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INFORMATION CONTENT OF ANALYSTS' COMPOSITE FORECAST REVISIONS

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Motivation

In this paper we examine the relationship between revisions in analyst's composite earnings forecasts and contemporaneous stock price movements. Interest in forecasts of accounting earnings has centered around two issues: (1) whether earnings forecasts have information content; and (2) the impact of regulations and policies regarding earnings forecasts on resource allocation (Gonedes, Dopuch, and Penman, 1976). The second issue is substantive only if earnings forecasts contain information. Empirical evidence suggest some relationship between management earnings forecasts and security price behavior (Foster, 1973; Patell, 1976; Nicholas and Tsay, 1979; Penman, 1980). There have also been numerous prouncements by both the SEC and AICPA involving the regulation of management forecast disclosures and guidelines for evaluating the reasonableness of such disclosures.[1] Given the attention afforded historical accounting data by managers, investors, and policymakers, it is not surprising that predictions of accounting data (such as earnings and sales forecasts) are worthy of such attention.

One of the attributes of management forecasts that limits their usefulness is the irregularity with which they are made public and the absence of published forecast revisions when new information causes management to change its earnings prediction. Normally, corporations do not issue public forecasts of accounting data on a regular basis.[2] However, security analysts are in frequent contact with corporations in an effort to confirm information or obtain new information that will enhance their own predictions of accounting data (primarily next year's earnings forecast). Earnings forecasts published by many different brokerage firms and investment services are continuously being updated and sold to investors. It would appear as if these analyst forecasts both complement management forecasts in those periods when management publicly releases forecasts, and act as substitutes in periods when management does not issue public forecasts. The question of interest is whether these analysts' forecasts contain information. If so, what are the implications for accounting policymakers who are attempting to regulate management forecasts? Can management forecasts of earnings be effectively regulated in such an environment? If analysts' forecasts do not contain information, what then are the implications for accounting policymakers and investors who subscribe to forecasting services? Verification of the information content of analysts' forecasts as complements and substitutes for management forecasts could have a pivotal effect on issues related to management forecast regulations and market resource allocation.

Our experience suggests that each corporation, while dealing with the investment community as a whole, has rather close communications with a subset of key analysts or industry specialists. However, analysts issuing forecasts for a given entity are not confined to this subset. As a result of these characteristics, it is not likely that examining the information content of any single source of analysts' earnings forecasts will afford the most powerful test of their information content.

We extend the existing literature on the information content of financial forecasts in two ways. First, we examine the information content of changes in a group forecast produced by 35 independent brokerage firms. This "consensus" forecast is expected to provide a more powerful test of information content than examining any single analysts' forecast.[3] In evaluating these changes in consensus forecasts we control for public announcements made by management regarding actual and predicted earnings in an effort to observe the information content when analysts complement management, and when they substitute for management.

Second, the research investigates the relationship between the magnitude of several information metrics and the magnitude of unsystematic returns. In theory the amount of information, if appropriately measured, should be associated with the magnitude of the unsystematic returns. The previous absense of strong association may have been due to the way information was measured. We examine two information metrics and observe that the previously untested measure of information dominates the more traditional measure. Both extensions strengthen the evidence that earnings forcasts by analysts contain information important to establishing market prices.

METHODOLOGY

Sample Selection and Forecast Data

The sample consisted of revisions in analysts' forecasts of annual earnings per share for 70 companies. These forecasts were obtained from the Investment Brokers Estimate System (I/B/E/S) data base.[4] Each firm satisfied the following criteria:

(1) Monthly rate-of-return data for its common stock on the CRSP tapes during the period April 1970 to December 1978. This data was required for estimating unsystematic returns on each security.

(2) A fiscal year ending on December 31. This criterion ensured that the forecast horizons during any given month were the same for all firms.

(3) At least four analysts provided forecasts for each firm in any given month, and at least one forecast revision occured during the next month.

Only those earnings forecast revisions that were made between April and December of 1977 and 1978 were examined.[5] The forecast revisions from the sample of 70 firms that satisfied these criteria were from 23 industries (Standard Industrial Two-Digit Classification), and represented a sample of 1,002 firm-months of data. Of the 1,002 firm-months, 48 percent were in 1977 and 52 percent were in 1978; 551 represented upward forecast revisions and 451 represented downward forecast revisions. During 1977, there were 258 downward revisions and 226 upward revisions, whereas in 1978 there were 193 downward revisions and 325 upward revisions.[6]

Controlling for Management Disclosures

A potential problem that arises when assessing the information content of a specific type of information disclosure is the release of other information during the test period. One method of controlling for this problem is to delete an appropriate time period surrounding each confounding event (Foster, 1980, p. 55). Two potential confounding events in this study are (1) announcements of quarterly earnings and/or (2) disclosures of earnings forecasts by management. Previous studies have found that these information releases affect security prices (Brown and Kennelly, 1972; Foster, 1973; May 1971; Patell, 1976; Penman, 1980).

