Market Demand for Conservative Analysts

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Abstract

Sell-side analysts, on balance, have incentives to emphasize good company news and downplay the bad, resulting in inefficient forecasts. We conjecture that this behavior generates a demand for forecasts from conservative analysts who unwind this pattern, at least in part, resulting in more efficient forecasts. To investigate, we introduce a measure of analyst conservatism and assess the market reaction to analysts' forecast revisions conditioned on their past levels of conservatism. We find a stronger market reaction to forecast revisions by more conservative analysts, and that this result is heightened for companies with greater institutional investor following.

JEL Classification: G14; G24; M4.

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1. INTRODUCTION

As information intermediaries, sell-side analysts gather and process public and private company information for the benefit of client investors who lack the resources to conduct such analyses on their own. The research of analysts, therefore, is geared towards first creating an information asymmetry between analysts and client investors, and then, ideally, reducing this asymmetry through systematic and objective disclosures. However, analysts, on balance, tend to emphasize good and downplay bad company news, resulting in inefficient forecasts (Easterwood and Nutt, 1999). While many incentives documented in prior work are consistent with the norm of aggressive research (Hong and Kubik, 2003; D'Avilio et al., 2002; Francis and Philbrick, 1993; Das et al., 1998; among others), we argue that some analysts will forecast conservatively to garner influence with equity investors, who should value more efficient earnings research.[1] Accordingly, we conjecture that equity investors in general—and institutional investors in particular—will respond more strongly to research from conservative analysts, who generate more informative forecasts by unwinding, at least in part, the aggressive research practices representative of analysts more generally.[2]

Since it is not clear if analysts act conservatively in an absolute sense, we introduce a conservatism measure that is conditional (based on an individual analyst's reactions to news direction, bad versus good) and relative (based on an individual analyst's reactions relative to those of peer analysts).[3] This measure is formed each calendar year by ranking, among the analyst population, the average asymmetry in an individual analyst's forecast revisions to bad

[1] Our study focuses on analysts' earnings research due to its importance to market participants in gauging firm performance (Brown, 1993), and its effect on other analyst research outputs such as target price forecasts (Bandyopadhyay et al., 1995; Bradshaw, 2002; Bradshaw, 2004; Gleason et al., 2008) and stock recommendations (Loh and Mian, 2006; Ertimur et al., 2007).
[2] In addition to institutional investor attention, analyst research quality has been shown to help attract underwriting business, Krigman et al. (2001), as well as generate trade, Jackson (2005).
[3] For ease of exposition, we refer to the 'conditional and relative conservatism' as simply 'conservatism.'

versus good news; an analyst making stronger revisions in response to bad news versus good news relative to her peers is considered more conservative.[4] After this modeling, we examine how analyst characteristics vary with the conservatism measure. We find that more conservative analysts are more likely to be *Institutional Investor* award winners, work for larger investment houses, have more experience, and yield forecasts that are more accurate and more persistent, i.e., more predictive of future earnings revisions. Collectively, these findings suggest that more conservative analysts are more able, have better resources, and provide more efficient forecasts than less conservative analysts.

In order to examine the market response to conservative analysts' earnings research, we assess the short-window market responses to analysts' forecast revisions in the testing period (i.e., current year) conditioned on their level of conservatism established during the estimation period (i.e., prior year). We find a stronger market response to forecast revisions of conservative analysts; the response is 22% larger for the revisions of the top versus the bottom conservatism quintile. Extending the analysis to the setting of institutional ownership, we conjecture and find that institutions' fiduciary duties coupled with their greater ability to discern efficient earnings research result in a relatively stronger market response to conservative analysts' forecasts.

To further the interpretation of our results, we conduct two additional analyses. The first tests a prediction of Bayesian investor learning that analysts develop a more precise reputation for conservatism as their estimation period lengthens, thereby allowing the market to observe a

[4] There are several reasons underlying our introduction of this measure. First, most prior work on related concepts, such as pessimism, rely on realized earnings. The primary concern with such an approach is that these measures are based on *ex post* realizations, which are subject to management discretion, as well as firm and industry shocks— all of which are beyond the analyst's control. We avoid the *ex post* problem by explicitly modeling the analyst's forecast revisions throughout a calendar year in relation to news surrounding the revisions. This modeling also enables us to observe analysts' reactions to company news multiple times during the year, and therefore to form a clearer impression of their forecasting patterns. Second, while a relative *ex ante* pessimism measure may overcome the *ex post* problem as well, such a measure can only reveal the extent of the universal downward bias of an analyst and does not consider an analyst's asymmetric response to *bad versus good news*. We argue that what is more relevant to investors is how an analyst's earnings expectation changes in response to bad relative to good news, not the analyst's forecasting bias irrespective of news content.

greater number of forecasts (Chen et al., 2005). We find that the market reaction to conservative analysts' forecasts monotonically increases as the estimation period is set longer, from one to four years. The second analysis investigates whether the market response to our measure of analyst conservatism aggregates to the brokerage level; we find stronger market responses to revisions of analysts employed by more conservative brokers.[5]

Several efforts towards enhancing the reliability of our measure and robustness of our findings are made. Among them, we verify that the conservatism measure is not subsumed by alternative constructs such as brokerage size, *Institutional Investor* award status, past forecast accuracy, or past bias. Also, altering design choices, such as varying the requirements of our conservatism measure, or substituting a stock returns-based news proxy for our analyst-based news proxy, yield inferentially similar results.

Our findings offer several insights to the empirical literature on the economics of financial intermediation. First, we document a significant cross-sectional variation in individual analysts' asymmetric response to bad versus good news—which by definition forms the degree of analysts' conservatism—and show a stronger market response to more conservative analysts' earnings research. This response is rational in the sense that earnings forecasts of more conservative analysts are *ex post* more accurate and more informative about longer-term earnings revisions. While prior literature has extensively documented the benefits to producing aggressive research, our findings of greater price impact for conservative analysts' earnings research suggest the existence of an alternative incentive—market influence—to produce conservative research.

Second, we find a stronger market response to conservative analysts' earnings research in the presence of greater institutional ownership. Institutional investors shape analysts' careers in

[5] We use the term "brokerage" or "broker" to refer to a broad array of research firms who are more precisely described as full-service investment banks, underwriters, brokers, or pure research firms (Cowen et al., 2006).

the sense that they vote for the best analysts in the annual *Institutional Investor* survey (Stickel, 1992). In addition, institutions pay for research either directly or indirectly, when they allocate their trading across brokerage firms, thereby generating trading commissions for the analysts' employers (Ljungqvist et al. 2007). Our empirical results are consistent with more conservative analysts' research having enhanced credibility in this important setting, as well as institutional investors' ability to discern efficient earnings research.

We organize our paper as follows. The next section describes the related literature and develops our hypotheses. Section 3 describes our conservatism measure and specifies empirical tests. Section 4 explains our sample selection and provides results, and section 5 concludes.

2. RELATED LITERATURE AND HYPOTHESES DEVELOPMENT

We conjecture that the information asymmetry between analysts and investors coupled with analysts' incentives to produce aggressive research induces a demand for the more efficient earnings expectations of conservative analysts. We first outline the economic rationale underlying the existence of a demand for conservative analysts and then extend our reasoning to institutional ownership, where we predict a stronger demand for conservative analysts.

2.1. The market demand for conservative analysts

Easterwood and Nutt (1999) examine competing hypotheses for why analysts fail to incorporate new information efficiently into their forecasts. In particular, they investigate whether analyst behavior is more descriptive of under-reacting to news, over-reacting to news, or reacting to news conditional on its nature—good versus bad. Consistent with the latter description, their findings indicate that analysts generally over-react to good news and under-react to bad news. Easterwood and Nutt (1999) suggest that incentives for this conditional aggressiveness—empirically supported in related work—include career concerns at brokerages geared toward gaining underwriting business through promoting stocks (Hong and Kubik, 2003; D'Avilio et al., 2002) and reliance on maintaining positive management relations (e.g., Francis and Philbrick, 1993; Das et al., 1998).[6]

We conjecture that equity investors can benefit in at least two ways from conservative analysts who unwind the aggressiveness, which is representative of the analyst population. First, conservative analysts' forecasts should be more accurate since these forecasts mitigate the inefficiency due to upward bias generally found in analysts' forecasts (Easterwood and Nutt, 1999). Second, more conservative analysts' forecast revisions should be more persistent in the sense that they better explain future company earnings beyond one year.[7] The above discussion serves as a basis for our first hypothesis.

