

Based on Sunder's [1975] results, perhaps the partitioning should be done on industry membership in order to assess whether there is a unique relationship between information effect and industry membership. However, the partitioning on risk classes can be viewed as similar to a partitioning based on industry membership since there is a pronounced bunching on $\hat{\beta}_{it}$ of the present sample of firms which share the same industry membership.[4] Moreover, due to (*a*) the properties of the statistical tests performed here and (*b*) the relatively short time span

[4] For example, the average $\hat{\beta}_{it}$ of steel, auto-related, oil, and chemical firms in the sample are 1.087, 1.307, 1.025, and 1.133, respectively. The variances of these groups' average $\hat{\beta}_{it}$ are .160, .103, .069, and .067, respectively. The average $\hat{\beta}_{it}$ of all firms in the sample is 1.274, and the variance around this average is .195.

over which informational effects are measured, it was not feasible to partition the firms into more than two separate categories. Therefore, firms are partitioned into two risk classes, high and low, in order to determine whether the informational effects of ACs are dependent upon risk class. The resulting statistical tests are multivariate in nature to give effect to the simultaneous comparison of the returns on high-risk change and nonchange firms, and on low-risk change and nonchange firms.

Data and Data Selection

Results are reported for tests performed on four separate categories of accounting changes: discretionary changes with positive effects on net income, (+,*D*); discretionary changes with negative effects on net income, (−,*D*); nondiscretionary changes with positive effects on net income (+,*N*); and nondiscretionary changes with negative effects on net income, (−,*N*). In this study, a nondiscretionary AC is one that a firm made in response to an opinion of the Accounting Principles Board or as a result of an Internal Revenue Service tax ruling tied to the firm's financial reporting. Accounting changes that do not meet either of these criteria are judged to be discretionary.

Firms were identified as having made accounting changes from the 1969–73 issues of *Accounting Trends and Techniques*, which provides an annual tabulation of the numbers of firms (from its 600) which had various types of disclosures in their annual reports of the previous calendar year. All the firms in this sample are New York Stock Exchange firms. One additional restriction was that monthly return data for all firms had to be included on the tape developed by the Center for Research in Security Prices (CRSP).

The 766 accounting changes enumerated in *Trends* over the sample period of 1968–72 provided the beginning point in the selection of firms making ACs. These ACs were identified with 675 AC firms, which were reduced to 280 firms for the reasons listed in table 1. The conclusions of this research are based on these 280 accounting change firms. Throughout this study, an AC was construed to be two different accounting treatments accorded the same set of circumstances in two successive periods. Poolings were eliminated because they involve different circumstances in the two successive periods. The effect of the business combination precludes the possibility of the reporting entities in the two periods being the same, because prior to the combination the two entities had separate ownership, whereas subsequently they had common ownership.

After a change firm was initially identified, its annual report or its 10–K report was examined in order to categorize the change on the basis of (*a*) the directional effect of the accounting change on net income, and (*b*) whether it was discretionary or nondiscretionary. If the change firm had two or more accounting changes with different directional effects on

TABLE 1

Reasons Firms Were Eliminated from Original Sample of Accounting Trends and Techniques Change Firms, 1968–72

Total number of *Trends* change firms		675
Less:		
Firms that made both discretionary and nondiscretionary changes in the same year	43	
Firms whose discretionary changes either had a zero or immaterial effect on net income or the net income effect was not disclosed	161	
Firms whose nondiscretionary changes either had a zero or immaterial effect on net income or the net income effect was not disclosed	64	
Firms that were not listed on the New York Stock Exchange	70	
Firms for which either the annual report or the 10-K report was not available or the change could not be identified from reading the report	60	
Firms that had an insufficient number of monthly returns on the CRSP tape	6	
Firms that could not be suitably matched	6	
Firms whose changes were poolings of interest	8	418
Trends firms included in the sample		257
Non-*Trends* firms included in the sample		23
Total sample firms		280

net income, the firm was classified as positive or negative, depending on the net effect of all ACs it made during the year.

Hereafter, boldface notations refer to vectors to give effect to the multivariate assumption; also, the time subscript, t, is dropped for convenience. To insure that the vector of estimated relative risks of accounting change firms (ACFs), $\hat{\beta}_1$, equals the vector of estimated relative risks of nonchange firms (NCFs), $\hat{\beta}_2$, matching was done at the individual firm level (i.e., if $\hat{\beta}_1 = \hat{\beta}_2$ for each pair of individual firms, it would also hold for the two groups of firms). The $\hat{\beta}_i$ for the ACFs' were taken from monthly issues of Merrill Lynch, Pierce, Fenner & Smith's *Security Risk Evaluation*, selected as follows.[5] For each firm, the $\hat{\beta}_i$ was selected from the issue for the month which preceded the firm's fiscal year-end by four months. The fourth month prior to year-end is six months prior to the assumed year-end disclosure month ($t = 0$), so that the sample impact period began with the sixth month before $t = 0$. Each NCF's estimated β_i was taken from the same issue of *Security Risk Evaluation* as the $\hat{\beta}_i$ of the matched change firm to insure that their relative risks were estimated over the same time period.

