Conference calls and information asymmetry†

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Abstract

We hypothesize that conference calls are voluntary disclosures that lead to long-term reductions in information asymmetry among equity investors. Cross-sectional and time-series tests show that information asymmetry is negatively associated with conference call activity. Firms initiating a policy of regularly holding conference calls experience statistically and economically significant and sustained reductions in information asymmetry, in contrast to one-time callers, who experience no significant decline in asymmetry. Since prior work shows that the cost of equity capital is increasing in the level of information asymmetry, our results suggest that firms that more frequently hold conference calls have lower costs of capital.

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Key Words: Information asymmetry; Voluntary disclosures; Conference calls; microstructure; Probability of informed trade

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1. Introduction

We examine whether voluntary disclosures, as measured by conference calls, are negatively associated with the level of information asymmetry among equity investors. Information asymmetry is a key concept because it drives the demand for financial disclosures. Economic theory (Diamond, 1985; Verrecchia, 2001) suggests that disclosure reduces information asymmetry through two channels: disclosure directly lowers the amount of private information relative to publicly-available information, and it indirectly reduces private information search incentives. Therefore, we predict that the information asymmetry for firms holding more frequent conference calls is lower than for others.

Information asymmetry is important to firms because it is positively related to the cost of equity capital. This positive relation arises because uninformed investors demand a return premium to compensate for their risk of trading with privately-informed investors. The risk is not diversifiable since uninformed investors are always at a disadvantage relative to informed investors.¹

We focus on conference calls for two primary reasons. First, the prior literature (Bushee et al., 2003; Frankel et al., 1999) documents that conference calls generally convey material information to the market as evidenced by significant increases in return volatility during the actual call periods. The disclosure of information to the market is a necessary condition for conference calls to reduce the level of information asymmetry. Second, Brown et al. (2003) document substantial variation in conference call activity, both across firms and over time. This high degree of cross-sectional and time-series variation allows for relatively powerful tests of the association between disclosure frequency and information asymmetry.

¹ Also see O’Hara (2003).
We measure the level of information asymmetry using the Probability of Informed Trade (PIN), which is based on the EKO market microstructure model developed by Easley, Kiefer, and O’Hara (1997). While new to accounting, this methodology has already been successfully used in the finance literature. The PIN is a firm-specific estimate of the probability that a particular trade order originates from a privately-informed investor, and hence, directly captures the extent of information asymmetry among investors in the secondary market.

Our use of a direct measure of information asymmetry contrasts with indirect spread-based proxies of asymmetry used in prior studies, such as Welker (1995) and Leuz and Verrecchia (2000). Using a direct measure avoids the numerous econometric problems and interpretation difficulties that occur when using spread-based proxies (Callahan et al., 1997; O’Hara, 1995). An additional advantage of the PIN methodology is that it enables us to analyze the sources of the underlying relation between conference call activity and information asymmetry. As discussed in Section 2, greater disclosure can either reduce the relative amount of informed trading and/or decrease how frequently certain investors obtain private information. Such analyses are not possible using spread-based proxies of information asymmetry.

We find strong cross-sectional and time-series evidence supporting our hypothesis that a firm’s conference call activity is negatively related to its level of information asymmetry. Our cross-sectional tests indicate that there is a negative and highly significant association between the number of conference calls held in one quarter and the level of information asymmetry during the subsequent quarter. The relation is significant in the pooled regression and in each of 11 quarterly regressions. The magnitude of the pooled coefficient suggests that each conference call is associated with a roughly three percent decrease in the average level of information.

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2 For example, the PIN has been used to investigate how informed trading varies across different stock exchanges (Easley et al., 1996) and types of securities (Easley et al., 1998), and the inter-temporal pattern of informed trading (Brockman and Chung, 2002).
asymmetry. Combined with the results in Easley et al. (2002), this finding suggests that a firm holding one conference call each quarter reduces its annual cost of equity capital by at least 15 basis points.