H1: The market will react more strongly to the forecast revisions from more conservative analysts.

Although conservative analysts' forecasts can benefit investors—via efficiency gains resulting in improved forecast accuracy and persistence—this prompts a supply-side question. How does conservative forecasting benefit the analyst? Analysts who produce better quality research (e.g., more accurate forecasts) have been found to attract new underwriting business (Krigman et al., 2001), generate trade (Jackson 2005), and garner votes to the *Institutional Investor* awards (Ljungqvist et al., 2007). Groysberg et al. (2008) provide evidence of a positive association between each of these outcomes and analyst compensation.

[6] Even in the post Regulation FD environment, both anecdotal and empirical evidence point to the continuing importance of analysts' favorable views in maintaining management relations (Greenberg, 2007; Mayew, 2008).

[7] Prior empirical financial work has shown that the price response to earnings news is positively related to its persistence (Kormendi and Lipe, 1987; Easton and Zmijewski, 1989; among others), and the forecasting literature documents greater forecast efficiency associated with analyst adjustments for the persistence of earnings in their forecasts (Gu and Chen, 2004; Ettredge et al., 1995). We provide evidence on the link between conservatism and persistence (and accuracy) in Section 4.5.

The mentioned benefits of conservatism raise a natural question. If conservatism is beneficial to analysts, then, in equilibrium, why are not all analysts conservative or at least more conservative?[8] We believe the answer to this question involves the trade-offs between the incentives for being aggressive versus those for building a conservative reputation. For example, analysts at investment firms that garner underwriting business through promoting stocks will rationally act aggressively, assuming they have career concerns. Alternatively stated, analysts employed at firms who do not use a reputation strategy to gain underwriting and trading business are unlikely to be compensated for building a conservative reputation. Still other analysts may be aggressive if they rely heavily on access to private information and, thus, are in greater need of currying favor with management. In particular, analysts who lack the human capital necessary to conduct independent analysis may find it difficult to sacrifice access to management and, therefore, will need to maintain positive management relations. These scenarios and others reasonably lead to cross-sectional variation in levels of conservatism.

2.2. The market demand for conservative analysts by institutional investors

Our second hypothesis posits that institutional investors will react more strongly to earnings forecasts of more conservative analysts for two reasons. First, relative to retail investors, institutional investors are more capable of evaluating relevant dimensions of brokerage research (Malmendier and Shanthikumar, 2007; Ljungqvist et al., 2007). As one example, Malmendier and Shanthikumar (2007) show that small traders literally follow analysts' stock recommendations, while large traders adjust their trading response to known biases. Among the factors contributing to such ability of institutional investors are the greater resources and

[8] Of course, not all analysts can be conditionally conservative by our construction, which focuses on relative conservatism among the analyst population. However, the more pertinent question is why do we observe substantial cross-sectional variation in conditional conservatism? In other words, why would analysts not forecast in a uniformly conservative manner to render our conservatism measure negligible?

investment experience associated with the management of larger portfolios (Cready, 1988; Brous and Kini, 1994).

Second, institutional investors consider the legal implications of their investment decisions due to their fiduciary responsibilities (Del Guercio, 1996; Badrinath et al., 1989; Gompers and Metrick, 2001); therefore, we reason that they will also consider carefully the nature of their investment inputs, i.e., earnings expectations. For example, institutions are known to rely on analyst research to satisfy standards of fiduciary responsibility and have cited the use of analyst reports "as evidence of care and prudence" (O'Brien and Bhushan, 1990, p. 56). Overall, we reason that, relative to retail investors, institutions possess increased resources and economic incentives to seek more conservative inputs for evaluating firm performance. This discussion serves as a basis for our second hypothesis.

H2: Institutional investors will react more strongly to the forecast revisions from more conservative analysts.

3. SPECIFICATION OF EMPIRICAL TESTS

3.1. Conservative analyst measure

We view conservative analysts as those who unwind—at least in part—the asymmetric weighting in favor of good news representative of analysts generally. In the spirit of the setup in Basu (1997), we quantify an individual analyst's conservatism annually by regressing the analyst's forecast revisions during a year on the news, an indicator for negative news, and the interaction between the news and the negative news indicator. We measure the news by the average revision of two neighboring analysts, one forecasting within one trading week before and the other one week after the analyst's revision.[9] When there is more than one preceding or

[9] Since the specific events driving analyst forecast revisions are unknown, we opt to use revisions of the analysts' neighbors as a proxy for news. That is, we utilize an output-related measure, the actual actions of other analysts, rather than attempting to predict which events are most relevant to financial analysts. In a supplemental analysis, however, we verify that our results are qualitatively similar if we use either the past one-month or past one-quarter

succeeding revisions from neighboring analysts, we use the forecast revision closest to the revision of interest. In order to compute the forecast revisions from the analyst trio (i.e., the analyst of interest, the preceding analyst, and the succeeding analyst), we employ the same consensus benchmark, which is the average of all outstanding forecasts issued within 30 days before the forecast of the preceding analyst. This design allows all the revisions to be based on contemporaneous news and a common consensus benchmark. To enhance the reliability of our measure, we require that an analyst make at least eight forecast revisions in a calendar year, including a minimum of two upward and two downward revisions. We estimate the following analyst- and year-specific regression model to examine the sensitivity of an analyst's forecast revisions to bad versus good news:

$$REV_{j,t}^{i} = \alpha_0 + \alpha_1 BADNEWS_{j,t}^{N} + \beta_0 REV_{j,t}^{N} + \beta_1 BADNEWS_{j,t}^{N} * REV_{j,t}^{N} + \varepsilon_{j,t}^{i} \quad (1)$$

The variables are defined as follows:

$REV_{j,t}^{i}$ = Analyst i's earnings forecast revision, calculated as analyst i's forecast at time t for firm j minus the mean consensus forecast for firm j scaled by nearest preceding monthly stock price. The mean consensus forecast is based on forecasts made within 30 days prior to the first control analyst's forecast.

$REV_{j,t}^{N}$ = Average revision of analyst i's closest two neighbors (one preceding and the other succeeding analyst i), where each revision is calculated as the neighboring analyst's forecast for firm j minus the mean consensus forecast for firm j scaled by nearest preceding monthly stock price.

$BADNEWS_{j,t}^{N}$ = Bad news indicator which is equal to 1 for $REV_{j,t}^{N} < 0$ and equal to 0 when $REV_{j,t}^{N} >= 0$.

$\varepsilon_{j,t}^{i}$ = Error term.

We measure an analyst's (conditional) conservatism as $(\beta_0 + \beta_1) / \beta_0$, which scales the analyst's average forecast response to bad news by her average response to good news. In order

stock returns as news proxies.

to have a clean benchmark in measuring an analyst's incremental response to bad news, we require that the analyst react positively, on average, to good news in the calendar year, i.e., we require analyst-year observations with positive β_0 coefficients.[10] Finally, we quintile rank the conservatism measures computed during the estimation period (year t-1) to serve as our primary conservatism proxy, $CONSERV_{t-1}$, for our market response tests in year t.

3.2. General market demand for conservative analysts

Prior research on analysts' earnings forecasts documents the information content in forecast revisions by association with short-window market reactions (e.g., Givoly and Lakonishok, 1979; Lys and Sohn, 1990; Park and Stice, 2000; Clement and Tse, 2003; Frankel et al., 2006). From the earnings response literature (Collins and Kothari, 1989; Easton and Zmijewski, 1989; Kormendi and Lipe, 1987), we control for revision magnitude, loss firms, size, book-to-market, and systematic risk. From the analyst forecasting literature (Stickel, 1992; Clement, 1999; Jacob et al., 1999; Park and Stice, 2000; Clement and Tse, 2003; Chen et al., 2005), we control for *Institutional Investor* award status, broker size, experience, forecast frequency, past accuracy, and past optimism.

We estimate the following model to examine the incremental market reaction to the earnings forecasts of conservative analysts:

$$\begin{aligned} CAR_{j,t} = \; & \delta_0 + \delta_1 \, REV^{i}_{j,t} + \delta_2 \, CONSERV^{i}_{t-1} + \delta_3 \, REV^{i}_{j,t} * CONSERV^{i}_{t-1} \\ & + \Sigma_m \varphi_m \, Controls + \Sigma_m \gamma_m \, REV^{i}_{j,t} * Controls + \omega_{j,t} \end{aligned} \quad (2)$$

[10] In section 4.1, we discuss a number of sensitivity analyses regarding our conservatism measure; these do not materially affect the empirical results.