In estimating $\hat{\beta}_i$, Merrill Lynch applies Ordinary Least Squares pro-

[5] *Security Risk Evaluation* contains two types of estimates of firms' β_i. One type is the unadjusted $\hat{\beta}_i$, and the other is Merrill Lynch's attempt to adjust for the regression tendency of extreme estimates of β_i. The unadjusted $\hat{\beta}_i$ are used in this study. Also, the $\hat{\beta}_i$ of 1968 change and nonchange firms were not taken from *Security Risk Evaluation* because the publication was not begun until 1969. Instead, 1968 firms' β_i were estimated in the same manner that Merrill Lynch estimates β_i, all over sixty months' data.

cedures to the realized form of the market model, given by:

$$\tilde{R}_{it} = a_i + b_i \tilde{R}_{mt} + e_{it} \tag{4}$$

where R_{it} and R_{mt} are realizations of the expectations of R_{it} and R_{mt}, respectively, from (2); a_i and b_i are estimates of α_i and β_i, respectively, from (2); e_{it} is the realized residual return on security i for period t. Standard and Poor's Index of 500 firms is used as the proxy for R_m. Most of the firms' β_i in *Security Risk Evaluation* are estimated over periods of sixty months. Although others are estimated over shorter periods, no biases should result from including such firms since they only constituted about 5 percent of the present sample.

It was not possible to select each NCF with a $\hat{\beta}_i$ identically equal to the $\hat{\beta}_i$ of its matched ACF, except in a very few cases. The average absolute difference between the two $\hat{\beta}_i$ of a matched pair is around .085, and the median absolute difference is around .06. The largest acceptable difference between the two $\hat{\beta}_i$ of a matched pair was arbitrarily set at .4. The average standard error of the $\hat{\beta}_i$ is around .33 for firms which are classified as high-risk; the average standard error of low-risk firms' $\hat{\beta}_i$ is around .24. The standard errors are large in relation to the differences between the two $\hat{\beta}_i$ of a pair. These large errors in estimating β_i provide the main motivation for grouping firms. The standard errors of the average group estimated beta, $\hat{\beta}_g$, are estimated to fall between .05 and .11 for high-risk groups and between .035 and .08 for low-risk groups in this study.[6] These estimates are in line with the standard errors of portfolio $\hat{\beta}_p$ in Gonedes' study.[7] But they are a little larger, perhaps because the $\hat{\beta}_g$ in this study are somewhat larger than the $\hat{\beta}_p$ of his portfolios.

Firms were also matched on fiscal year-end. Given a discrete impact period common to both members of a pair, this matching holds constant the timing of the impact of the two firms' general year-end disclosures. Such a control feature takes on added importance in light of evidence about seasonality in monthly return distributions provided by Rozeff

[6] These estimates of the standard errors of groups' $\hat{\beta}_g$ were not computed in the normal manner, that is, from a time series of $\hat{\beta}_g$ which would be taken from a regression of $\tilde{R}_{gt}$ on $\tilde{R}_{mt}$ over t time periods. This method is not feasible here because firms that made accounting changes in different time periods, t, are aggregated into each group, g. Consequently, there is more than one $\tilde{R}_{mt}$ pertaining to each $\tilde{R}_{gt}$. Fortunately, however, Fama and MacBeth have developed a way of approximating the standard error of a group (or in their case, a portfolio) $\hat{\beta}_g$ ($\hat{\beta}_p$). They averaged the standard errors of firms' $\hat{\beta}_i$ in their respective portfolios. Then they computed the standard errors of the $\hat{\beta}_p$ of their portfolios and found that the ratio of the simple average standard error of the $\hat{\beta}_i$ in a portfolio to the standard error of the portfolio $\hat{\beta}_p$ was between three and seven. This means that the β_p of portfolios (of around size forty in their sample) can be estimated with between three and seven times the precision of the average of the $\hat{\beta}_i$ in a portfolio. See Fama and MacBeth [1973, pp. 615–21].

[7] See table 6 in this study and table 2 in Gonedes [1975] for a comparison of group $\hat{\beta}_g$ and portfolio $\hat{\beta}_p$, respectively. For the standard errors of Gonedes' $\hat{\beta}_p$, see Gonedes [1975, p. 235].

and Kinney [1976]. They found that, on the average, the monthly returns of New York Stock Exchange firms were unusually high in January and June and unusually low in February.

Tables 2–6 provide profiles of the various aspects of the firms in the sample. Table 2 gives the numbers of ACFs selected from each of the five years examined here. Tables 3 and 4 show the main industry concentrations of change and nonchange firms. Note that approximately 30 percent of the accounting change firms in the sample came from four industries—steel, auto-related, oil, and chemicals. Thus, the industry matching should be worthwhile, since a randomly selected control group would probably have exhibited greater differences on the underlying production-investment and financing decisions than does this sample. I should also mention that for each ACF, the selection of matched NCFs was first attempted at the three-digit Standard Industrial Classification (SIC) Code number level. If no NCF with the same year-end and similar $\hat{\beta}_i$ was available, an attempt was made to locate an NCF from the same two-digit SIC category.