Prior studies, including Botosan (1997), Botosan and Plumlee (2002), and Sengupta (1998), examine the relation between an aggregate measure of disclosure activity and the cost of capital. Economic theory suggests that disclosure is related to the cost of capital either through disclosure’s effect on information asymmetry, or its effect on estimation risk, or both. Our findings suggest that at least part of the association between disclosure and the cost of capital is due to the effect of voluntary disclosure on information asymmetry.

Using a time-series framework, we examine changes in information asymmetry in the periods surrounding a firm’s initial conference call. We find a significant reduction in asymmetry only for firms initiating a policy of holding periodic conference calls; one-time callers experience no significant decline in asymmetry. The magnitude of the coefficient suggests that firms on average experience a 1/20th decrease in the level of information asymmetry after adopting a policy of regularly holding conference calls. These results suggest that firms must demonstrate a commitment to a higher level of disclosure in order to benefit from the decreases in information asymmetry. Further analyses suggest that neither our time-series nor our cross-sectional findings are a result of the potentially endogenous relation between the level of information asymmetry and conference call activity.

The next section develops the expected relation between conference calls and information asymmetry, while Section 3 describes the EKO market microstructure model and the PIN

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3 Investors are compensated for the non-diversifiable risk involved with estimating an asset’s unknown risk and return distribution parameters (Coles et al., 1995).
estimation procedure. In Section 4, we discuss our cross-section analyses, and our time-series
tests are presented in Section 5. Section 6 concludes the paper.

2. Background and hypothesis development

Information asymmetry exists when investors are differentially informed about a firm’s value
and it allows investors with superior information to trade profitably at the expense of other
investors. To compensate for these expected losses, uninformed investors demand a return
premium that increases in the risk of trading with privately-informed investors (O’Hara, 2003).
Easley et al. (2002) provide evidence that the level of information asymmetry is positively
associated with firms’ cost of capital.

Economic theory suggests that the risk of trading against a privately-informed investor
depends on two factors: (1) how frequently such investors obtain private information and (2)
how intensely they trade on their private information relative to the level of uninformed trading.
Therefore (as explained more fully below), an increase in disclosure frequency reduces the level
of information asymmetry either by leading to a reduction in the frequency with which investors
obtain private information (“private information events”) or decreasing the relative amount of
informed trading, or both.

Conference calls lead to a reduction in how frequently investors become aware of private
information in two ways. First, disclosure directly preempts some future private information
events by releasing the information publicly before it can be discovered privately. Conference
calls allow firms to summarize and disseminate a large quantity of information that otherwise
would likely have been privately discovered and traded on at some later date(s). Second, more
frequent disclosures allow investors to rationally anticipate that the firm is following an ex ante
policy of frequently disseminating information and hence, it is more likely to promptly release
material information in the future. Such a disclosure policy reduces the expected benefits of private information search activities. This indirect effect corresponds to that discussed in King et al. (1990), who posit that managers reduce the incentives for private information acquisition when they issue an earnings forecast.

Conference calls also reduce information asymmetry by reducing the relative amount of informed trading in the equity markets. Merton (1987) and Fishman and Hagerty (1989) describe models in which more informative disclosures reduce the costs associated with processing and assimilating public information. As a result, greater disclosure induces more investing by uninformed liquidity traders. Diamond and Verrecchia (1991) find that under certain conditions, the amount of uninformed trading by large investors increases as the firm discloses more information. These results suggest that uninformed investors are more likely to trade in firms with higher disclosure levels. We expect, therefore, that as a firm holds conference calls more frequently, there will be more uninformed trading. Ceteris paribus, more uninformed trading decreases the risk of trading against a privately-informed investor.

However, this ceteris paribus condition is unlikely to hold since previous research indicates that increases in uninformed trading will be associated with more informed trading. Kyle (1985) demonstrates that if investors are risk neutral and are not capital constrained, then the amount of informed trading varies proportionately with the expected amount of uninformed trading. Thus, the relative amount of informed trading is unchanged when the expected amount of uninformed trading changes. In practice, however, informed traders are likely to be risk averse and/or capital

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4 Consistent with these models, Leuz and Verrecchia (2000) find a significant increase in trading volume for German firms committing to higher disclosure levels. Since bid-ask spreads also decrease, their evidence suggests that uninformed investors generate much of the increased trading.