The model variables are defined as follows:

$CAR_{j,t}$	=	Market response, calculated as the five-day size-adjusted excess return for firm j centered on analyst i's forecast revision.
$REV^{i}_{j,t}$	=	Analyst i's earnings forecast revision, calculated as analyst i's forecast at time t for firm j minus the mean consensus forecast for firm j scaled by nearest preceding monthly stock price.
$CONSERV^{i}_{t-1}$	=	The quintile rank of analyst i's conservatism score in year t-1. The conservatism score measures the analyst's incremental forecast response to bad news scaled by her response to good news during year t-1, and is defined as $(\beta_0 + \beta_1) / \beta_0$ from the coefficient estimates in Eq. (1).
Response controls:		
$ABSREV^{i}_{j,t}$	=	Absolute value of analyst i's earnings forecast revision, where the revision is calculated as analyst i's forecast at time t for firm j minus the mean consensus forecast for firm j scaled by nearest preceding monthly stock price.
$LOSS_{j,t-1}$	=	Loss firm, an indicator variable equal to 1 for firms with negative actual earnings and 0 otherwise.
$FSIZE_{j,t-1}$	=	Firm size, calculated as the market capitalization at the end of year t-1 (Compustat annual data item #24 * Compustat annual data item #25).
$BM_{j,t-1}$	=	Book-to-market ratio, calculated as Compustat annual data item #60 divided by FSIZE at the end of year t-1.
$BETA_{j,t-1}$	=	Firm beta, obtained from a firm-specific regression of the firm's daily return on the value-weighted market index daily return using trading days in year t-1.
Analyst controls:		
$AWARD^{i}_{t-1}$	=	*Institutional Investor's All-American Research Team* indicator equal to 1 if analyst i was named to the list during year t-1 and 0 otherwise.
$BSIZE^{i}_{t-1}$	=	Brokerage house size, calculated as the natural logarithm of the number of analysts employed by analyst i's brokerage firm at the end of year t-1.
$GEXP^{i}_{t-1}$	=	Number of years since analyst i first appeared in I/B/E/S database.

$FREQ^{i}_{t-1}$	=	Number of earnings forecasts made by analyst i during year t.
$AVABSFE^{i}_{t-1}$	=	Average of ex-post absolute forecasts errors, i.e., actual earnings – forecast earnings, of analyst i during year t deflated by share prices ten days prior to the forecasts. (This measure is negatively related to accuracy.)
$AVFE^{i}_{t-1}$	=	Average of ex-post forecasts errors, i.e., actual earnings – forecast earnings, of analyst i during year t deflated by share prices ten days prior to the forecasts.
$\omega_{j,t}$	=	Error term.

Eq. (2) tests our first hypothesis, H1, which predicts that the market response will be greater for revisions of more conservative analysts. If the evidence is consistent with H1, then the term of interest, the interaction between the forecast revision and the conservatism measure, REV*CONSERV, will be positive. Since information content resides in both the sign and magnitude of the forecast revision itself, the forecast revision, REV, is a control variable and expected to be positive.

3.3. Institutional investor demand for conservative analysts

We investigate whether there is a stronger market response to conservative analysts' forecast revisions in the presence of greater institutional ownership. Relying on prior literature, we use the following firm characteristics to proxy for institutional investor ownership: the percentage of shares held by institutions, the number of institutions owning shares, and the number of shares held by institutions (Walther, 1997; Bartov et al., 2000; Bonner et al., 2003; Callen et al., 2005).[11] To examine the incremental extent to which institutional investors value the earnings forecasts of conservative analysts, we estimate the following model.

$$CAR_{j,t} = \lambda_0 + \lambda_1 REV^{i}_{j,t} + \lambda_2 CONSERV^{i}_{t-1} + \lambda_3 INST^{i}_{j} + \lambda_4 REV^{i}_{j,t} * CONSERV^{i}_{t-1} + \lambda_5 REV^{i}_{j,t} * INST_{j,t-1} + \lambda_6 CONSERV^{i}_{t-1} * INST_{j,t-1} \quad (3)$$

[11] For brevity, the latter two institutional proxies are used only in untabulated robustness tests.

$$+ \lambda_7 REV^{i}_{j,t} * CONSERV^{i}_{t-1} * INST_{j,t-1} + \Sigma_m \pi_m \text{Controls}$$

$$+ \Sigma_m \phi_m REV^{i}_{j,t} * \text{Controls} + \upsilon^{i}_{j,t}$$

where the only variable not previously defined is

$INST_{j,t-1}$ = The quintile level of institutional investment in firm j, defined as the percentage of institutional shares at the end of year t-1.

$\upsilon_{j,t}$ = Error term.

Eq. (3) tests our second hypothesis, H2, which predicts that the institutional market segment will react more strongly to the forecast revisions from conservative analysts. To support H2, we expect a positive coefficient estimate for the interaction REV*CONSERV*INST.

4. SAMPLE SELECTION AND RESULTS

4.1. Sample selection

In order to compute the analyst- and year-specific conservatism measure, Equation (1) is estimated using annual earnings forecasts in the I/B/E/S database between January 1989 and December 2005 for each analyst and year. We make a number of sample requirements to improve the reliability of our computation. First, stock price data from the CRSP database must exist to deflate forecast revisions, and stock prices must be greater than $1 to avoid small denominators and inferences based on "penny stocks."[12] Second, each forecast must be preceded and succeeded by forecasts issued by peer analysts each within one week from the forecast of interest. Third, a consensus forecast must be calculated using outstanding forecasts issued within 30 days prior to the forecasts of interest. This consensus forecast serves as the common revision benchmark for both the forecast of interest and the neighboring forecasts.[13] Fourth, a minimum

[12] An examination of excluded companies (those with analyst forecasts but no stock prices) show that they are primarily trusts, investment funds, small AMEX companies, and ADR's.

[13] To clarify the common benchmark, we calculate a mean consensus forecast from the forecast detail file based on analysts who have forecasted within 30 days of the first control analyst's forecast. Our results are similar if, alternatively, we use the nearest preceding consensus forecast from the I/B/E/S summary file.

level of analyst activity (eight forecasts per year, with at least two higher and lower than the consensus forecast) is required in order to allow the market a sufficient number of observations to gauge an analyst's level of conservatism.[14] Finally, we limit the influence of spurious conservatism estimation regressions with a forecast revision outlier treatment and by requiring positive β_0 estimates in Eq. (1).[15] The first column in Table 1 reports the resulting sample used in our conservatism estimation tests and includes 620,132 earnings forecasts made by 6,290 analysts, covering 7,541 companies.

The market response tests use earnings forecasts between January 1990 and December 2006. We use 1990 as the starting point since the I/B/E/S database may have suffered a time lag in recording forecasts until the mid-1980's (Keane and Runkle, 1998). Parallel to the sample used to construct our conservatism measure, we require available CRSP and Compustat data, a consensus forecast to serve as the revision benchmark and an analyst- and year-specific conservatism score from our estimation model. As shown in second column of Table 1, the resulting sample used in our market response tests includes 832,363 earnings forecasts made by 5,714 analysts, covering 9,702 companies.[16]

[Insert Table 1 about here]

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[14] The results are insensitive to changing the level of minimum activity to 6 or 10 forecasts per year (and 1 or 3 upward and downward minimum revisions).

[15] Our outlier treatment is based on forecast revision observations with absolute studentized residuals greater than three. However, omitting this outlier treatment, or using one of the following alternative treatments: omitting observations with Cook's D greater than 1 or regression estimations in the lowest R^2 decile, yields similar results. Further, we also retain very similar results if we relax the requirement of positive β_0 observations.

[16] There are two reasons why the market test sample is larger than the conservatism measure estimation sample. First, we estimate conservative reputation in one year and test the market response in the subsequent year, resulting in a one-year lag between the samples. This lag produces more observations in the market response sample since there are more forecasts made in 2006 (the year unique to the market response tests) than there were in 1989 (the year unique to the conservatism estimation sample). Second, we do not impose peer benchmark forecasts or analyst activity-level constraints (necessary to estimate the conservatism measure) on the market response sample; therefore, all analyst-years in t associated with a valid analyst conservatism score estimated in year t-1 qualify for inclusion in the market response tests.