TABLE 2
Calendar Years from Which Accounting Change Firms Were Selected

Year	Number of firms
1968	34
1969	55
1970	52
1971	77
1972	62
Total	280

TABLE 3
Industries of Sample Firms

Three-digit SIC code	Industry	Change firms	Nonchange firms
331	Steel	26	20
371	Automobiles, Parts and Accessories	23	21
291	Oil	20	17
280	Chemicals	14	16
367	Electronics	10	15
357	Office and Business Equipment	9	7
372	Aerospace	8	7
220	Textiles	9	5
999	Conglomerates	7	7
356	Industry Machinery and Pollution Control	5	7
374	Railroad Equipment	6	6
541	Retail Food Chains	7	4
355	Specialty Machinery	5	6
260	Paper	6	5
		155	143
	All others (ten or less)	125	137
	Total	280	280

TABLE 4
Industries of Sample Firms

Two-digit SIC code	Industry	Change firms	Nonchange firms
370	Transportation manufacturers	38	34
350	Machinery and equipment	28	30
330	Iron and steel; smelting and refining	31	24
280	Chemicals and drugs	24	27
360	Electrical and electronics	22	29
290	Oil and building materials	25	22
		168	166
	All others (less than twenty)	112	114
	Total	280	280

Table 5 gives the numbers of accounting changes which affected the various accounts, along with the categorizations of ACs on the basis of relative discretion and directional effect on net income. There are sixty-six more changes listed in table 5 than there are accounting change firms because some firms made more than one change in a given year. Also, in a few of these cases, the extra change had a zero (0) or an undisclosed (?) effect on net income. However, if a change firm was categorized as discretionary (nondiscretionary) in a given year, all its ACs that year were discretionary (nondiscretionary). It is clear from table 5 that the discretionary-positive-effects ACs are dominated by four types: depreciation changes to straight-line; equity investment changes; changes to capitalization of costs or to cash basis for expenses; and changes in pension actuarial assumptions. Also, around 80 percent of the nondiscretionary-positive-effects ACs are changes from cost to equity for long-term investments in stock.

Table 6 gives the sample sizes and average group $\hat{\beta}_i$, $\hat{\beta}_g$, (with $g = 1, 2$ for change and nonchange, respectively), for the various hypothesis groups on which formal tests were performed. In general, the $\hat{\beta}_g$ matches in table 6 are relatively close. The largest difference of .038 occurs between the low-risk firms for which the firms' changes were positive-nondiscretionary, that is, ACF (+, N). In order to determine whether the $\hat{\boldsymbol{\beta}}_g$ difference (of .02 for high-risk firms and .038 for low-risk) could cause spurious results, the statistical tests (to be described later) were performed a second time to correct for the differences in estimates of $\boldsymbol{\beta}_g$. The results of this cross-validation test were quite similar to those of the principal test conducted. By implication, then, the $\hat{\boldsymbol{\beta}}_g$ differences in the remaining test groups of table 6 should also not cause spurious results. The details of this cross-validation test are discussed in conjunction with the results of the test of ACF (+, N)–NCF. Table 6 also indicates that groups of sample firms tend to have above-average $\hat{\beta}_g$. This may be due either to the exclusion of utilities from the sample (the Standard and Poor's 500 Index includes utilities), or the fact that firms which made ACs during 1968–72 had above-average relative risk.

TABLE 5

Types of Accounting Changes by Account Affected, Discretion of Management in Making the Change, and Directional Effect on Net Income

	Discretionary				Nondiscretionary				Total
	+	−	0	?	+	−	0	?	
Equity Investment									
To Equity from Cost	12		2		50	4			68
To Inclusion	14	2	7	3					26
To Exclusion	1		1	2					4
Depreciation									
To Accelerated	1								1
To Straight-Line	51								51
Service Lives	13		1	1	3				18
Other	5	1	1		1				8
Inventory									
From LIFO	16			1	1				18
To LIFO		2							2
Other		1	1						2
Expense Recognition									
Intangibles		2							2
To Capitalization—To Cash	23	1	2		1				27
To Expense—To Accrual	4	7	1						12
Other	3	1			1				5
Revenue Recognition									
Construction Contracts	4		1						5
Other		3				1			4
Income Taxes									
To Tax Allocation					3				3
Investment Tax Credit	14	5							19
On Undistributed Subsidiary Income	1				2	4	1		8
Pensions									
Actuarial Assumptions	45	4	4	2					55
Actuarial Methods	1		1						2
Amortization Period				1					1
Foreign Currency Translation	2		1						3
Discount Receivables/Pay						1			1
Other		1							1
Totals	210	30	23	10	62	10	1	0	346

TABLE 6

Numbers of Pairs of Firms and Average $\hat{\beta}_g$ for Hypothesis Groups

Hypothesis group[a]	Number of pairs	$\hat{\beta}_g$ of high-risk		$\hat{\beta}_g$ of low-risk	
		ACF	NCF	ACF	NCF
ACF(+, *D*)-NCF	183	1.575	1.564	.924	.940
ACF(−, *D*)-NCF	27	1.915	1.898	.870	.870
ACF(+, *N*)-NCF	60	1.560	1.580	.951	.989
ACF(−, *N*)-NCF	10	1.810	1.800	1.188	1.196

[a] Each hypothesis group is identified by the accounting change firms (ACF) that made the type of change noted in parentheses and the matched nonchange firms (NCF), hence the notation ACF-NCF.

Hypotheses

Table 6 lists the four hypothesis groups on which tests are performed. In each case, the null hypothesis is given by:

$$H_0\colon \boldsymbol{\mu}_d|\boldsymbol{\theta} = \boldsymbol{\mu}_0 = 0 \tag{5}$$

where: $\boldsymbol{\mu}_d|\boldsymbol{\theta} =$ a 2×1 column vector of average return differences of $\bar{\mathbf{R}}_1|\boldsymbol{\theta}_1 - \bar{\mathbf{R}}_2|\boldsymbol{\theta}_2$, conditional upon the information variable, $\tilde{\boldsymbol{\theta}}_i$.