The Wall Street Journal Index was examined to determine during which months in the test period management announced quarterly earnings and/or disclosed earnings forecasts. Out of the entire sample of 1,002 firm-months, there were 357 firm-months during which one or both of the above-mentioned confounding information releases occurred. The reduced sample of 645 observations, including 359 upward revisions and 286 downward revisions, was expected to represent instances where analysts were substituting for management forecast revisions. The breakdowns for both the full sample and the subsample are summarized in Table 1.[7]

Information Variables

The distribution of earnings forecasts summarizes the expectations of security analysts regarding future earnings. These expectations are conditional upon the set of information available to analysts at a given point in time. An increase in this information set may result in a different mapping and, hence, in a revision of the earnings forecast distribution. If the information set remains unchanged from one period to the next, there will be no change in the earnings forecast distribution.

A revision in the parameters of the forecast distribution indicates that new information has been received by analysts. Unfortunately, there is no accepted theory that identifies how forecasted accounting earnings should be related to security prices. Empirical results regarding management earnings forecasts, however, suggest that unexpected earnings forecast revisions are associated with security returns (Patell, 1976). We expected similar results for analysts' forecast revisions because the communications between managers and a large group of security analysts faciliate analysts acting as compliments and substitutes for managers' forecast revisions.

We define the unexpected portion of security analysts' earnings forecasts as the change in the mean of the distribution of analysts' annual earnings forecasts from one month to the next. It may be expressed as:

$$\Delta\tilde{f}_{it} = \tilde{f}_{it} - f_{i,t-1} , \qquad (1)$$

where

$\Delta\tilde{f}_{it}$ = the unexpected component of analysts' earnings forecasts for firm i in month t,

$\tilde{f}_{it}$ = the mean of the distribution of security analysts' earnings forecasts for firm i in month t, and

$f_{i,t-1}$ = the mean of the distribution of security analysts' earnings forecasts for firm i in month t-1.

This definition of unexpected forecasted earnings is consistent with the assumption that the average analyst forecast for each month follows a martingale process. If there is no unexpected component, then $E(\Delta\tilde{f}_{it}) = 0$.

The information variables used in this study are the unexpected forecasted earnings component scaled in two different ways. The first information variable (I^1) adjusts only for the level of the forecast and is defined as follows:

$$\tilde{I}_{it}^{1} = \frac{\Delta\tilde{f}_{it}}{|f_{i,t-1}|}, \qquad (2)$$

The second information variable (I^2) adjusts for the consensus amongst analysts and is defined as follows:

$$\tilde{I}_{it}^{2} = \frac{\Delta\tilde{f}_{it}}{s(f_{i,t-1})}, \qquad (3)$$

where

$s(f_{i,t-1})$ = the standard deviation of the distribution of security analysts' earnings forecasts for firm i in period t-1.

Information variable I^1 represents the percentage change in the average earnings forecast from one month to the next. By scaling $\Delta\tilde{f}_{it}$ by the absolute value of the average forecast in period t-1, it is possible to compare forecast revisions across firms and over time. However, defining information variable I^1 in the manner described above poses at least one major problem. The earnings streams of some firms may be more volatile than those of others and therefore more difficult to forecast. This may result in larger forecast revisions for those firms and therefore in higher values for information variable I^1. Scaling the change in average forecasts by the standard deviation of the previous month's forecast distribution amounts to defining information variable I^2 as the number of standard deviations by which the mean forecast was revised from one period to the next. Hence, I^2 deflates the change in the composite forecast by a measure of the relative dispersion (or lack of consensus) about the prediction.

While the comparative accuracy of the forecast metric does not influence the good news/bad news classification of the forecast revision, it does influence the magnitude of the information variable, and hence the results of the association tests. The first information variable is analogous to that used in prior studies.[8] The second metric is used in an effort to provide what might be considered a more appropriate relative information metric across firms. Related research regarding forecast accuracy revealed that the error metric may significantly influence tests of the comparative accuracy of forecast agents (Imhoff and Paré, 1982). Using the same arguments, we suggest that I^2 may be a better relative measure of the amount of information conveyed, since it incorporates a measure of the underlying uncertainty as the deflator term.[9]

Measurement of Unsystematic Returns

Unsystematic returns were generated from a market model,

$$\tilde{R}_{it} = a_i + b_i \tilde{R}_{mt} + \tilde{e}_{it}, \qquad (4)$$

The coefficients (a_i and b_i) were estimated using monthly stock returns for the 84 months preceding the month during which the forecast revisions are made. These estimates, $\hat{a}_i$ and $\hat{b}_i$, were then used to determine monthly unsystematic returns such that:

$$w_{it} = R_{it} - (\hat{a}_i + \hat{b}_i R_{mt}), \qquad (5)$$

where

w_{it} = the estimated unsystematic return on security i in month t.