To gain insight as to the impact of sample attrition due to our conservatism measurement requirements, we compare the analysts and covered firms that would have been included—except for these requirements—with the analysts and firms that actually enter our final sample. Untabulated analyses reveal that compared to the potential sample (13,276 unique analysts), the final sample (6,290 unique analysts) is characterized by larger broker size, higher percentage of award analysts, more firm coverage, more experienced analysts, more frequent forecasting, longer forecast age, greater accuracy and pessimism. Collectively, the differences between samples appear to be driven either directly or indirectly by analyst activity level, proficiency, and broker-employer size. While we do not see any apparent direct biases since we estimate the cross-sectional variation in conservatism *within* the final sample, we acknowledge that the differences between samples may affect the generality of our inferences. In particular, our results may not generalize well to analysts who are less active, less proficient, or to those employed at smaller brokers.

With respect to the effect of attrition on firm-year characteristics, untabulated analyses reveal that compared with the potential sample (10,938 unique firms), the final sample (7,541 unique firms) is characterized by larger market capitalization, smaller book-to-market, higher beta, fewer loss firms, greater profitability, larger analyst following, and greater analyst activity. Again, we recommend caution in generalizing our findings, particularly to smaller and less profitable companies.

Table 2, Panel A reports descriptive statistics regarding the conservatism estimation of Equation (1). Median β_0, restricted to be non-negative, is 0.885 and median β_1 is -0.106, suggesting that the median analyst revises her forecast by 89% of the positive revisions of fellow analysts, and (0.885-0.106=) 78% of the negative revisions of fellow analysts.[17] Accordingly, the median conservatism score, defined as $[(\beta_0+\beta_1)/\beta_0]$, is 89%. As described in Section 3, we rank the conservatism score into quintiles, $CONSERV_{t-1}$, for use in all market response tests in year t.

Table 2, Panel B reports descriptive statistics regarding the conservatism estimation of Equation (1) across $CONSERV_{t-1}$ quintiles. Consistent with intuition, the median β_0's, indicating responses to good news, decrease from 1.027 in the least conservative quintile to 0.304 in the most conservative quintile.[18] The median β_1's, indicating incremental reactions to bad news, monotonically increase from -1.338 in the least conservative quintile to 0.784 in the most conservative quintile. These patterns indicate that analysts in the more conservative quintiles respond incrementally more strongly to bad news than do analysts in the less conservative quintiles. Considering the total reaction to bad news, $\beta_0+\beta_1$, results indicate a slightly positive response to bad news by the least conservative analyst quintile, as well as overall stronger responses to bad news by the most conservative analyst quintiles.

[Insert Table 2 about here]

[17] We discuss medians of the parameter estimates (and use the conservatism rank quintiles in all tests) to mitigate the influence of outliers due to small denominators.

[18] All mean and median coefficient estimates of betas are statistically significant ($p < 0.01$) based on t-tests and Wilcoxon signed rank tests, respectively.

4.2. Descriptive statistics

Table 3 compares the lowest and highest conservatism quintiles with respect to analyst-year forecasting characteristics during the sample period.[19] Relative to less conservative analysts, more conservative analysts tend to work at larger brokerage houses (66.1 versus 57.8 analyst employees),[20] are more likely to be winners of the *Institutional Investor* All-Star award (17.9% versus 13.4%), have greater general experience (6.74 versus 6.22 years), and forecast more frequently (61.6 versus 59.8 issued forecasts).[21] More conservative analysts are not significantly different on dimensions of number of firms covered (16.9 versus 16.6 firms), firm experience (3.54 versus 3.35 years), or forecast age (165.7 versus 166.9 days prior to the earnings release).

[Insert Table 3 about here]

Table 4 reports Pearson and Spearman correlations between selected variables used in the primary models in the subsequent short-window market reaction tests. The primary measure of analyst conservatism, $CONSERV_{t-1}$, is positively associated with alternative conservatism measures such as the measure that relies on four years of forecast revisions, $CONSERV^*_{t-1}$, (Pearson correlation coefficient $r = 0.577$, $p < 0.01$) and the conservatism measure at the brokerage-level, $BCONSERV_{t-1}$, ($r = 0.319$, $p < 0.01$). We find that $CONSERV_{t-1}$ is negatively related to average price-deflated absolute forecast errors, $AVABSFE_t$, ($r = -0.023$, $p < 0.01$), indicating that more conservative analysts provide more accurate forecasts. We also find that $CONSERV_{t-1}$ is positively related to the deflated signed forecast errors, $AVFE_t$ ($r = 0.019$, $p < 0.01$), indicating that analyst conservatism is positively related to forecast

[19] The signs and significances of the differences are similar when the two most conservative quintiles are combined and compared with a combination of the least two conservative quintiles.
[20] Consistent with prior literature, the natural log of brokerage size, not the nominal number of analyst employees, is used as the control variable in the multivariate models.
[21] We report means in the text and interpret differences based on the t-tests. Unless otherwise noted, results based on medians and Wilcoxon tests are consistent.

pessimism. Since these pair-wise associations raise the possibility that we may be capturing either an analyst's past accuracy or pessimistic forecast patterns in our conservatism measure, we control for these measures—and variations of these measures—in our multivariate analyses.

[Insert Table 4 about here]

4.3. The market demand for conservative analysts

Table 5 reports the results from estimating Eq. (2) to evaluate the short-window market reaction to forecast revisions conditioned on analyst conservatism. We evaluate the market reaction to forecast revisions in a base model, with earnings response controls, and with both earnings response controls and analyst controls. For brevity, we discuss results of the base model, although the results are inferentially similar across models.[22] Throughout the multivariate analyses, we report OLS coefficient estimates and t-statistics in parentheses based on robust standard errors adjusted for both heteroscedasticity and intra-analyst error correlation (Rogers, 1993).[23] The estimated coefficient on the scaled forecast revision term, REV, is positive and statistically significant ($\delta_1 = 0.8250$, $t = 22.81$), indicating that the short-window market reaction around the revision is associated with the signed magnitude of the revision. Consistent with our first hypothesis, the interaction term, REV*CONSERV, is positive and significant ($\delta_3 = 0.0486$, $t = 4.48$), indicating that market participants react more strongly to the forecast revisions of more conservative analysts.

[22] Continuous variables are winsorized at the 1st and 99th percentiles.

[23] Since our data may produce residuals that are correlated across analyst observations, and generic OLS standard errors may incorrectly estimate the actual variation in coefficient estimates in such cases, we follow the recommendations in Petersen (2009). Specifically, we use White standard errors that are robust to within cluster correlation—also referred to in the literature as clustered or Rogers standard errors (Williams, 2000; Rogers, 1993). In particular, the covariance between residuals within cluster is estimated and the standard error is adjusted accordingly. As we implement it, this approach adjusts the standard errors for both heteroscedasticity and dependence in the residuals at the analyst-level; we then verify that alternative within-cluster adjustments at the followed-company and brokerage firm levels yield similar results.

To provide an interpretation of the economic significance of our result, we evaluate the differential effect of our conservatism measure by comparing results from the most and least conservative quintiles. For the least conservative quintile, the aggregate coefficient of market reaction is (0.8250+1*0.0486=) 0.8736. For the most conservative quintile, the aggregate coefficient of market reaction is (0.8250+5*0.0486=) 1.068, or 22% larger than that of the least conservative quintile.

[Insert Table 5 about here]

4.3.1. Non-overlapping sample

We limit the observations in the market response tests to a non-overlapping sample to mitigate the possibility of dependence influencing our results. Specifically, we use only observations with no other forecast revisions during the short-term window [-2, +2]. Although inferences drawn from the non-overlapping sample may seem more reliable than those from the full sample due to increased independence within the accumulation windows, the non-overlapping sample eliminates observations surrounding important information events; that is, events that induce multiple analysts to revise. This feature limits analyst interpretations of material events such as earnings announcements and management press releases. (Due to this concern, we rely on the full sample in the primary analyses.) Untabulated results from estimating Eq. (2) on the non-overlapping sample reveal the estimated coefficient on the interaction term, REV*CONSERV, remains positive and significant, indicating that market participants react more strongly to the forecast revisions of conservative analysts in windows which omit concurrent forecasts.