5 Consistent with this argument, accounting regulators and practitioners argue that more and better disclosure by firms leads to less information asymmetry among investors and “levels” the playing field so that the capital markets are more attractive to uninformed investors (AICPA, 1993; FASB, 2001; Levitt, 1998).
constrained, resulting in changes in the intensity of informed trading that are less than fully proportional to the changes in uninformed trading. Consequently, the net effect of more frequent conference calls is relatively less informed trading.

Based on the above discussion, we expect that by holding conference calls more frequently, a firm reduces the information search incentives of (potentially) privately-informed traders and alters the trading behavior of both uninformed and informed investors. Altering these search incentives leads to a lasting impact on the level of information asymmetry that extends well beyond any temporary effects that occur during the call period (Bushee et al., 2003; Frankel et al., 1999). Thus, we test the following hypothesis, stated in alternative form:

H1: Conference call activity is negatively associated with the level of information asymmetry.

Our ability to reject the null hypothesis of no association in our cross-sectional tests (Section 4) is made more difficult by the potentially endogenous relation between disclosure and information asymmetry. In addition to disclosure affecting information asymmetry, a manager’s decision to hold conference calls is likely to be influenced by the concurrent level of information asymmetry. For example, a manager could decide to hold a conference call because the current level of information asymmetry is high, in the belief that doing so would reduce the level of asymmetry. This potential endogeneity will cause the cross-sectional association between call frequency and asymmetry to be less negative because firms with higher levels of (pre-call) asymmetry are more likely to hold calls. However, the endogeneity should not induce significant results when in fact there are none.

In contrast, endogeneity may magnify the expected negative relation between asymmetry and conference call activity in our time-series tests (Section 5). This strengthening will occur when a firm experiences an unusual event, such as a merger, that leads to a temporary increase in
uncertainty and causes the firm to hold a conference call to discuss the event. After the event-related uncertainty has been resolved, information asymmetry will decline but the decrease cannot be unambiguously attributed to the call. In order to ascertain whether our time-series (and cross-sectional) tests are biased by this potential effect, we conduct additional analyses designed to mitigate the effects of endogeneity on our results. Specifically, we separately analyze firm-quarter observations where firms are following a policy of regularly holding conference calls. For these firms, the decision to hold a conference call is unlikely to result from temporary increases in the level of information asymmetry and hence, endogeneity is unlikely to be a factor.

3. The PIN measure of information asymmetry

3.1 Theoretical development

Information asymmetry manifests itself when investors trade on the basis of their private information. While it is not possible to identify which trades are based on private information, the presence of privately-informed traders in the market can be (imperfectly) inferred from large imbalances between the number of buy and sell orders. This observability provides the intuition behind the EKO model of information asymmetry (Easley, Kiefer, and O’Hara, 1997). The EKO model is a learning model in which the market maker draws inferences about the presence and type of private information based on the observed order flow. Over a trading day, prices converge to their full information levels as the private information is fully revealed through the trading activity of informed investors. Thus, one can estimate the probability of information-based trading (PIN) for a given stock over a particular period based on the daily order flow during the period.

The basic structure of the EKO model can be seen in the game tree in Figure 1. At the
beginning of each day, nature determines whether one or more “informed” investors acquire private information. Private information events (news) occur with probability \( \alpha \); no news days occur with probability \( 1 - \alpha \). When it occurs, private information contains either “good” news with probability \( 1 - \delta \) or “bad” news with probability \( \delta \), where good (bad) news indicates the asset is currently undervalued (overvalued). The asset is correctly valued on no-news days.

Trade orders are assumed to arrive sequentially to the market according to Poisson processes. Orders from uninformed buyers (sellers) arrive randomly at the daily rate \( \varepsilon_b \) (\( \varepsilon_s \)) every trading day. Orders from risk neutral and competitive informed traders arrive randomly at the daily informed arrival rate \( \mu \) only on good- and bad-news days.\(^6\) The model implicitly assumes that the informed arrival rate \( \mu \) is determined by some unmodelled strategic process that results in a fraction of trades equal to \( \mu / (\mu + \varepsilon_b + \varepsilon_s) \) arriving from informed traders on private information event days (O’Hara, 1995). All of the arrival processes are assumed to be independent and their parameters are common knowledge across all traders and the market maker. The market maker sets prices to buy or sell shares at each point in time based on her current information set, and executes orders as they randomly arrive.