Incorrect specification of the process that generates security returns may cause unsystematic returns to be biased upwards in some months and downwards in others. If the process that generates security returns is correctly specified, then, for a random sample, the average unsystematic return will not be significantly different from zero. For a nonrandom sample, however, there may be positive or negative average unsystematic security returns during some months. These nonzero average cross-sectional returns may be caused by information other than security analysts' earnings forecast revisions. To avoid attributing these nonzero unsystematic returns to earnings forecast revisions, they are standardized in two ways.

In the first method, monthly unsystematic returns, w_{it}, are adjusted to remove the cross-sectional mean return as follows:

$$x_{it} = w_{it} - \bar{w}_t , \qquad (6)$$

where

x_{it} = the standardized monthly unsystematic return on security i in month t,

$\bar{w}_t = \frac{1}{n_t} \sum_{i=1}^{n_t} w_{it}$,

n_t = the number of firms in the sample during month t, and

w_{it} is as defined in equation (5).

This process ensures that the average transformed return for that month is equal to zero, and that the average x_{it} for the entire sample is zero.

This first method of standardization does not ensure that the variances of each cross-sectional distribution are equal. Moreover, the relative magnitude of the information content should be adjusted for "normal" changes in prices. The second method assumes that both the means and the variances of each cross-section are different. The standardization is performed as follows:

$$y_{it} = \frac{w_{it} - \bar{w}_t}{s_t(w)} , \qquad (7)$$

where

y_{it} = the standardized monthly unsystematic return on security i in month t,

$s_t(w) = [\frac{1}{n_t - 1} \sum_{i=1}^{n_t} (w_{it} - \bar{w}_t)^2]^{1/2}$, and

w_{it}, $\bar{w}_t$ are as defined earlier.

This method transforms the unsystematic returns so that each cross-section and, hence, the entire sample have y_{it} with a mean equal to zero and variance equal to one, thereby providing the properties required for the t-test (Mandelker, 1974).

RESULTS

Tests were conducted to determine if the unsystematic returns for the good news firms (I^K>0) were greater than those of the bad news firms. In addition, tests were run to determine if there was a significant relationship between the magnitude of I^1 and/or I^2 and the magnitude of the unsystematic returns.

Table 2 presents summary descriptive statistics for revisions in average earnings forecasts for both information measures for the full sample and for the subsample. Summary descriptive statistics for the security return regression models during the preforecast period are presented in Table 3.[9] The data for this period consists of monthly return data for the 84 months prior to the month during which forecasts were made.

Tests of Information Content

Table 4 presents summary descriptive statistics for the three measures of unsystematic returns. The average unsystematic returns for the 551 upward-revision firm-months, presented in column 2 of Table 4, are positive for all

Z_{it}. For the 451 downward-revision firm-months reported in column 5, the average unsystematic returns are negative for all three measures. The t statistic is the appropriate statistic for testing whether the return measure for either the upward revision firms or downward revision firms differs from zero. For the upward-revision firms the w_{it} return measure, was not greater than zero at the 0.05 level of significance. However, all other Z_{it} were significantly different from zero (in the expected direction) at the .01 level.

Column 8 of Table 4 presents values of the t-statistics for testing the significance of differences in Z_{it} for the upward and downward forecast revisions. The values presented in column 8 indicate that, for each variation of unsystematic returns (w_{it}, x_{it}, y_{it}), the differences are significant at the .0005 level. These results are consistent with the notion that revisions in security analysts' earnings forecasts have information content despite the potential confounding effects of other information releases that are occurring at the same time, such as management earnings and dividend announcements.

Table 5 presents summary descriptive statistics for the subsample of firm months without management releases for the three measures of unsystematic returns used in the study. The mean unsystematic returns were significantly greater than zero for two of the three Z_{it} for the upward-revision group and significantly less than zero for all three Z_{it} for the downward-revision group. The results for the subsample are similar to those obtained for the whole sample, and are once again consistent with the notion that revisions in security analysts' earnings forecasts contain information and are associated with changes in stock prices. This result is important in that it is consistent with the theory that analysts act as substitutes for management in periods where management does not make a public announcement. If the results were not significant it could have implied measurement problems or a lack of information content. The latter possibility is not appealing since it would suggest that the collective efforts of these analysts in gathering and processing data to develop and publish earnings forecast revisions is futile.

Note that the results for the full sample were stronger but consistent with those of the subsample. In this sense we suggest analysts' forecast revisions complement the news releases of management. If analysts' revisions conflicted with the news released by managers (given the results from the subsample) we would expect the full sample results to be less powerful than those of the subsample.