4.3.2. Earnings forecasting time-horizon

Prior literature has shown the benefits to firms of meeting and beating earnings benchmarks (Bartov et al., 2002; Brown and Caylor, 2005) and argued that analysts walk down their forecasts to allow management to exceed expectations (Richardson et al., 2004). Presumably, analysts are motivated to curry favor with management (Ke and Yu, 2006). Since the literature normally defines beating expectations relative to the last forecast issued prior to the earnings announcement (Bartov et al., 2002), it is not obvious how many analysts need to lower their expectations to enable management to beat expectations. Further, if a peer analyst has already revised downward, opportunism on the part of a particular analyst may result in a non-forecast rather than an additional lowered forecast. Overall, it is not apparent how widespread or material the incentives for lowering forecasts are vis-à-vis those for raising expectations, thus we view the potential time-horizon effect on our results as an empirical question. To address this issue, we re-perform our analyses after excluding near-term earnings forecasts, in which incentives for reducing expectations may be high. In particular, we exclude all analyst forecasts within 90 days of the earnings announcement and re-estimate both our conservatism measure and the subsequent market responses. Untabulated results indicate that the interaction term, REV*CONSERV, remains positive and significant.

4.3.3. Additional robustness efforts

We make a number of efforts to enhance the robustness of our results. Due to a lack of theory to guide the measurement of analyst conservatism, we use quintile ranks as a middle ground; however, we repeat our analyses using continuous and dichotomous measures. In addition to controlling for dependence econometrically through adjustments to standard errors, we limit our sample to a single firm-year revision (the most recent or mid-year) per analyst. One concern with relating analysts' forecast revisions to changes in share price is that other concurrent disclosures or economic events may affect price and may even cause analysts to revise (Chen et al., 2005). In an effort to address the concern of returns within our accumulation window "leading" the forecast revisions of our conservative analysts, we evaluate more narrow market responses with three-day [-1, +1] and single-day excess return [0, 0] windows. Summarizing, the results of these alternative analyses are inferentially similar to the original analysis and consistent with Hypothesis 1, which predicts stronger market reaction to the forecast revisions of conservative analysts.

4.4. The market demand for conservative analysts by institutional investors

Table 6 reports results from estimating Eq. (3) to evaluate the market reaction to forecast revisions conditioned on analyst conservatism and institutional ownership. The estimated coefficient on REV is positive and statistically significant ($\lambda_1 = 0.4378$, $t = 7.40$), supporting that the short-window market reaction around the revision is associated with the signed magnitude of the revision after controlling for other variables. The estimated coefficient on REV*INST, is positive and statistically significant ($\lambda_5 = 0.1865$, $t = 7.47$), indicating that institutional ownership increases the short-window market reaction to analyst forecast revisions. With respect to the term of main interest, REV*CONSERV*INST, we find a positive and significant interaction coefficient, ($\lambda_7 = 0.0189$, $t = 2.49$), indicating a relatively stronger market response to more conservative analysts' forecast revisions in the presence of higher institutional ownership. These findings suggest that the stronger market reaction to conservative analysts is driven to a greater degree by institutional investor presence.[24]

[Insert Table 6 about here]

4.5. Information-economic explanations underlying market results

Our results indicate a stronger market response to more conservative analysts' earnings research. In order to provide an economic interpretation of these results, we examine our conjecture that conservative analysts provide efficient forecasts, i.e., more accurate and more persistent earnings forecasts, where persistence implies a stronger association of forecast revisions with longer-term earnings revisions. To examine accuracy, we evaluate the association between the conservatism measure and *ex post* forecast accuracy while controlling for past accuracy and bias in a base model. We then extend this analysis using control variables from the forecasting accuracy literature such as broker size, award status, number of firms followed, general experience, and forecast frequency (Stickel, 1992; Mikhail et al., 1997; Clement, 1999; Jacob et al., 1999; Brown, 2001).

$$\begin{aligned} AVABSFE_{i,t} = \; & \kappa_0 + \kappa_1 CONSERV_{i,t-1} + \kappa_2 AVABSFE_{i,t-1} + \kappa_3 AVFE_{i,t-1} + \kappa_4 BSIZE_{i,t-1} \\ & + \kappa_5 AWARD_{i,t-1} + \kappa_6 NFIRMS_{i,t-1} + \kappa_7 GEXP_{i,t-1} + \kappa_8 FREQ_{i,t-1} \\ & + \kappa_9 FCAGE_{i,t} + \psi_{i,t} \end{aligned} \tag{4}$$

As reported in Table 7, Panel A, the conservatism variable $CONSERV_{t-1}$ is negatively and significantly associated with future absolute forecast errors (κ_1= -0.0116, t = -2.12), after controlling for past accuracy and past bias; the second model with the additional

[24] In untabulated results, we evaluate alternative institutional proxies, i.e., the number of institutional investors and the number of institutional shares, and nominal values of institutional ownership instead of the quintiles, with inferentially similar results.

control variables yields similar findings (κ_1= -0.0109, t = -2.00). These results suggest that analyst conservatism adds incrementally to past accuracy and past optimism to explain earnings forecast accuracy in the future year.

Examining revision persistence, we associate conservative analysts forecast revisions to long-term earnings realizations. In particular, we calculate two-year-ahead revisions, LT_EPSREV, as the actual two-year-ahead earnings minus the two-year-ahead consensus earnings expectation as of the analyst forecast revision date, scaled by the price at the end of the month preceding the forecast. We regress LT_EPSREV on analysts' forecast revisions, the conservatism measure, and controls consistent with the revision-based market response models.

$$LT\_EPSREV_{j,t} = \varphi_0 + \varphi_1 REV^{i}_{j,t} + \varphi_2 CONSERV^{i}_{t-1} + \varphi_3 REV^{i}_{j,t} * CONSERV^{i}_{t-1} + \Sigma_m \varphi_m Controls + \Sigma_m \gamma_m REV^{i}_{j,t} * Controls + \omega_{j,t} \quad (5)$$

In Table 7, Panel B, the base model reveals a positive and significant estimate on REV (φ_1 = 1.3516, t = 12.30) indicating that the annual forecast revision is positively related to longer term earnings revisions after controlling for other variables. Of primary interest, the interaction term, REV*CONSERV, is positive and significant (φ_3 = 0.1664, t = 5.14). The coefficient estimate (φ_3 = 0.1217, t = 3.73) remains significant in the second model, which includes the same firm and analyst control variables as the market tests. The positive and significant coefficient estimates on the interaction terms suggest that more conservative analysts forecasts are more strongly associated with longer-term earnings realizations.[25]

[25] In untabulated tests, we examine Equation (2) while including the *ex post* accuracy and persistence of analysts' forecasts. The inclusion of these forecast dimensions reduces the magnitude of the coefficient estimates of interest, the interaction between conservatism and revision, by 10%, indicating that the

4.6. Supplemental analyses

4.6.1. Conservatism across estimation windows

Since our findings suggest that investors react more strongly to forecasts of analysts with a high conservatism reputation in a prior period, we reason that Bayesian investors will react more strongly to analysts who develop more precise reputations based on longer estimation periods (Chen et al., 2005). We test this prediction by evaluating market responses to analyst conservatism levels separately estimated using periods of one, two, three, and four years. The untabulated results show that the estimated coefficients on the interaction term, REV*CONSERV, are positive, significant, and increasing monotonically in the estimation period length (0.0383, 0.0632, 0.0722, 0.0734 for periods of one, two, three, and four years, respectively). In this analysis, we constrain our sample only to those analysts forecasting for four contiguous years in order to isolate the variation across models to the composition of the sample; however, we find similar results after relaxing this constraint. In sum, our results suggest that the precision of an analyst's reputation for conservatism increases in her forecasting period and the market attends to the improvement in the conservatism signal.

4.6.2. Conservatism at broker-level

We next evaluate conservatism at the brokerage-level. We predict that brokers accrue—and investors react to—conservatism reputations based on the past forecasting behavior of their analysts. Similar reasoning appears in Barber et al. (2006), who analyze the distribution of stock ratings at the brokerage level in an effort to characterize the recommendations of their analysts; in Hong and Kubik (2003), who proxy for broker reputation by virtue of the firms' employed *Institutional Investor* award analysts; and in Cowen et

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aggregate market response to conservative analysts forecasts is, at least in part, due to their association with greater future accuracy and persistence.

al. (2006), who consider broker reputation for optimism based on their analysts' forecasts. To investigate, we modify Eq. (2) by replacing CONSERV with BCONSERV, defined as the annual quintile rank for a brokerage firm constructed from all analysts' forecasts working at that broker during year t-1. We impose an additional constraint, to preserve consistency, that analysts do not change firms during the year t-1. The untabulated results show estimated coefficient on the interaction term, REV*BCONSERV, is positive and significant (θ_3=0.1923, t=7.38), suggesting that our empirical analyses aggregates upward to the brokerage level.