$\tilde{\boldsymbol{\theta}}_i$ = the information variable, an accounting change in this case, with $i = 1$ or 2 representing the presence or absence, respectively, of an AC.

The two-sided alternative hypothesis is given by:

$$H_1\colon \boldsymbol{\mu}_d|\boldsymbol{\theta} \neq \boldsymbol{\mu}_0 = 0. \tag{6}$$

In words, the null hypothesis states that the conditional mean return vector of ACFs, $\bar{\mathbf{R}}_1|\boldsymbol{\theta}_1$ is not different from the conditional mean return vector of nonchange firms, $\bar{\mathbf{R}}_2|\boldsymbol{\theta}_2$. The sample conditional mean difference vector, $\bar{\mathbf{d}}|\theta$, is used to estimate the population parameter, $\boldsymbol{\mu}_d|\boldsymbol{\theta}$, in the tests performed here. The conditional difference vector is used in the context of a single-sample test because the matching of ACFs and NCFs induces a dependency between the two groups, which implies that the appropriate test on mean vectors is a matched-pairs test.

For each of the four hypothesis groups, the null hypothesis of no information effect associated with accounting changes is tested over the thirteen-month period surrounding the post-year-end month when most firms issue their preliminary earnings and/or complete annual reports. If $t = 0$ is the second month after a firm's fiscal year-end, the period runs from six months prior to $t = 0$ through six months subsequent to $t = 0$.

The months prior to year-end are of interest for the following three reasons: (*a*) some firms announce their ACs in their quarterly earnings summaries; (*b*) it is possible that third- and fourth-quarter stock price movements induce some firms to make discretionary ACs; and (*c*) firms compelled to make nondiscretionary ACs are generally known prior to year-end due to the publicity given the pronouncement that forces the change. The months subsequent to year-end are of interest because the year-end preliminary or complete annual reports contain more information than the quarterlies regarding certain changes (e.g., changes in estimates and the effect of changes). Therefore, this combined thirteen-month period should contain the bulk of any unique market effect associated with ACs. At the same time, the period is short enough to avoid any substantial overlap with the hypothesized market effects associated with other ACs made by the sample firms in prior and subsequent years. Archibald [1972], Kaplan and Roll [1972], and Sunder [1975] all observed some market effect during this period of time. In addition to the formal statistical tests performed here, plots of the

cumulative average return differences are also presented in order to reveal the relative timing of the market effects.

Procedures for Grouping Firms and Computing Returns

For each of the four hypothesis groups, the $\hat{\boldsymbol{\beta}}_i$ of ACFs were ranked in descending order. The high-risk group is made up of the ACFs with the $\hat{\beta}_i$ in the top half of the ranked $\hat{\beta}_i$. The low-risk group of ACFs have the $\hat{\beta}_i$ in the bottom half of the ranked $\hat{\beta}_i$. The $\hat{\beta}_g$ for each risk group is the arithmetic average of all the $\hat{\beta}_i$ in the group. If the total number of pairs is odd, the low-risk group is one larger than the high-risk group. The high- and low-risk groups of NCFs are the firms matched with the ACFs that fall in the respective high- and low-risk groups of ACFs.

This method of partitioning firms into risk classes is known to induce a regression effect with respect to the estimates of relative risk (see Black, Jensen, and Scholes [1972] and Fama and MacBeth [1973]). However, this phenomenon should not impinge upon the ability of the tests to produce strong inferences about the informational effect of accounting changes because of the careful matching procedures. Recall that *firms* were matched on $\hat{\beta}_i$. Therefore, the "regression toward the mean" of extreme $\hat{\beta}_1$ should be substantially offset by the impact of the same phenomena on the extreme (matched) $\hat{\beta}_2$.

Once the high-risk and low-risk groups of ACFs (and NCFs) are formed, total monthly returns are computed on each ACF and each NCF from $t = -6$ through $t = +6$. Returns on the ACFs (NCFs) within each risk group are averaged for each of the thirteen months of the impact period, yielding an ACF (NCF) group return for each month.

Letting month $t = 0$ be the second month after each firm's fiscal year-end and not an identical calendar month for all change firms in a hypothesis group represents a difference between the estimation procedure used here and that used by Gonedes [1975]. He equated $\hat{\beta}_p$ at March of each year and computed returns from April through March of the next year, thereby ignoring the different fiscal year-ends of firms in the hypothesis group. As a result, it is possible that he was capturing in $\bar{\mathbf{d}}|\boldsymbol{\theta}$ some "noise" that was unrelated to the information variable $\bar{\boldsymbol{\theta}}_i$. The procedure used here, on the other hand, is designed to measure the most pronounced effects of accounting changes and AC disclosures, assuming that such effects are revealed during the thirteen-month period around the post-year-end date when most ACs are disclosed to the public.

After the returns of AFCs and NFCs within each risk class have been averaged for each month, thirteen monthly conditional return difference vectors, $\mathbf{d}|\boldsymbol{\theta}$, are computed by taking the difference between conditional changes and nonchange returns, $\mathbf{R}_1|\boldsymbol{\theta}_1 - \mathbf{R}_2|\boldsymbol{\theta}_2$, for $t = -6$ through $t = +6$. These thirteen conditional group return difference vectors are then averaged to get the conditional *average* return difference vector, $\bar{\mathbf{d}}|\boldsymbol{\theta}$, which is used in the statistical tests.