The trading process described above will lead to one of three general patterns of trade orders. On a no-news day, the model predicts a roughly equal number of buyer- and seller-initiated trade orders.\(^7\) On a good- (bad-) news day, there will be a large imbalance in the order flow, with

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\(^6\) This assumption implies that there are no binding short sales constraints.

\(^7\) Although we do not restrict \( \varepsilon_b = \varepsilon_s \), our results indicate that these parameters are generally quite close in magnitude.
buyer-initiated (seller-initiated) trades predominating. Each of these patterns corresponds to a branch of the game tree shown in Figure 1.

Although unaware of the branch chosen by nature on any given day, the market maker knows the probability of each branch and the expected order process associated with each. She uses the observed number of buys and sells to update her beliefs throughout the trading day via Bayes Rule. Although her beliefs during the day depend on the actual sequence of trade orders on that day, the probability of informed trade at the beginning of the day has the following simple form:

\[
P(IN) = \frac{\alpha \mu}{\alpha \mu + \varepsilon_b + \varepsilon_s}
\]  

The PIN measure represents the expected fraction of trades that are information-based since the numerator is the expected number of orders from privately-informed investors and the denominator is the expected total number of orders each day. Thus, the ratio of the two is the ex ante probability that the first trade of the day is based on private information. Consistent with economic intuition, Eq. (1) shows that the level of information asymmetry is increasing when there are more frequent private information events (\(\alpha\)), more informed trading (\(\mu\)), and is decreasing in the willingness of uninformed investors to trade in the stock (\(\varepsilon_b\) and \(\varepsilon_s\)).

The likelihood function induced by the EKO model for a trading day, conditional on the parameter vector \(\theta = (\alpha, \delta, \varepsilon_b, \varepsilon_s, \mu)\), is determined by a mixture of the three possible components (no news, good news, and bad news) with weights reflecting the probabilities of their occurrence in the data. Letting \(B\) and \(S\) equal the daily number of buyer- and seller-initiated trades, respectively, the likelihood function for a particular trading day is:
As can be seen from Eq. (2), the daily numbers of buy and sell orders are sufficient statistics for the data. As one observes the order flow over an increasing number of days, one can estimate the parameter vector $\theta$ with increasing precision, assuming that the parameter vector $\theta$ is stationary.\(^8\) Essentially, the model uses the normal level of buying and selling to identify $\varepsilon_b$ and $\varepsilon_s$. Abnormal buy or sell order volume is interpreted as information-based and identifies $\mu$. The number of days on which there is abnormal buy or sell volume identifies $\alpha$ and $\delta$. The maximum likelihood estimates for each firm’s parameter vector $\theta$ over a particular period allows us to calculate firm- and period-specific PINs via Eq. (1). For example, consider a stock for which on 60% of the days there are 50 buys and 50 sells, on 20% of the days there are 80 buys and 50 sells; and on 20% of the days there are 50 buys and 80 sells. The EKO model parameters would be identified as $\varepsilon_b = \varepsilon_s = 50$, $\mu = 30$, $\alpha = .40$, and $\delta = .50$. The corresponding PIN equals 10.7%.

3.2 Empirical implementation

In order to estimate the PIN parameter vector $\theta$, we require the daily numbers of buy and sell orders over each estimation period for each firm. The number of observations in each estimation period depends on the specific analysis, with the minimum being about 30 trading days and the maximum is 62 trading days, representing one calendar quarter. Trade order data for NYSE, AMEX and Nasdaq traded stocks are obtained from the Trades and Quotes (TAQ) database.\(^9\) We classify each trade as buyer- or seller-initiated using the standard Lee-Ready algorithm (Lee

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\(^8\) Evidence in Easley et al. (2002) suggests that this assumption is reasonable.