Cross-sectional Correlation and Time Series Tests

The presence of nonzero cross-sectional correlation (i.e., $cov(w_{it}, w_{jt}) \neq 0$, $i \neq j$) between unsystematic returns causes a bias in the estimate of the standard error. This bias causes tests of significance which assume independence to be biased in a manner that depends upon the nature of the correlation. The tests on the t statistics (reported in columns 4 and 7 of tables 4 and 5) assume independence between unsystematic returns. Lack of independence will cause the results of this test to be biased. To remedy this potential problem, an approach is used that pools unsystematic returns over time but not across firms.

Each month, firms were divided into two groups on the basis of the direction of revision of the analysts' earnings forecasts. Two portfolios were formed each month, one consisting of firms with positive forecast revisions and the other of firms with negative forecast revisions. Each portfolio was held for one month. An investment strategy that consisted of buying shares of firms that had upward revisions and short-selling an equal dollar amount of shares of firms that had downward revisions was employed. Investments were made at the end of the month preceding the month during which forecasts were revised. Each portfolio was held until the end of the forecast revision month. The data consisted of forecast revisions that were made during 18 months.

Let t be a month during which upward (downward) revisions were made for $m_t^1(m_t^2)$ firms. A portfolio consisting of $m_t^1(m_t^2)$ stocks was formed at the end of month t-1. The unsystematic portfolio return during month t was calculated as follows:

$$w_{pt}^k = \frac{1}{m_t^k} \sum_{i=1}^{m_t^k} w_{it}, \qquad k = 1, 2,$$

where

w_{pt}^k = the unsystematic portfolio return during month t,

m_t^k = the number of securities in the portfolio during month t,

k = a superscript representing the upward-revision portfolio when k = 1 and the downward-revision portfolio when k = 2, and

w_{it} is as defined in equation (2).

The monthly unsystematic return in month t, using the investment strategy described earlier, was then calculated as follows:

$$w_{pt} = w_{pt}^1 - w_{pt}^2.$$

The statistic $t = \bar{w}_p \times \sqrt{18}/s(w_{pt})$ (where $\bar{w}_p$ and $s(w_{pt})$ represent the mean and standard deviation, respectively, of the time series of w_{pt}) is distributed as a t-distribution with 17 degrees of freedom. The monthly unsystematic return using our investment strategy was 1.622 percent and the t-statistic was 5.199, a value that is significantly greater than zero at the .0001 level of significance. These results provide support for the earlier tests, and are consistent with the premise that revisions in security analysts' earnings forecasts convey information.

Relationship between Unsystematic Returns and Information Variables

The results reported above are based on the classification of forecast revisions as good news and bad news alone with no recognition for the magnitude of the information measure. We also considered the relationship between the information measure and the market response to be important. In the absence of a relationship between the magnitude of information metric and the market response one could argue that the information (as defined) is not driving the results but is simply a reasonable proxy for something else that is responsible for the results.

First we examined the general nature of the relationship. Spearman rank-order correlation coefficients were computed to examine the ordinal relationship between unsystematic returns and information variables I^1 and I^2. Table 6 reports Spearman rank-order correlation coefficients for the full sample (of 1,002 firm-months) and the subsample, in Panels A and B respectively. For all three measures of unsystematic returns, there is a positive relationship with both I^1 and I^2.[11]

Figures in parentheses in Table 6 represent t values. These values enable the null hypothesis (r_s=0) to be rejected at the 0.05 level of significance for all three unsystematic return measures. The results indicate a positive rank-order correlation between the magnitude of unsystematic returns and the magnitude of information variables I^1 and I^2, adding to the evidence suggesting that both information measures capture "news" that is reflected in security prices. Note that I^2 is more highly correlated with the unsystematic returns in every form, thereby dominating I^1.

We also examined the possibility of a linear relationship between the information metrics and unsystematic returns. Table 7 reports results of pooled time-series and cross-sectional regressions. The null hypothesis underlying the regressions is that the coefficient of the information variable is less than or equal to zero. In Table 7, t values are given in parentheses below the estimates of the regression coefficients. In panel A, for information variables I^1 and I^2, the t values for $\hat{\beta}$ are large enough to reject the null hypothesis at the .001 level of significance. Regression tests for the subsample (reported in Table 7 Panel B) provided similar results.

The relationship between the magnitude of each I^k_{it} observation and its corresponding error term (ε^k_{it}) were evaluated for heteroscedasticity and serial correlation. We observed significant heteroscedasticity problems in most cases for I^1 but only in one case for I^2, suggesting that the linear relationship is most appropriate when applied to I^2.[12] No serial correlation was observed with either metric.[13] These results suggest that while there is a significant positive linear relationship between unsystematic security returns and both information variables, I^2 dominates I^1 and the assumptions of linearity are also more appropriate in the case of I^2.