Statistical Tests

The tests employ Hotelling's T^2 statistic, a multivariate analog to the standard t-test. The single-sample test was performed because of the dependency between ACFs and matched NCFs, so the test is a multivariate matched-pairs t-test. The information conditioning argument, $\boldsymbol{\theta}$, is omitted from equations (7) and (8) below to avoid clutter.

The T^2 statistic has the form:

$$T^2 = N(\bar{\mathbf{d}} - \boldsymbol{\mu}_0)' S_d^{-1} (\bar{\mathbf{d}} - \boldsymbol{\mu}_0). \tag{7}$$

All terms in (7) are as previously defined except N, which is the sample size (i.e., the thirteen monthly return observations used to estimate $\bar{\mathbf{d}}|\boldsymbol{\theta}$), and S_d, which is the sample covariance matrix of the mean difference vector $\bar{\mathbf{d}}|\boldsymbol{\theta}$. The elements of S_d are defined in (8) where H denotes high-risk and L denotes low-risk.

$$S_d = \begin{bmatrix} \mathrm{Var}(d_H) & \mathrm{Cov}(d_H, d_L) \\ \mathrm{Cov}(d_H, d_L) & \mathrm{Var}(d_L) \end{bmatrix}. \tag{8}$$

A condition necessary for virtually all statistical applications is that the sampling units be independent of each other. In this case, the sampling units are monthly return observations, and the weak-form tests of market efficiency provide evidence that monthly price observations (and price changes) are very nearly serially independent. A further discussion of weak-form market efficiency can be found in Fama [1970].

When the null hypothesis (5) is true,

$$F = \frac{N - R}{R(N - 1)} T^2 \tag{9}$$

has the F distribution with degrees of freedom R (= 2, for the number of risk classes) and $N - R$ (where N = 13). Departures of $\boldsymbol{\mu}_d|\boldsymbol{\theta}$ from $\boldsymbol{\mu}_0$ increase the mean of T^2, so the decision rule is to accept the null hypothesis if the observed F-value falls beneath the area of the F distribution corresponding to one's desired level of significance. Otherwise, one rejects the null hypothesis and infers that the ACs have information content. A thorough description of the T^2 statistic can be found in both Morrison [1967] and Anderson [1958].

Risk and Information Effects

In order to determine whether accounting changes possess information for particular risk classes, the two risk-class group return differences (of ACF-NCF) used in each statistical test were cumulated and plotted in a time series. The cumulative return difference (CRD) is defined as:

$$\widetilde{CRD}_t = \sum_{t=-6}^{+6} \bar{\mathbf{d}}_t|\boldsymbol{\theta} \tag{10}$$

where all terms have been previously defined. The *CRD* plots thus portray the unique market effects associated with the accounting changes made by the two risk classes of firms. If different cumulative return differences result for the two risk classes, this will indicate the existence of a risk-dependency of the AC information. On the other hand, similar *CRD* plots will signal the absence of risk dependency.

Care must be exercised in using the figures to interpret the results of the statistical tests for two reasons. First, the figures plot cumulative return differences and not the average return differences used in the tests. As a consequence, the same amount of variability plotted in a time series of *CRD* is more likely to render it statistically insignificant than the related average return difference. That is, the variability visually represented in a cumulative return difference plot affects the entire cumulation and not merely the time series of the underlying average. Second, since the related statistical tests compare the returns of change and nonchange firms across high-risk and low-risk portfolios, the two risk-class *CRD* plots should be visually compared to the horizontal zero-return line and not to each other.

Results

Table 7 presents the summary statistics for the four hypothesis tests, each of which is identified with a unique line number. For each test, the table gives the average return differences on high-risk and low-risk firms, the standard errors of the estimates of average return difference, and the resulting *F*-values for the statistical tests. Figures 1–4 also present the plots of the two risk-class *cumulative* return differences in each test.

The tests of discretionary ACs that increase net income, that is, AC (+, *D*), and of nondiscretionary ACs that increase net income, that is,

TABLE 7
Summary Statistics for Tests

Line number	Hypothesis group	Number of pairs	Average return difference		*F*-value[a]
			High-β	Low-β	
(1)	ACF(+, *D*)–NCF	183	−.00208	−.00797	4.837
			(.00377)[b]	(.00264)	
(2)	ACF(−, *D*)–NCF	27	−.00610	−.00632	1.311
			(.01043)	(.00585)	
(3)	ACF(+, *N*)–NCF	60	.01361	.00295	4.961
			(.00473)	(.00253)	
(4)	ACF(−, *N*)–NCF	10	−.00910	−.01200	.644
			(.01109)	(.01674)	

[a] Selected fractiles of $F_{2,11}$ are:

Fractile	Value of $F_{2,11}$
.900	2.86
.950	3.98
.975	5.26
.990	7.21

[b] Standard error of average return difference estimate immediately above.