\(^9\) In addition to including indicator variables in our analysis to control for exchange related differences, we also conduct sensitivity analyses by either excluding Nasdaq firms or limiting the sample to Nasdaq firms (because the exchange does not use market makers). None of our inferences are changed for these analyses.
and Ready, 1991). The algorithm classifies any trade that takes place above (below) the midpoint of the current quoted spread as a buy (sell) because trades originating from buyers (sellers) are most likely to be executed at or near the ask (bid). For trades taking place at the midpoint, a “tick test” based on the most recent transaction price is used to classify the trade. Following standard practice, we use a five second lag on reported quote times to adjust for differences in reporting times between quotes and trades on the NYSE. Following the recommendation in Ellis et al. (2002), we do not adjust quote times for Nasdaq and AMEX trades. Additionally, large trades are often broken down and matched against multiple investors. TAQ reporting conventions often classify such transactions as multiple trades. Following Hasbrouck (1988), we classify all trades occurring within five seconds of each other as a single trade.

We eliminate observations with extreme EKO parameter estimates using the following filters: (1) if $50\mu > \varepsilon$ or $50\varepsilon > \mu$, where $\varepsilon = \varepsilon_b + \varepsilon_s$; (2) if $\alpha < 0.02$ or $\alpha > 0.98$; (3) if $\delta < 0.02$ or $\delta > 0.98$; and (4) if $\min(\varepsilon, \mu) < 1$. These filters result in the elimination of 5,393 firm-quarter observations in the pooled, cross-sectional sample (representing 14% of the initial sample) and 1,981 observations in the time-series sample (19% of the initial sample).10 As a robustness check, we replicate each analysis reported in the tables below without applying the filters. In none of the analyses did the results change qualitatively. Specifically, all of the conference call coefficients retain the same sign and remain significant at the 5% level or better when the coefficients reported in the tables are significant.

4. Cross-sectional analyses

In this section, we examine the cross-sectional association between the frequency of conference

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10 The restrictions on $\delta$ are responsible for 63% and 69%, respectively, of the excluded observations.
calls and information asymmetry. The advantage of a cross-sectional analysis over a time-series approach is that the variation in conference call activity is potentially higher, thereby permitting more powerful tests. By the same token, other factors correlated with information asymmetry will also vary more across firms, so it is important to ensure that adequate control variables are employed.

4.1. Regression specification

The regression equation is shown in Eq. (3), with \( i \) and \( t \) denoting the firm and calendar quarter, respectively. We use calendar quarters because our data show a distinct seasonal pattern in conference call frequency within each quarter. Our sample begins in January 1999 and extends through December 2001; it includes all NYSE, AMEX and Nasdaq firms that had at least one conference call during the period.

\[
PIN_{it} = \beta_0 + \beta_1 \text{Calls}_{it-1} + \beta_2 \text{Size}_{it} + \beta_3 \text{Inside}_{it} + \beta_4 \text{InstOwn}_{it} + \beta_5 \text{Analysts}_{it} + \beta_6 \text{Consensus}_{it} + \epsilon_{it}
\]

In Eq. (3), \( \beta_0 \) is a vector of indicator variables for the stock exchange, calendar quarter, and 48 industries based on the Fama and French (1997) classification. \( \text{Calls} \) is the number of conference calls held during the prior quarter, \( \text{Size} \) is firm size, \( \text{Inside} \) is stock ownership by insiders, \( \text{InstOwn} \) is stock ownership by institutions, \( \text{Analysts} \) is the number of stock analysts covering the firm, and \( \text{Consensus} \) is the degree of consensus among analysts. Detailed variable descriptions are provided in Sections 4.2 and 4.3 below.

Measuring \( PIN \) in the calendar quarter after the quarter in which the conference calls take place (if any) ensures that the market had access to the information conveyed during the calls before the PIN estimation period begins. This procedure also allows us to avoid a potential confounding effect caused by a temporary increase in information asymmetry caused by the
release of information during the conference call.\textsuperscript{11} Public disclosures lead to an increase in information-based trading when investors have varying abilities to process information, and skilled investors transform the public but noisy signal of firm value into private information (Kandel and Pearson, 1995; Kim and Verrecchia, 1994). We expect that conference calls will lead to relatively large order imbalances on the day of the call and hence, be identified as private information events. Thus, we avoid any confounding relation between conference call frequency and PINs (through a higher value of $\alpha$) by measuring PINs after any calls take place.