The results reported in Tables 6 and 7 enable us to draw several general conclusions. First, the significance of the results for I^2 is uniformly stronger than that for I^1 for all measures of Z_{it}. The I^2 variable was expected to provide a stronger mapping into price revisions and its dominance over I^1 seems to be consistent with our expectations. The use of a measure of analysts' forecast variability as a relative deflator produced results that dominated the more common information measure (I^1) in associations with unsystematic security returns. These empirical results suggest that the information metric influences the results, and that I^2 may be a more powerful measure of the unexpected portion of the forecast.

The three measures of Z_{it} provided results that generally support the use of some standardization of residuals. The standardized measures (x_{it} and y_{it})

generate more significant associations with the information measures than the unstandardized returns (w_{it}). The standardized measure that adjusts the mean residual to zero (x_{it}) generates the highest t statistics in most instances (see Tables 6 and 7), but the differences between x_{it} and y_{it} are slight.

SUMMARY

Accounting data are of great concern to preparers, users, and regulators in both historical and prospective form. Rules governing the preparation and dissemination of both are of continuous interest to the profession. Recently, prospective data have received a great deal of attention from both the SEC and the AICPA. Results presented here demonstrate that disclosure rules should consider analysts as well as management forecasts in formulating policy since analysts forecasts contain information which complements and substitutes for management disclosures regarding future earnings. Moreover, we found a significant relationship between the <u>magnitude</u> of unsystematic security returns and the <u>magnitude</u> of forecast information variables.

The results of the tests for information content revealed that upward revisions in average predicted earnings are accompanied by positive unsystematic returns and downward revisions by negative unsystematic returns. It appears that positive revisions are perceived by investors as conveying favorable information about firms' production-investment activities, and therefore cause investors to bid up the prices of these firms' shares of stock. The converse is true for firms with negative earnings forecast revisions.

The reported results were significant for the subsample as well as for the full sample, suggesting that revisions in analysts' forecasts both complement and substitute for the disclosures of management. The results of the association tests revealed that the forecast information variable which was deflated by the variability in analysts' forecasts (I^2) dominated the more commonly used information measure. While both information measures were significantly associated with returns, the I^2 measure explained a greater portion of the unsystematic security returns. The fact that we observed a significant linear relationship between the magnitude of the I^2 forecast information metric and the magnitude of the unsystematic returns, and that the linear form of the relationship appears to be descriptive effectively offers a new basis for future research until a better theory is developed.

FOOTNOTES

[1]The SEC has been actively involved in forecast disclosure policies since 1972. The AICPA has issued a Guide, a Statement of Position, and an SAS (forthcoming), all since 1980. For a discussion of these pronouncements and their current and potential impact on the accounting profession, see Danos and Imhoff, 1983.

[2]Much speculation exists as to why corporate forecasts are issued on such an irregular basis. Two explanations provided by corporate managers for their public forecasts are: 1) to correct for inaccurate expectations (either too optimistic or too pessimistic) of market participants due to "noise" created by other information (i.e. analysts forecasts, contract negotiations, labor disputes, material shortages, etc.), and; 2) because of unplanned responses by corporate officers to questions raised by the news media at press conferences where other news (such as quarterly sales or earnings results) was to be announced. In the second situation the unplanned releases were sometimes made against the forecast disclosure policies of the companies involved.

[3]The relevance of using multiple forecasts to represent a more appropriate signal is well documented in the literature (Beaver, 1981a). Also, recent Wall Street Journal articles have noted that consensus forecasts consistently outperform individuals (April 6, 1983, p. 48, November 3, 1981, p. 31). Findings also suggest that composite estimates based on small groups (such as three or four) are able to achieve most of the incremental representativeness provided by composite estimates of larger groups (Einhorn, Hogarth, and Klempner, 1977).

[4]The I/B/E/S data base contains monthly summary data, including the mean forecast, the standard deviation, and a list of the individual estimates of earnings per share made by approximately 35 contributing institutional brokerage firms. Forecasts are made for several hundred publicly held corporations. Each brokerage firm estimates per-share earnings for a large subset of the total number of companies covered by the I/B/E/S data base. The number of forecasts reported for each company in any given month may vary between 1 or 2 and 35. Brokerage firms revise their forecasts frequently. The total number of revisions by all the brokerage firms varies from month to month and from firm to firm.

[5]Revisions made between January and March were not examined, since "forecasts" for December 31 firms during the early part of the calendar year referred to the previous year's earnings. The majority of forecasts made in January and February 1977 (1978) are for the fiscal year ending December 1976 (1977).

[6]A test of the null hypothesis of an equal probability of an upward and a downward revision was conducted using the binomial test. The null hypothesis was rejected at the 5 percent significance level for forecast revisions during 1978 and for the entire sample. These annual differences in the distribution of good news and bad news are not unusual, and they say nothing about the magnitude of the "news," which is accounted for in the information metric itself.

[7]Note that most WSJ entries were for earnings announcements. Public forecasts by management are seldom systematically released. The fact that few management forecasts are published and fewer still without accompanying actual earnings and dividend announcements prohibited us from conducting a better test of the complementing effect of analysts' forecasts.