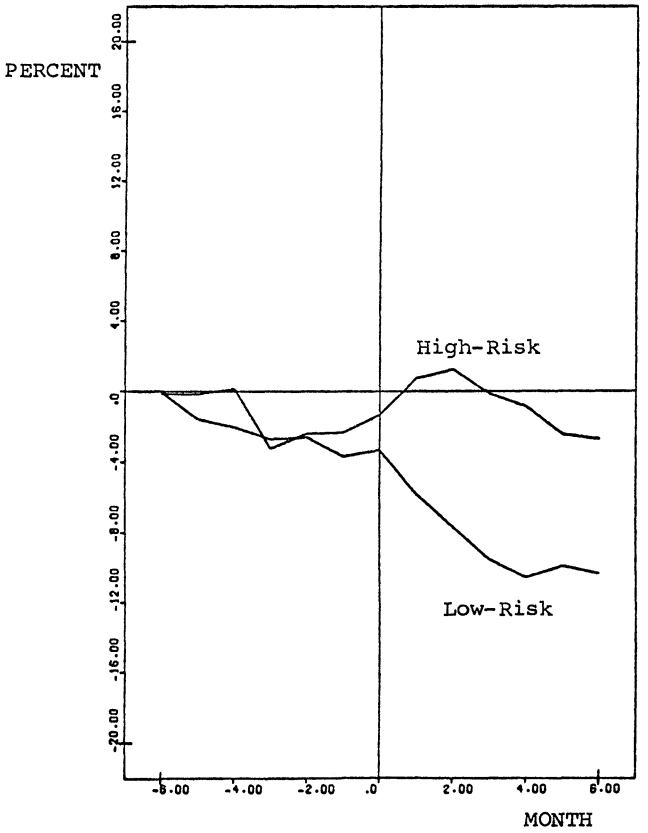

FIG. 1.—Cumulative return difference of ACF (+, D) - NCF; 183 pairs of firms

AC (+, N), both yielded statistically significant F-values. By contrast, the tests of discretionary ACs that decrease net income, that is, AC (−, D), and of nondiscretionary ACs that decrease net income, that is, AC (−, N), appear not to be associated with unique market behavior.

The F-value of 4.837 in the test of discretionary ACs that increase net income is significant at less than .05. Furthermore, the average return differences in this test are negative for both risk classes. Taken at face value this means that, on the average, the market has assigned lower relative values to firms that make discretionary ACs. The test conducted here does not offer clues as to the reason for the negative relative assessment of the change firms, but recent evidence by Bremser [1975] sheds some light on this subject. He found that firms that made ACs (+, D) exhibited a poorer pattern of EPS and return on stockholders' equity than a randomly selected control group of nonchange firms.

Whether ACs actually convey value-pertinent information is not directly addressed here inasmuch as the test covers the entire thirteen-month period around the year-end disclosure month. However, figure 1 does provide some insight into the timing of the market effect. From the

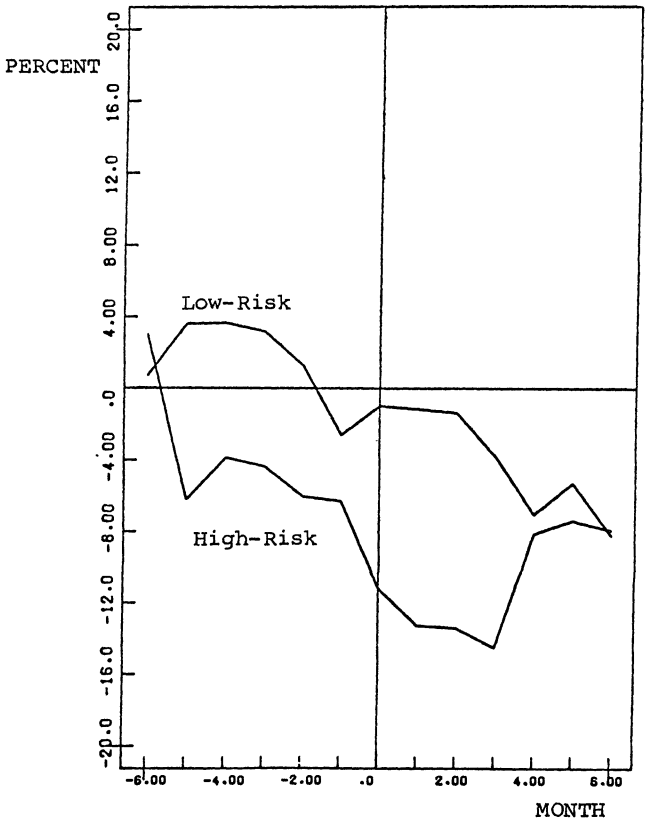

FIG. 2.—Cumulative return difference of ACF (−, D) - NCF; 27 pairs of firms

figure it appears that in general, the average return differences are much closer to zero during the period prior to year-end than during the post-year-end period. This in turn suggests that the more detailed year-end disclosures have information content. While some may interpret this as an evidence of market inefficiency, I prefer to think that the year-end disclosures simply provided possibly additional information for investors. For example, it is very difficult to predict the full net income effect of many ACs prior to year-end,[8] in which case a disclosure of the magnitude of the net income effect of an AC (+, D) could have significant information content.

The average return differences in the test of ACF (+, N)–NCF produce an F-value of 4.961, which is statistically significant at less than .05. In this test, the average return differences are both positive. The

[8] To illustrate, the income effect of inventory changes depends, among other things, on the quantity of the ending inventory balance. A change to the equity method of accounting for investments depends on the net income of the investee. Changes in depreciation methods depend on the ages of depreciable assets and current-year acquisitions.