\subsection*{4.2. Conference calls}

For each calendar quarter, we record the number of conference calls made by each firm. We include calls that were listed on any of the following sources that track conference calls: BestCalls.com, CCBN.com, and First Call. BestCalls.com aggregates information from a variety of sources about conference calls that are open to the public and then makes this information freely available on its web site. CCBN.com is the leading provider of earnings-related webcasts of conference calls and currently services more than 3,000 calls in a typical quarter. First Call is a leading provider of financial information that has long tracked conference call activity.

Firms use additional types of voluntary disclosures beyond conference calls, such as management forecasts and press releases, to disseminate information to the market. To the extent that conference calls and other types of voluntary disclosures are substitutes, this reduces our ability to detect the effect of conference calls on the level of information asymmetry. Alternatively, if conference calls are positively correlated with other voluntary disclosure activities, then conference calls should be interpreted as being indicative of firms’ overall propensity to disclose information voluntarily. Miller (2002) finds that disclosure frequency is

\textsuperscript{11} Frankel et al. (1999) and Bushee et al. (2003) find evidence consistent with conference calls leading to a temporary increase in information-based trading during the immediate call period.
positively correlated across different types of disclosures.

4.3. Control variables

To isolate the effect of conference calls on information asymmetry, we include variables in the regression to control for other aspects of the firm’s information environment that are likely to be associated with both PIN and the number of conference calls. The first variable we include is Size, which is defined as the natural log of the firm’s market value of equity measured at the beginning of quarter \( t \) using data from CRSP. Based on the results in Atiase (1985), Bamber (1987), and Diamond and Verrecchia (1991), we expect a negative association between Size and PIN.

We include three variables to control for the presence of sophisticated and potentially informed market participants. Inside (InstOwn) is the proportion of the firm’s shares held by insiders (institutions) at the beginning of quarter \( t \).\(^ {12} \) Ayers and Freeman (2000) and Jiambalvo et al. (2002) find that current returns reflect future earnings to a greater extent when the percentage of institutional and inside ownership is higher. These findings suggest that there is more current informed trading based on future earnings information for these firms, and hence, more information asymmetry. Accordingly, we expect the coefficients on Inside and InstOwn will be positive. Analysts is the number of analysts making forecasts during quarter \( t \) about the following quarter’s earnings and are obtained from First Call. Since Easley et al. (1998) find that PINs are negatively associated with the level of analyst following, we expect the coefficient on Analysts will be negative.

Consensus is used as a proxy for the degree of certainty about the expected level of earnings.

\(^{12} \) Insider ownership data come from Lancer Analytics and include direct and indirect holdings, along with derivative holdings such as options, held by all officers, directors, and other top executives. Institutional ownership data come from the CDA/Spectrum 13F institutional holdings database.
Zhang (2001) finds that the amount of earnings uncertainty is positively associated with both the level of information asymmetry and disclosure quality. Thus, we expect Consensus to be negatively associated with PIN. We define Consensus as $\ln(1 + 1/\text{STD})$, where STD is the standard deviation of EPS forecasts averaged over the quarter. We use the quarterly average since the standard deviation changes each time a forecast is added, dropped, or revised. When there are fewer than two forecasts, Consensus is set to the arbitrarily low value of zero.\footnote{We discuss the robustness of our results with respect to this assumption below. We choose not to use STD as an alternative to Consensus due to the difficulty of specifying a particular value for STD when there are fewer than two forecasts.}

4.4. Sample description

Our sample is made up of all firms that have at least one conference call between January 1, 1999 and December 31, 2001 according to any of the sources described above, together with required data from TAQ, CRSP, Spectrum and Lancer Analytics. While excluding firms that had no calls during our sample period will reduce the power of our tests, this restriction ensures that the types of firms included in our analysis are not substantially different from each other along unanticipated dimensions that we cannot control for in the regression. Our final sample consists of 34,035 firm-quarters representing 5,754 individual firms. Descriptive statistics are reported in Table 1, Panel A.