[8]Patell (1976), Nichols and Tsay (1979), and Beaver, Clarke and Wright (1979) use information variables defined as the percent deviation of a forecast from the mean forecast obtained from models based on the past actual earnings series, which is similar to I^1.

[9]The I^2 measure deflates by uncertainty given the time period over which information is measured. It is conceivable that within each period (one month), analysts agree completely to two different forecasts at two different points in time. Operationally, this is simply not the case.

[10]If $\tilde{R}_{it}$ and $\tilde{R}_{mt}$ are bivariate normal, there will be a linear relationship between them and

$$E(\tilde{R}_{it}/R_{mt}) = a_i + b_i R_{mt},$$

such that

$$E(\tilde{e}_{it}/R_{mt}) = 0.$$

Ordinary least squares regression on the time series of $\tilde{R}_{it}$ and $\tilde{R}_{mt}$ was used to estimate the coefficients a_i and b_i in equation (1). This estimation procedure assumes that:

(i) a_i and b_i are constant over time,

(ii) $E(\tilde{e}_{it}) = 0$,

(iii) $\sigma(\tilde{e}_{it}, \tilde{R}_{mt}) = 0$, and

(iv) $\sigma(\tilde{e}_{is}, \tilde{e}_{it}) = \begin{cases} 0 & \text{for } s \neq t, \\ \sigma_i^2 & \text{for } s = t. \end{cases}$

[10]The mean $\hat{\beta}$ of 1.18 results from using a value-weighted market index to estimate the individual $\hat{\beta}$'s. Had an equal-weighted index been used, the mean $\hat{\beta}$ would have been lower. For example, Beaver, Clarke, and Wright (1979) report a mean β of 1.203 when a value-weighted index was used and a mean β of 0.950 when an equally weighted market index was used.

[11]To interpret the significance of the observed rank-order correlation coefficients, certain assumptions about the distributional properties of unsystematic returns and information variables must be satisfied. If each joint observation of Z_{it} and I^k_{it} is independent and drawn from the same distribution, then for large samples ($N \geq 10$), the statistic $t = r_s[\frac{N-2}{1-r_s^2}]^{1/2}$ is distributed as a Student's t with N-2 degrees of freedom (Siegel, 1956, p. 212). r_s represents the Spearman rank-order correlation coefficient and N is the sample size. To the extent that these assumptions are satisfied, tests of significance will be unbiased.

[12]A Goldfeld-Quandt test (Goldfeld and Quandt, 1965) was used to examine the homoscedasticity of the 12 relationships (six each for I^1 and I^2, each with errors from w, x, and y for both the full sample and subsample). For I^1, only the errors from y_{it} for the subsample failed to reject the null of homoscedasticity. For I^2, all but the errors from w_{it} failed to reject the null.

[13]The Durbin-Watson statistics for the 12 relationships ranged from 1.96 to 2.08, suggesting no serial correlation between the magnitude of the I^k_{it} and their corresponding residuals generated by w, x, and y.

REFERENCES

Ball, R. and P. Brown, "An Empirical Evaluation of Accounting Income Numbers," Journal of Accounting Research (Autumn 1968), pp. 159-178.

Beaver, W. H., Financial Reporting: An Accounting Revolution (Englewood Cliffs, N.J.: Prentice Hall, 1981a), pp. 160-163.

________, "Econometric Properties of Alternative Security Return Methods," Journal of Accounting Research (Spring 1981b), pp. 163-184.

________, R. Clarke and W. F. Wright, "The Association between Unsystematic Security Returns and the Magnitude of Earnings Forecast Errors," Journal of Accounting Research (Autumn 1979), pp. 316-340.

Brown, P. and J. W. Kennelly, "The Information Content of Quarterly Earnings: An Extension and Some Further Evidence," Journal of Business (July 1972), pp. 403-415.

Danos, P. and E. A. Imhoff, Jr., "Factors Affecting Auditors' Evaluations of Forecasts," Journal of Accounting Research (Autumn 1983) in press.

Einhorn, H. J., R. M. Hogarth and E. Klempner, "Quality of Group Judgment," Psychological Bulletin (1977), pp. 158-172.

Foster, G., "Stock Market Reaction to Estimates of Earnings per Share by Company Officials," Journal of Accounting Research (Spring 1973), pp. 25-37.

________, "Accounting Policy Decisions and Capital Market Research," Journal of Accounting and Economics (March 1980), pp. 25-62.

Givoly, D. and J. Lakonishok, "The Information Content of Financial Analysts' Forecasts of Earnings," Journal of Accounting and Economics (January 1979), pp. 165-185.

Goldfeld, S. M. and R. E. Quandt, "Some Tests of Homoscadicity," Journal of the American Statistical Association (June 1965), pp. 539-547.