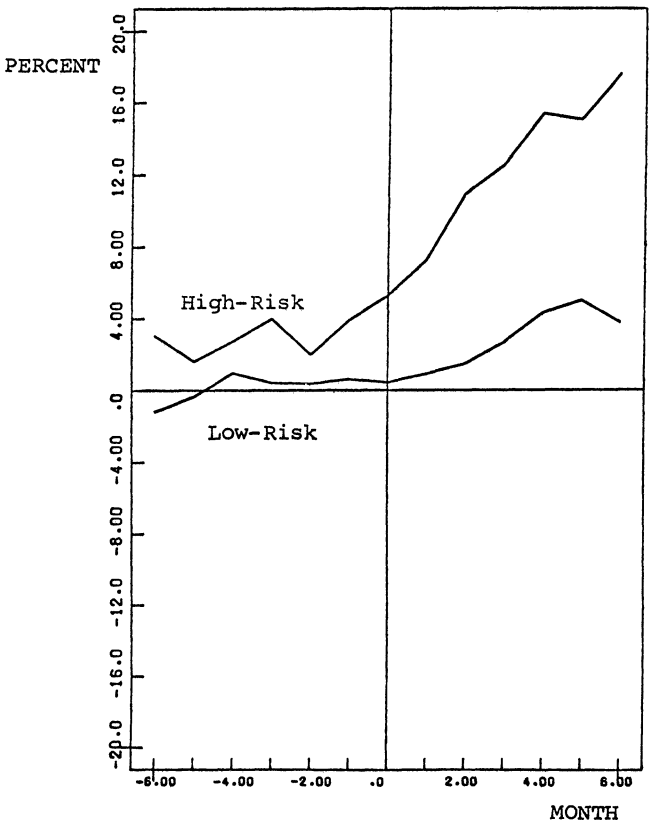

FIG. 3.—Cumulative return difference of ACF (+, N) - NCF; 60 pairs of firms

signs of these return differences are in marked contrast to those in the test of ACF (+, D). Therefore, since the signs of the income effects are the same, the discretion characteristics of the changes appear to be related to the opposite results. Another possible explanation could be that the types of ACs examined in the two tests are different. The results of this test are essentially a test of nondiscretionary changes from cost to equity whereas the test of ACF (+, D) includes several different types of ACs. Unfortunately, I can only speculate about whether any type of bias exists.

In the present test, the demonstrably unique market behavior seems to begin in month $t = -2$ (see fig. 3), which is the fiscal year-end of each firm. While the relative timing of this effect is generally consistent with the advance notice that investors have of nondiscretionary accounting changes, figure 3 reveals substantially more nonzero return differences after year-end than before. But again, to the extent that (a) the net income effect of the change is important to investors and (b) the fourth quarter's results affect the full year net income, the year-end disclosures could convey information to investors.

Recall from table 6 that the largest $\hat{\beta}_g$ difference (between ACF and

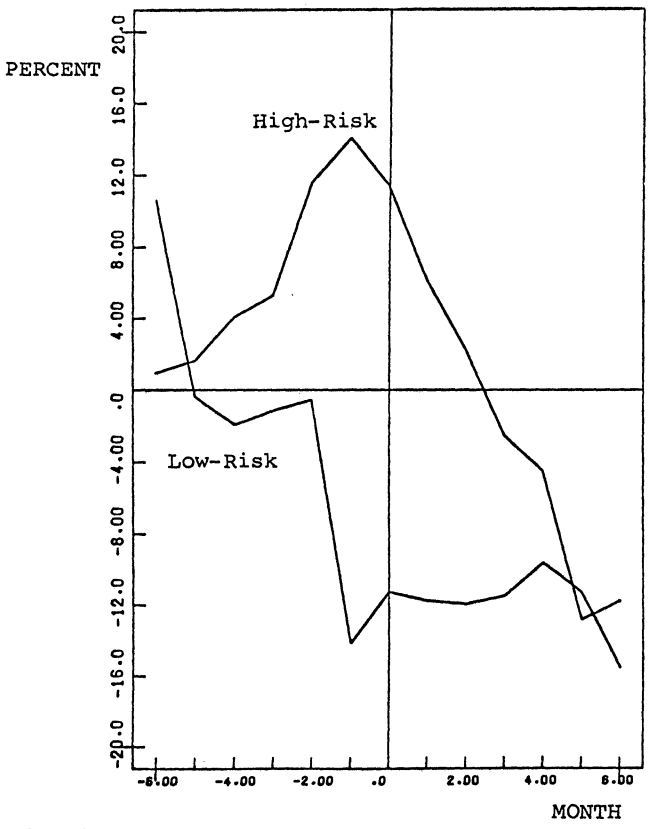

FIG. 4.—Cumulative return difference of ACF (−, N) - NCF; 10 pairs of firms

NCF) existed for the sample firms in this test. Furthermore, for both risk classes, the $\hat{\beta}_g$ of the nonchange firms exceeded the $\hat{\beta}_g$ of the change firms. Consequently, one would expect the returns on NCFs to exceed the returns on ACFs. And since this result did obtain, there exists the possibility that the differences in the (true) $\boldsymbol{\beta}_g$ alone could be sufficient to have rendered the above results spurious.

In order to test whether this was indeed the case, a cross-validation test was performed. The effect of this test was to correct for the effect that differences in the *estimates* of $\boldsymbol{\beta}_g$ might have had on the test results. There were three steps involved. The first step was to inflate the existing $\hat{\boldsymbol{\beta}}_g$ difference (of .02 for high-risk firms and .038 for low-risk firms) to the maximum that a 95 percent confidence interval would allow. Thus, it was necessary to impose a normality assumption on the distribution of β_g. The resulting $\hat{\beta}_g$ differences were .3728 for high-risk firms and .2732 for low-risk firms.