Over the sample period, the total number of calls is 31,652 and the mean (median) number of calls per firm per quarter is 0.93 (1). The mean (median) PIN is 18.24 (17.16), which indicates that there is a roughly 18\% probability that the opening trade on any given day is based on
private information. The mean and median $\alpha$ demonstrate that private information events occur on almost 40% of trading days. The average value of $\ln(\mu/\varepsilon)$ implies that $\mu/\varepsilon = 0.55$. This value indicates that informed trades are about half the level of uninformed trades and represent about 1/3 of total trades on information event days. For the average firm, 30% of its outstanding shares are owned by insiders and 43% are owned by institutions. Consistent with previous studies, the extent of analyst following is quite skewed, with the mean being over 2 while the median firm-quarter has no analyst coverage.

Summary statistics on the number of conference calls during each quarter are provided in Table 1, Panel B, and show that both the absolute number of calls and the average number of calls per firm increase dramatically over the sample period. The number of calls each quarter increases from less than 1,500 at the beginning of 1999 to over 3,300 in 2001. At the same time, the mean number of calls per firm per quarter increases from 0.56 at the beginning of our sample period to over 1.0 in each quarter in 2001. The upsurge in call activity does not appear to be directly related to Regulation FD, as calling activity increased steadily over our sample period and there is no pronounced jump in the fourth quarter of 2000 when Regulation FD became effective.

As shown in Table 2, there is a significant negative correlation between the number of conference calls in period $t-1$ and $PIN$ in period $t$ (-0.25), providing preliminary evidence supporting our hypothesis. The strongest correlations between $PIN$ and the control variables are with $Size$ (-0.76), $InstOwn$ (-0.44), and $Inside$ (0.35). All of the correlations with $PIN$ are of the
expected sign with the exception of \textit{InstOwn}. This unexpected result may be caused by the combination of the large positive correlation between \textit{Size} and \textit{InstOwn} (0.58) and the large negative correlation between \textit{Size} and \textit{PIN}.

\textbf{4.5. Conference calls and information asymmetry}

To investigate the cross-sectional association between the frequency of conference calls in one quarter and the level of information asymmetry in the subsequent quarter, Table 3 presents the results of estimating Eq. (3) for the entire pooled sample as well as for each of the eleven quarterly time periods. Overall, the explanatory power of the regressions is quite high, as the adjusted-$R^2$ is 59\% in the pooled regression and ranges from 53\% to 65\% in the quarterly regressions. The results for the pooled regression demonstrate that the number of conference calls in quarter $t-1$ is negatively and significantly associated with the level of information asymmetry in quarter $t$ ($t$-statistic = $-14.3$).\textsuperscript{14} The results for the quarterly regressions are generally consistent with the pooled regression, although the significance levels are predictably lower given the smaller sample sizes. All eleven of the quarterly conference call coefficients are significant at the 0.001 level or better, with the exception of the second quarter of 1999, where the coefficient is significant at the 5\% level. The magnitude of the \textit{Calls} coefficient in the pooled regression is $-0.59$ and ranges from $-0.30$ to $-0.91$ in the quarterly regressions. No temporal trend is evident in the quarterly coefficient magnitudes. Not surprisingly given the individual significance of the quarterly coefficients, the mean of the quarterly \textit{Calls} coefficients, $-0.62$, is also significantly negative as the Fama-MacBeth $t$-statistic is $-12.6$.

\begin{table}[h]
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\caption{Table 3 about here.}
\end{table}

\textsuperscript{14} All reported $t$-statistics are based on Huber-White standard errors.
The *Calls* coefficient in the pooled regression indicates that PINs in one quarter are 0.59 percentage points lower for each conference call held during the prior quarter. This represents a 3.2% decrease relative to the mean *PIN* of 18.24, which in our view represents a moderate and economically plausible effect on the level of information asymmetry. Combined with the findings in Easley et al. (2002) on the association between PINs and the cost of equity capital, our results suggest that holding a conference call each quarter is associated with a 15 basis point reduction in the annual cost of equity capital.