Gonedes, N. J., "Risk, Information, and the Effects of Special Accounting Items on Capital Market Equilibrium," Journal of Accounting Research (Autumn 1975), pp. 220-256.

________ and N. Dopuch, "Capital Market Equilibrium, Information Production and Selecting Accounting Techniques: Theoretical Framework and Review of Empirical Work," Studies on Financial Accounting Objectives: 1974, Supplement to Journal of Accounting Research, Vol. 12, pp. 48-129.

________, N. Dopuch and S. H. Penman, "Disclosure Rules, Information-Production and Capital Market Equilibrium: The Case of Forecast Disclosure Rules," Journal of Accounting Research (Spring 1976), pp. 89-137.

Imhoff, E. A. and P. Paré, "Analysis and Comparison of Earnings Forecast Agents," Journal of Accounting Research (Autumn 1982), pp. 429-439.

Mandelker, G., "Risk and Return: The Case of Merging Firms," Journal of Financial Economics (1974), pp. 331-333.

May, R., "The Influence of Quarterly Earnings Announcements on Investor Decisions as Reflected in Common Stock Price Changes," Empirical Research in Accounting: Selected Studies, 1971, Supplement to Journal of Accounting Research, Vol. 9, pp. 119-163.

Mood, A. M., F. A. Graybill and D. C. Boes, Introduction to the Theory of Statistics (New York: McGraw Hill, Inc., 1974).

Nichols, D. R. and J. J. Tsay, "Security Price Reactions to Long-Range Executive Earnings Forecasts," Journal of Accounting Research (Spring 1979), pp. 140-155.

Patell, J. M., "Corporate Forecasts of Earnings per Share and Stock Price Behavior: Empirical Tests," Journal of Accounting Research (Autumn 1976), pp. 246-276.

Penman, S. H., "An Empirical Investigation of the Voluntary Disclosure of Corporate Earnings Forecasts," Journal of Accounting Research (Spring 1980), pp. 132-160.

Securities and Exchange Commission, Release No. 5992 (Securities Act of 1933), (SEC, November 7, 1978).

________, Release No. 6084 (Securities Act of 1933); Release No. 15944 (Securities Act of 1934), (SEC, June 25, 1979).

________, Release No. 6413 (Securities Act of 1933), (SEC, June 24, 1982).

Siegel, S., Nonparametric Statistics (New York: McGraw Hill Book Company, 1956).

The Wall Street Journal, November 3, 1981, p. 31.

________, April 6, 1983, p. 48.

Winkler, Robert L., "The Quantification of Judgment: Some Experimental Results," Proceedings of the American Statistical Association, 1967, pp. 386-395.

________, "Probabilistic Prediction: Some Experimental Results," Journal of the American Statistical Association (December, 1971), pp. 675-685.

TABLE 1

DESCRIPTION OF SAMPLE

		Full Sample	Subsample[a]
1977	Upward Revisions	226	148
	Downward Revisions	258	165
1978	Upward Revisions	325	211
	Downward Revisions	193	121
	Totals	1,002	645

[a]Excluding firm-months during which there were quarterly earnings announcements and/or earnings forecast disclosures by management.

TABLE 2

DESCRIPTIVE STATISTICS FOR INFORMATION VARIABLES

		Full Sample n=1002	Subsample n=645
I^1	Minimum	-0.3536	-0.3536
	Maximum	0.2547	0.2547
	Mean	- .0036	- .0038
	Standard Deviation	0.0428	0.0453
I^2	Minimum	-2.5115	-2.5115
	Maximum	2.8962	1.3543
	Mean	.0201	.0198
	Standard Deviation	0.4110	0.4330

TABLE 3

DESCRIPTIVE STATISTICS FOR SECURITY RETURN REGRESSION MODELS

		Full Sample	Subsample
$\hat{\beta}$	Minimum	0.4728	0.4728
	Maximum	2.1517	2.1131
	Mean	1.1755	1.1787
	Standard Deviation	0.3240	0.3240
R^2	Minimum	0.1228	0.1254
	Maximum	0.6695	0.6695
	Mean	0.3782	0.3784
	Standard Deviation	0.1280	0.1300

TABLE 4

SUMMARY STATISTICS FOR UNSYSTEMATIC SECURITY RETURNS ON PORTFOLIOS FORMED CONDITIONAL UPON THE DIRECTION OF FORECAST REVISION

	Upward-Revision Firms--H_0^1 n=551			Downard-Revision Firms--H_0^2 n=451			H_0^3
(1) Z_{it}	(2) Mean ($\bar{Z}_1$)	(3) Standard Deviation (S_1)	(4) $t_{\bar{Z}_1}$	(5) Mean ($\bar{Z}_2$)	(6) Standard Deviation (S_2)	(7) $t_{\bar{Z}_2}$	(8) $t_{\bar{Z}_1-\bar{Z}_2}$
w_{it}	.0026	0.0599	1.0200	-.0113	0.0589	-4.0610	3.6730
x_{it}	.0069	0.0592	2.6920	-.0082	0.0583	-3.0260	4.0430
y_{it}	0.1156	0.9740	2.7880	-0.1413	0.9960	-3.0130	4.1130
	Estimated average systematic risk for firms with $I_{it} > 0 = 1.2022$ and for $I_{it} < 0 = 1.1429$.						