The second step was to multiply these $\hat{\beta}_g$ differences by the weighted average[9] monthly risk premiums (of .00169 for high-risk and .00299 for

[9] The weights used as the proportions of firms in this particular test which made their accounting changes during each of the years 1968–72.

low-risk) that investors actually had available to them during the sample period.[10] (Note that this is a direct application of the capital asset pricing model of (1).) The resulting product constituted the correction factor which was then added to the average return on ACF. Thus, returns on ACF were incremented to their theoretical equilibrium under the assumed return-generating model, thereby reducing the observed return differences.

The third and final step involved performing the T^2 test a second time, using the adjusted returns computed above. The resulting F-value was 4.159. While this is less than the unadjusted F-value of 4.961, it is still statistically significant at less than .05. On this basis, it seems reasonable to conclude that the results of the test of ACF $(+, N)$–NCF are not likely to have occurred solely as a result of $\boldsymbol{\beta}_g$ differences between change and nonchange firms. Furthermore, because the $\boldsymbol{\beta}_g$ difference in the test of ACF $(+, N)$–NCF was the largest of any of the tests conducted in this study, the inference is that the other tests were also substantially unaffected by $\boldsymbol{\beta}_g$ differences.

The two tests involving accounting changes that decreased net income both produced insignificant F-values. Furthermore, the average return differences for the full thirteen-month period were negative in both tests. Thus, it appears that investors assess firms that make ACs which reduce net income, whether discretionary or nondiscretionary, as having essentially the same value as similar firms that do not make ACs. Figures 2 and 4 indicate, however, that the paths of the average return differences in the two tests are quite different. It is likely that these two tests are substantially affected by sampling error due to the small numbers of pairs of firms involved. As a result, additional evidence needs to be examined before more definitive conclusions can be reached about the market effects of AC $(-, D)$ and AC $(-, N)$.

Risk and Information Effects

One of the hypotheses underlying the research design used in this study is that the market may react differently to the ACs made by firms in different risk classes. And because arguments can be advanced to predict widely different results, this poses an interesting empirical question. To assess whether there is some risk dependency in the accounting-change information, the cumulative return differences on both high-risk firms and low-risk firms are plotted in figures 1–4.

In general, the figures seem to indicate some differences. For example, the low-risk firms in the test of ACF $(+, D)$ appear to account for most of the multivariate return difference (see fig 1). In the test of ACF

[10] These weighted average monthly risk premiums used the weights referred to in n. 9 above applied to the realized monthly risk premiums available to investors during August 1968–August 1973. Ibbotson and Sinquefield [1976, p. 37] provide these data, wherein they used the return on the Standard and Poor's 500 Index as the surrogate for R_m and the return on ninety-day U.S. Treasury bills as the surrogate for R_z.

(+, N), the high-risk firms reflect the bulk of the unique market movement (see fig. 3). The results in the remaining two tests are mixed. That is, in both tests, the two risk-class return differences cumulate to essentially the same number, but they take quite different paths in arriving at the same points.

The reasons for predicting risk-class differences in the AC information chiefly concerned predictions of investors' assessments of shifts in the relative risks of one risk class of the accounting change firms. Indeed, in the two tests with significant F-values, the return differences may have been caused by differential (i.e., ACF–NCF) shifts in β_g. And in both tests, this phenomenon appears to have affected one risk class of firms to a greater extent than it did the other. The possibility that these accounting changes are uniquely associated with shifts in the β_i of ACFs provides a motivation for further research.

Both Ball [1972] and Sunder [1975] observed changing $\hat{\beta}_i$ of the change firms in their studies. However, neither presented any evidence on a control group of nonchange firms. Consequently, it is not known whether the $\hat{\beta}_i$ changes they observed were uniquely related to the ACs they examined. The evidence presented here suggests a similar phenomenon which I plan to investigate in a subsequent study.

Summary and Conclusions

The hypotheses that motivated this study are that the relative discretion of management and the sign of the net income effect of an accounting change have information content. In testing these hypotheses, change firms were matched with nonchange firms of similar risk and industry membership. The matched pairs of firms were then divided into high-risk and low-risk groups for the multivariate test.

The test results indicate that both discretionary and nondiscretionary ACs that increase net income are associated with concurrent and unique stock market behavior. However, the negative return differences for discretionary changes, contrasted with the positive return differences for nondiscretionary changes, suggest that the discretion available to management in making the ACs possesses information content.

The tests of discretionary and nondiscretionary ACs that reduced net income both produced insignificant F-values. But the small sample sizes in these two tests render any interpretation of these results tenuous. Replications should be performed on larger samples before conclusions are reached.

The plots of cumulative return differences on the two risk classes of firms examined in this study suggest some risk dependency in the AC information, particularly in the tests with significant F-values. However, the nature of the risk dependency is still far from clear, and for this reason it is probably wise not to make too much of the risk dependency until the nature of the relationship is specifically investigated.

The results in the two tests that supported the information-content hypothesis could be due to either or both of two factors: changes in β_i or differences in the unsystematic component of security returns. If the difference results from changes in β_i, then the ACs would appear to convey value-pertinent information of a more lasting nature. If the difference results from differences in residual returns only, then the AC information would appear to be temporally limited in scope. This, too, will have to await future research.

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