5. CONCLUSION

There are many influences leading analysts to produce research in a less than conservative fashion. However, analysts may also have incentives to produce conservative research due to market incentives to build influence with clients. Anecdotally, among reasons for selecting analysts for the All-American team, fund managers frequently cite reasons such as "not endorsing everything," "having sober opinions and very little hyperbole," and "willing to present unpopular perspectives" (*Institutional Investor*, October 2006). Our empirical results are consistent with these anecdotes, in that we document a relatively stronger market response to more conservative analysts, as well as show this effect is greater in the presence of higher institutional ownership.

We document that analysts ranked higher on the conservatism measure have more experience and are more likely to be *Institutional Investor* All-Star award winners, perhaps indicating a reduced need to cultivate management relations. In addition, results show that conservative analysts forecast more efficiently. In other words, conservative analysts' forecasts are more accurate and they better predict long-term company earnings. In supplemental analyses, we document an increased market reaction to our measure of conservatism as we lengthen the estimation period and provide evidence that analysts' reputations for conservatism aggregate upward to the brokerage level.

Our findings make several contributions relevant to researchers and market participants. In an environment where company-specific uncertainty generally couples with analyst incentives to elevate expectations, one wonders if there is any reason to believe that analysts have incentives to produce research in a conservative manner. Through establishing a set of empirical links between conservatism in forecast patterns and the corresponding reaction of market participants—particularly those of institutional investors—

we support the argument that at least some analysts aim to forecast conservatively. In other words, our results suggest an equilibrium where many analysts produce aggressive research and are rewarded—as shown in prior empirical work and media reports, however, other analysts produce conservative—and on average more efficient—research and garner increased aggregate-market influence. Second, our empirical evidence offers an alternative perspective to those who view increased regulation as the primary solution to curtailing aggressive research practices. We show that investors—especially the institutional set—distinguish and reward the type of analysts that espoused regulations in the recent years try to bring about. Third, our results are relevant to investors who rely on analyst forecasts to form their earnings expectations, and brokerage houses, which use equity research outputs to establish credibility with clients (Brown, 1993).

We close with some limitations. In this study, we confine our analysis to sell-side analysts' earnings forecasts. Analysts perform other important tasks, such as writing reports, communicating with investors, making stock recommendations, and forecasting target prices (Asquith et al., 2005). Future research may consider if similar patterns of market observation of and response to analyst conservatism exist for these tasks. Due to conservatism estimation requirements, our results may not generalize well to analysts at smaller brokers, or less active and less proficient analysts. Further, we advise caution for generalization of our results towards smaller and less profitable firms, since these types of firms are under-represented in our final sample.

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Appendix
Variable Definitions

Variable	Definition
$ABSREV^{i}_{j,t}$	Absolute value of analyst i's earnings forecast revision, where the revision is analyst i's forecast at time t for firm j minus the mean consensus forecast for firm j scaled by nearest preceding monthly stock price. The mean consensus forecast is based on forecasts made within 30 days ahead of the first control analyst's forecast.
$AVABSFE^{i}_{t}$	Average of ex-post absolute forecasts errors, i.e., actual earnings – forecast earnings, of analyst i during year t deflated by share prices ten days prior to the forecasts. (This measure is negatively related to forecasting accuracy.)
$AVFE^{i}_{t}$	Average of ex-post forecasts errors, i.e., actual earnings – forecast earnings, of analyst i during year t deflated by share prices ten days prior to the forecasts. (This measure is negatively related to optimism.)
$AWARD^{i}_{t-1}$	*Institutional Investor* All-American award winner indicator equal to 1 if analyst i was named to the list during year t-1 and 0 otherwise.
$BADNEWS^{N}_{j,t}$	Bad news indicator which is equal to 1 for $REV^{N}_{j,t} < 0$ and 0 otherwise.
$BCONSERV^{i}_{t-1}$	Brokerage conservatism, calculated as the annual quintile rank for brokerages based on their analyst employees' CONSERV scores in year t-1.
$BETA_{j,t-1}$	Firm beta, obtained from a firm-specific regression of the firm's daily return on the value-weighted market index daily return using trading days in year t-1.
$BM_{j,t-1}$	Book-to-market ratio, calculated as Compustat annual data item #60 divided by FSIZE at the end of year t-1.

$BSIZE^{i}_{t-1}$	Brokerage house size, calculated as the natural logarithm of the number of analysts employed by analyst i's brokerage firm at the end of year t-1.
$CAR_{j,t}$	Market response, calculated as the five-day market-adjusted excess return for firm j centered on analyst i's forecast revision.
$CONSERV^{i}_{t-1}$	The quintile rank of analyst i's conservatism score in year t-1. The conservatism score measures the analyst's incremental forecast response to bad news scaled by her response to good news during year t-1, and is defined as $(\beta_0 + \beta_1) / \beta_0$ from the coefficient estimates in Eq. (1): $REV^{i}_{j,t} = \alpha_0 + \alpha_1 BADNEWS^{N}_{j,t} + \beta_0 REV^{N}_{j,t} + \beta_1 BADNEWS^{N}_{j,t} * REV^{N}_{j,t} + \varepsilon^{i}_{j,t}$
$FCAGE^{i}_{t}$	The average number of days between analyst i's earnings forecasts and the companies' earnings announcement dates during year t.
$FEXP^{i}_{t}$	The number of years in which analyst i made at least one forecast for the company covered during year t.
$FREQ^{i}_{t}$	Number of earnings forecasts made by analyst i during year t.
$FSIZE_{j,t-1}$	Firm size, calculated as the market capitalization at the end of year t-1 (Compustat annual data item #24 * Compustat annual data item #25).
$GEXP^{i}_{t}$	Number of years since analyst i first appeared in I/B/E/S database.
$INST_{j,t-1}$	The quintile level of institutional investment in firm j, defined as the percentage of institutional shares at the end of year t-1.

$LOSS_{j,t-1}$	Loss firm, an indicator variable equal to 1 for firms with negative actual earnings and 0 otherwise.
$NFIRMS^{i}_{t}$	Number of firms analyst i follows during year t.
$REV^{i}_{j,t}$	Analyst i's earnings forecast revision, calculated as analyst i's forecast at time t for firm j minus the mean consensus forecast for firm j scaled by nearest preceding monthly stock price. The mean consensus forecast is based on forecasts made within 30 days ahead of the first control analyst's forecast.
$REV^{N}_{j,t}$	Average revision of analyst i's closest two neighbors, where each revision is calculated as the neighboring analyst's forecast for firm j minus the mean consensus forecast for firm j scaled by nearest preceding monthly stock price.

Table 1
Sample Selection

	Earnings Forecasts (Unique Analysts) [Unique Companies]	
	Conservatism Estimation Tests	Market Response Tests
I/B/E/S annual earnings forecasts – U.S. universe, Jan 1989 - Dec 2005	2,226,108 (20,336) [25,467]	
I/B/E/S annual earnings forecasts – U.S. universe, Jan 1990 - Dec 2006		2,474,168 (23,364) [29,009]
After CRSP merge and a minimum $1 stock price restriction	1,677,218 (13,708) [11,920]	1,730,837 (13,895) [11,977]
After prior consensus forecast requirement	1,435,177 (13,276) [10,938]	1,471,127 (13,435) [11,051]
After the requirement of a preceding and a succeeding forecast from peer analysts, each within one week	868,868 (11,773) [8,275]	
After the restriction for minimum analyst activity and directional consistency.	620,132 (6,290) [7,541]	
After valid conservatism scores		832,363

		(5,714) [9,702]
Final sample	620,132 (6,290) [7,541]	832,363 (5,714) [9,702]

Table 1 provides the sample selection for the number of earnings forecasts, unique analysts, and unique companies. The first column provides sample selection for the conservatism estimation tests or Equation (1), and the second column for the market response tests or Equation (2). Section 4.1 provides details on sample restrictions and requirements.