Notes:

$$t_{\bar{Z}_1-\bar{Z}_2} = \frac{[n_1 n_2/(n_1+n_2)]^{1/2}(\bar{Z}_1-\bar{Z}_2)}{[\{(n_1-1)S_1^2+(n_2-1)S_2^2\}/(n_1+n_2-2)]^{1/2}},$$ where n_1 and n_2 are the

number of firm-months during which $I_{it} > 0$ and $I_{it} < 0$, respectively; Z_1 and Z_2 are the mean unsystematic returns for firm-months during which $I_{it} > 0$ and $I_{it} < 0$, respectively; S_1 and S_2 are the standard deviations of the unsystematic returns for firm-months during which $I_{it} > 0$ and $I_{it} < 0$, respectively.

TABLE 5

SUMMARY STATISTICS FOR UNSYSTEMATIC SECURITY RETURNS ON PORTFOLIOS FORMED CONDITIONAL UPON THE DIRECTION OF FORECAST REVISION (SUBSAMPLE)

	Upward-Revision Firms--H_0^1 n=359			Downard-Revision Firms--H_0^2 n=286			H_0^3
(1) Z_{it}	(2) Mean ($\bar{Z}_1$)	(3) Standard Deviation (S_1)	(4) $t_{\bar{Z}_1}$	(5) Mean ($\bar{Z}_2$)	(6) Standard Deviation (S_2)	(7) $t_{\bar{Z}_2}$	(8) $t_{\bar{Z}_1-\bar{Z}_2}$
w_{it}	-.0006	0.0587	-0.1880	-.0118	0.0588	-3.3820	2.4010
x_{it}	.0055	0.0582	1.7990	-.0074	0.0580	-2.1520	2.8020
y_{it}	.0950	0.9520	1.8900	-0.1261	1.0000	-2.1250	2.8600

Estimated average systematic risk for firms with $I_{it} > 0 = 1.2089$ and for $I_{it} < 0 = 1.1407$.

TABLE 6

SPEARMAN RANK-ORDER CORRELATIONS BETWEEN UNSYSTEMATIC SECURITY RETURNS AND INFORMATION VARIABLES

Panel A: Full Sample

Z_{it} / I^k	w_{it} r_s (t value)	x_{it} r_s (t value)	y_{it} r_s (t value)
I^1	0.1020* (3.2424)	0.1190* (3.7900)	0.1180* (3.7577)
I^2	0.1120* (3.5642)	0.1260* (4.0165)	0.1240* (3.9517)

Panel B: Subsample

Estimates based on 645 observations.

Z_{it} / I^k	w_{it} r_s (t value)	x_{it} r_s (t value)	y_{it} r_s (t value)
I^1	0.1020* (3.2424)	0.1190* (3.7900)	0.1180* (3.7577)
I^2	0.1120* (3.5642)	0.1260* (4.0165)	0.1240* (3.9517)

*Significant at $\alpha \leq .05$.

TABLE 7

SUMMARY OF POOLED TIME-SERIES AND CROSS-SECTIONAL REGRESSIONS

Equation: $Z_{it} = \alpha^k + \beta^k I^k_{it} + \varepsilon^k_{it}$, k=1,2

Panel A: Full Sample

	Z_{it}	$\hat{\alpha}$ (t value)	$\hat{\beta}$ (t value)
I^1	w_{it}	-.0031 (-1.65)	0.1473 (3.35)
	x_{it}	.0006 (0.29)	0.1566 (3.60)
	y_{it}	.0093 (0.29)	2.6158 (3.59)
I^2	w_{it}	-.0041 (-2.16)	.0207 (4.54)
	x_{it}	-.0004 (-0.23)	.0214 (4.75)
	y_{it}	-.0073 (-0.23)	0.3609 (4.78)

Panel B: Subsample

	Z_{it}	$\hat{\alpha}$ (t value)	$\hat{\beta}$ (t value)
I^1	w_{it}	-.0052 (-2.24)	.0816 (1.59)
	x_{it}	.0002 (0.06)	.0935 (1.84)
	y_{it}	.0031 (0.07)	1.6100 (1.88)
I^2	w_{it}	-.0058 (-2.52)	.0154 (2.89)
	x_{it}	-.0005 (-0.23)	.0166 (3.14)
	y_{it}	-.0086 (-0.22)	0.2776 (3.13)