Table 2
Descriptive Statistics on the Conservatism Score Estimation

Panel A: Parameter estimates

	p1	p25	p50	mean	p75	p99	% positive
α_0	-0.014	-0.001	-0.000	-0.001	0.000	0.005	38.2%
α_1	-0.009	-0.000	0.000	0.001	0.001	0.017	62.0%
β_0	0.027	0.513	0.885	1.152	1.227	5.460	100.0%
β_1	-6.002	-0.728	-0.106	-0.415	0.374	2.496	44.2%
$(\beta_0+\beta_1)/\beta_0$	-4.975	0.339	0.887	$9.3*10^7$	1.635	31.587	85.8%
R^2's	4.4%	46.8%	73.3%	66.4%	89.9%	99.7%	
# of forecast observations in analyst-year regressions	8	14	21	27.4	34	108	

Panel B: Parameter estimates across conservatism quintiles

	Conservatism Quintiles, $CONSERV_{t-1}$				
	1	2	3	4	5
	N=4,413	N=4,422	N=4,424	N=4,422	N=4,418
	mean p50	mean p50	mean p50	mean p50	mean p50
β_0	1.878 1.027	1.676 1.255	1.071 1.056	0.777 0.753	0.360 0.304
β_1	-2.432 -1.338	-1.047 -0.608	-0.150 -0.109	0.320 0.283	1.229 0.784
$(\beta_0+\beta_1)/\beta_0$	$-1.3*10^{10}$ -0.242	0.459 0.473	0.874 0.886	1.447 1.403	$1.8*10^{10}$ 3.633

Notes to Table 2
This table reports descriptive statistics on the conservatism estimation based on Eq. (1) performed by each analyst and year:

$$REV^{i}_{j,t} = \alpha_0 + \alpha_1 BADNEWS^{N}_{j,t} + \beta_0 REV^{N}_{j,t} + \beta_1 BADNEWS^{N}_{j,t} * REV^{N}_{j,t} + \varepsilon^{i}_{j,t}$$

Panel A reports descriptive statistics of the parameter estimates from the 22,099 analyst-year regressions of Eq. (1). The conservatism score, defined as $(\beta_0 + \beta_1) / \beta_0$, measures an analyst's response to bad news scaled by her response to good news during year t-1. $CONSERV_{t-1}$, the variable used in all analyses, is the quintile rank of the conservatism score (1 = least conservative quintile and 5 = most conservative quintile). Panel B reports descriptive statistics of β_0, β_1, and $(\beta_0 + \beta_1) / \beta_0$ across $CONSERV_{t-1}$ quintiles. N denotes the number of analyst-year groups in each quintile. The variable definitions appear in the Appendix.

Table 3
Descriptive Statistics on Earnings Forecast Patterns

	Conservatism Quintiles, $CONSERV_{t-1}$					
	1 N=3,885	2 N=4,019	3 N=4,102	4 N=4,121	5 N=3,992	5-1
			mean (median)			t-test (wilcoxon z)
$BSIZE_t$	57.8 (38.0)	66.5 (46.0)	69.0 (51.0)	71.0 (53.0)	66.1 (48.0)	6.18 *** (8.36) ***
$AWARD_t$	0.134 (0.00)	0.175 (0.00)	0.200 (0.00)	0.204 (0.00)	0.179 (0.00)	5.41 *** (6.09) ***
$NFIRMS_t$	16.6 (15.0)	16.6 (15.0)	17.1 (15.0)	17.3 (16.0)	16.9 (15.0)	1.44 (3.81) ***
$GEXP_t$	6.22 (5.00)	6.60 (6.00)	6.62 (6.00)	6.79 (6.00)	6.74 (6.00)	5.40 *** (6.26) ***
$FEXP_t$	3.35 (3.00)	3.38 (3.00)	3.62 (3.00)	3.41 (3.00)	3.54 (3.00)	1.09 (0.51)
$FREQ_t$	57.9 (50.0)	58.6 (53.0)	60.7 (55.0)	61.7 (56.0)	61.7 (55.0)	4.19 *** (5.46) ***
$FCAGE_t$	166.9 (163.0)	166.7 (163.5)	167.9 (164.8)	166.5 (163.7)	165.7 (163.2)	1.35 (0.47)

Notes to Table 3
This table compares one-year-ahead analyst-year characteristics among the conservatism quintiles (1 = least conservative quintile and 5 = most conservative quintile). N is the number of analyst-year groups for each quintile. The variable definitions appear in the Appendix.

* Two-tailed $p < 0.10$; ** two-tailed $p < 0.05$; *** two-tailed $p < 0.01$.

Table 4
Pearson and Spearman Correlations

	$CONSERV_{t-1}$	CAR_t	REV_t	$INST_t$	$CONSERV^*_{t-1}$	$BCONSERV_{t-1}$	$AVABSFE_t$	$AVFE_t$
$CONSERV_{t-1}$		-0.004***	-0.027***	-0.012***	0.577***	0.319***	-0.023***	0.019***
CAR_t	-0.002*		0.144***	0.009***	-0.006***	-0.007***	-0.007***	0.002**
REV_t	-0.033***	0.175***		0.078***	-0.033***	-0.025***	-0.034***	0.010***
$INST_t$	-0.012***	0.011***	0.044***		-0.020***	-0.006***	-0.169***	0.072***
$CONSERV^*_{t-1}$	0.575***	-0.002***	-0.039***	-0.021***		0.214***	-0.023***	0.010***
$BCONSERV_{t-1}$	0.303***	-0.003***	-0.027***	-0.013***	0.198***		-0.038***	0.031***
$AVABSFE_t$	-0.022***	-0.008***	-0.007***	-0.169***	-0.022***	-0.029***		-0.436***
$AVFE_t$	0.019***	0.001	0.000	0.071***	0.012***	0.027***	-0.424***	

Notes to Table 4

This table presents the Pearson (Spearman rank) correlations above (below) the diagonal of the selected variables. CONSERV* = $CONSERV^*_{t-1}$ based on four years of forecast observations (as opposed to a single year). The remaining variable definitions appear in the Appendix.

* Two-tailed $p < 0.10$; ** two-tailed $p < 0.05$; *** two-tailed $p < 0.01$.

Table 5
Market Reaction to Forecast Revisions Conditioned on Analyst Conservatism

$$CAR_{j,t} = \delta_0 + \delta_1 REV^{i}_{j,t} + \delta_2 CONSERV^{i}_{t-1} + \delta_3 REV^{i}_{j,t} * CONSERV^{i}_{t-1} + \Sigma_m \varphi_m Controls + \Sigma_m \gamma_m REV^{i}_{j,t} * Controls + \omega_{j,t} \quad (2)$$

	Expected Sign	Model 1 N =832,363	Model 2 N =806,937	Model 3 N = 802,002
Intercept		-0.0014 ***	0.0061 ***	-0.0008
		(-5.41)	(8.16)	(-0.69)
REV	+	0.8250 ***	2.1826 ***	3.1543 ***
		(22.81)	(21.99)	(18.89)
CONSERV		0.0000	0.0001	0.0000
		(0.55)	(1.32)	(0.42)
REV*CONSERV	+	0.0486 ***	0.0471 ***	0.0403 ***
		(4.48)	(4.43)	(4.01)
Response controls:				
ABSREV			-0.6966 ***	-0.6780 ***
			(-32.99)	(-32.28)
REV*ABSREV	−		-36.9677 ***	-37.1796 ***
			(-39.36)	(-39.60)
LOSS			-0.0036 ***	-0.0037 ***
			(-7.31)	(-7.47)
REV*LOSS	−		-0.2149 ***	-0.2191 ***
			(-7.11)	(-7.10)
Log(FSIZE)			-0.0007 ***	-0.0007 ***
			(-9.51)	(-10.00)
REV*Log(FSIZE)	−		0.0056	-0.0188 **
			(0.57)	(-1.96)
BM			0.0059 ***	0.0058 ***
			(15.99)	(15.35)
REV*BM	−		-0.5515 ***	-0.4989 ***
			(-16.46)	(-14.90)
BETA			-0.0018 ***	-0.0017 ***
			(-7.67)	(-7.56)
REV*BETA	+		0.2636 ***	0.2405 ***
			(12.16)	(11.75)
Analyst controls:				
AWARD				-0.0005 *
				(-1.65)
REV*AWARD	+			-0.1355 ***
				(-3.02)
Log(BSIZE)				0.0005 ***

				(4.39)
REV*Log(BSIZE)	+			0.1571 ***
				(8.25)
GEXP				-0.0000
				(-0.06)
REV*GEXP	+			0.0312 ***
				(7.36)
Log(FREQ)				0.0013 ***
				(6.25)
REV*Log(FREQ)	−			-0.3421 ***
				(-10.83)
AVABSFE				0.0034
				(0.32)
REV* AVABSFE	−			-3.6640 ***
				(-4.81)
AVFE				0.0158
				(1.10)
REV*AVFE	-			-1.9704 *
				(-1.90)
R^2		2.10%	3.63%	3.83%
F-value		759 ***	303 ***	191 ***

Notes to Table 5

This table reports results from estimating Eq. (2) to evaluate the market reaction to analysts' forecast revisions conditioned on conservatism during the sample period January 1990 - December 2006. We report the coefficient estimates for the OLS regression and t-statistics in parentheses based on standard errors adjusted for both heteroscedasticity and intra-analyst error correlation. Log (.) indicates that natural logarithm of a variable. The variable definitions appear in the Appendix.

* Two-tailed $p < 0.10$; ** two-tailed $p < 0.05$; *** two-tailed $p < 0.01$.

TABLE 6
Market Reaction to Forecast Revisions Conditioned on Analyst Conservatism and Institutional Ownership

$$CAR_{j,t} = \lambda_0 + \lambda_1 REV^{i}_{j,t} + \lambda_2 CONSERV^{i}_{t-1} + \lambda_3 INST^{i}_{j} + \lambda_4 REV_{i,j,t} * CONSERV^{i}_{t-1} + \lambda_5 REV^{i}_{j,t} * INST^{i}_{j} + \lambda_6 CONSERV^{i}_{t-1} * INST^{i}_{j} + \lambda_7 REV^{i}_{j,t} * CONSERV^{i}_{t-1} * INST^{i}_{j} + \Sigma_m \pi_m Controls + \Sigma_m \phi_m REV^{i}_{j,t} * Controls + \upsilon_{j,t} \quad (3)$$

	Expected Sign	Model 1 N = 814,553	Model 2 N =796,528	Model 3 N = 791,654
REV	+	0.4378 ***	2.6995 ***	3.7262 ***
		(7.40)	(23.09)	(21.01)
CONSERV		-0.0001	-0.0001	-0.0002
		(-0.35)	(-0.45)	(-0.90)
INST		-0.0000	0.0004 **	0.0004 **
		(-0.02)	(2.05)	(1.97)
REV*CONSERV	+	0.0101	0.0113	0.0042
		(0.58)	(0.64)	(0.24)
REV*INST		0.1865 ***	0.1278 ***	0.1235 ***
		(7.47)	(4.48)	(4.45)
CONSERV*INST		0.0000	0.0001	0.0001
		(0.88)	(1.23)	(1.30)
REV*CONSERV*INST	+	0.0189 **	0.0158 **	0.0162 **
		(2.49)	(2.18)	(2.28)
			Parameter estimates for response and analyst controls in Models 2 and 3 suppressed for brevity.	
R^2		2.30%	3.66%	3.85%
F-value		336 ***	231 ***	165 ***

Notes to Table 6

This table reports results from estimating Eq. (3) to evaluate the market reaction to analysts' forecast revisions conditioned on conservatism and institutional ownership during the sample period January 1990 - December 2006. We report the coefficient estimates for the OLS regression and t-statistics in parentheses based on standard errors adjusted for both heteroscedasticity and intra-analyst error correlation. The variable definitions appear in the Appendix.

Table 7
Analyst Conservatism, Future Forecast Accuracy, and Long-term Earnings Realizations

Panel A: Analyst Conservatism and Future Forecast Accuracy

$$\begin{aligned} AVABSFE_{i,t} = \; & \kappa_0 + \kappa_1 CONSERV_{i,t-1} + \kappa_2 AVABSFE_{i,t-1} + \kappa_3 AVFE_{i,t-1} + \kappa_4 BSIZE_{i,t-1} \\ & + \kappa_5 AWARD_{i,t-1} + \kappa_6 NFIRMS_{i,t-1} + \kappa_7 GEXP_{i,t-1} + \kappa_8 FREQ_{i,t-1} \\ & + \kappa_9 FCAGE_{i,t} + \psi_{i,t} \end{aligned} \quad (4)$$

	Expected Sign	1 n = 20,090	2 n = 19,999
Intercept		1.5192 ***	1.3235 ***
		(34.22)	(16.39)
$CONSERV_{i,t-1}$	-	-0.0116 **	-0.0109 **
		(-2.12)	(-2.00)
$AVABSFE_{i,t-1}$	+	0.5501 ***	0.5390 ***
		(70.65)	(67.71)
$AVFE_{i,t-1}$	-	-0.0265 ***	-0.0269 ***
		(-3.72)	(-3.77)
$BSIZE_{i,t-1}$	-		-0.0276 ***
			(-3.09)
$AWARD_{i,t-1}$	-		-0.0846 ***
			(-3.81)
$NFIRMS_{i,t-1}$	-		-0.1000 ***
			(-3.62)
$GEXP_{i,t-1}$	-		-0.0081 ***
			(-4.23)
$FREQ_{i,t-1}$	+		0.1679 ***
			(7.03)
R^2		31.5%	31.9%
F-value		2,500 ***	984 ***

Notes to Table 7 Panel A
This panel reports results from estimating Eq. (4) to evaluate forecast accuracy during the sample period January 1990 - December 2006. We report the coefficient estimates for the OLS regression and t-statistics in parentheses based on standard errors adjusted for both heteroscedasticity and intra-analyst error correlation. The variable definitions appear in the Appendix.

* Two-tailed $p < 0.10$; ** two-tailed $p < 0.05$; *** two-tailed $p < 0.01$.

Panel B: Analyst Conservatism and Long-term Earnings Realizations

$$LT\_EPSREV_{j,t} = \varphi_0 + \varphi_1 REV^{i}_{j,t} + \varphi_2 CONSERV^{i}_{t-1} + \varphi_3 REV^{i}_{j,t} * CONSERV^{i}_{t-1} + \Sigma_m \varphi_m Controls + \Sigma_m \gamma_m REV^{i}_{j,t} * Controls + \omega_{j,t} \qquad (5)$$

	Expected Sign	Model 1 n =238,493	Model 2 n = 229,468
Intercept		-0.0176 ***	-0.0473 ***
		(-31.24)	(-16.86)
REV	+	1.3516 ***	2.0000 ***
		(12.30)	(4.99)
CONSERV		0.0003 **	0.0003
		(1.97)	(2.28)
REV*CONSERV	+	0.1664 ***	0.1217 ***
		(5.14)	(3.73)
FSIZE			0.0043 ***
			(33.44)
REV*FSIZE	+		-0.0358
			(-1.32)
BM			-0.0160 ***
			(-13.96)
REV*BM			-0.0936
			(-0.83)
BETA			-0.0058 ***
			(-14.53)
REV*BETA	-		-0.0500
			(-0.82)
AWARD			0.0000
			(0.06)
REV*AWARD	+		0.2718 **
			(2.25)
BSIZE			-0.0000
			(-0.16)
REV*BSIZE	+		0.1648 ***
			(3.16)
GEXP			-0.0000
			(-0.53)
REV*GEXP	+		0.0015
			(0.13)

FREQ			0.0041 ***
			(8.26)
REV*FREQ			-0.0006
			(-0.57)
AVABSFE			-0.0030 ***
			(-13.33)
REV* AVABSFE	-		-0.1508 ***
			(-3.23)
AVFE			0.0008 ***
			(3.41)
REV*AVFE	-		-0.1674 ***
			(-4.07)
R^2		6.57%	10.2%
F-value		496 ***	199 ***

Notes to Table 7 Panel B

This table reports results from estimating Eq. (5) to evaluate long-term earnings realizations during the sample period January 1990 - December 2006. We report the coefficient estimates for the OLS regression and t-statistics in parentheses based on standard errors adjusted for both heteroscedasticity and intra-analyst error correlation. The variable definitions appear in the Appendix.

* Two-tailed $p < 0.10$; ** two-tailed $p < 0.05$; *** two-tailed $p < 0.01$.