Brown et al. (2003) find that over a similar time period, more than 80% of firms that hold a conference call in one quarter hold at least one call during the subsequent quarter. This observation suggests that the negative association between the number of calls in period $t-1$ and the level of information asymmetry in period $t$ could be driven in part by the number of calls during period $t$, and thus, the results in Table 3 are potentially overstated. In order to investigate this possibility, we augment Eq. (3) with the number of conference calls held in period $t$ and repeat the analyses. Untabulated results show that all of the coefficients on $Call_{t-1}$ are negative, though slightly smaller in magnitude than their counterparts in Table 3, and all but one (second quarter of 1999) is significant at the 1% level. Additionally, all of the coefficients on $Call_t$ are negative and seven are significant at the 5% level. Thus, even after controlling for the number of contemporaneous calls, PINs are negatively associated with the number of conference calls during the prior quarter.

The coefficients for the control variables *Size*, *Inside*, and *Analysts* all have the expected signs and are significant at the 5% level or better in the pooled regression and in each of the
quarterly regressions. The results show that firm size is negatively associated with PIN.\textsuperscript{15} This result is consistent with our intuition that information asymmetry is less prevalent in large firms. Consistent with Ayers and Freeman (2000) and Jiambalvo et al. (2002), we find that firms with more ownership by insiders also have higher levels of information asymmetry. Additionally, all of the Analysts coefficients are negative and significant, indicating that greater analyst coverage is negatively related to the level of information asymmetry. These results are consistent with Easley et al. (1998) and Jiambalvo et al. (2002).

The results for the other two control variables, InstOwn and Consensus, are less consistent with our expectations. In the pooled regression, InstOwn is significantly positive (t = 2.05) as hypothesized, but the average of the quarterly coefficients is not significantly different from zero as the Fama-MacBeth t-statistic is 1.52. However, the quarterly regressions show that the pooled result masks the fact that InstOwn’s association with PIN has been inconsistent over time. The results show while the InstOwn coefficient is positive in 8 of the 11 quarterly regressions (and significant in 4 of them), it is (significantly) negative in 3 (2) of the quarterly regressions. All of the negative coefficients occur in 2001.

While the association between InstOwn and PIN varies over the sample period, the association between Consensus and PIN is consistently positive, which is contrary to our expectations. The association is highly significant in the pooled regression and in 7 of the 11 quarterly regressions. The mean of the quarterly coefficients is also significantly different from zero (t = 5.77). One possible explanation is that this result is driven by the high positive correlation between Analysts and Consensus (0.89). Untabulated results show that this is only a partial explanation as the pooled Consensus coefficient remains significantly positive (though at

\textsuperscript{15} Given the importance of Size in the regressions, we test whether the associations between PIN and conference calls are being driven by non-linearities in the relation between PIN and Size by including higher-order terms of Size in the regression. Augmenting the regression in this way does not alter any of our inferences.
a reduced level) when we exclude Analysts from the regression equation.\textsuperscript{16} We also examine whether our decision to assign the minimum Consensus score of zero to firms with less than two analyst forecasts is causing this result. However, restricting the sample to firms with two or more forecasts does not alter our inferences. Another interpretation of this result is that high levels of Consensus occur when analysts have been guided by management to expect a target level of accounting earnings; however, such earnings are more likely to be managed. In these cases, the difference between accounting and economic earnings is relatively large, which increases the incentives to search for private information.

Having found that the number of conference calls is negatively associated with the level of information asymmetry in the cross-section, we now investigate the sources of this association. Recall from Section 2 that voluntary disclosures are related to information asymmetry through two channels: a reduction in how frequently investors become aware of private information and a reduction in the relative trading intensity of informed investors. Therefore, we examine how conference call frequency is related to the frequency of private information events, $\alpha$, and the relative amount of uninformed trading, $\ln(\mu/\varepsilon)$.

The regression specification is similar to Eq. (3), except that the dependent variable is based on the underlying PIN parameters rather than the PIN itself. Panel B of Table 3 presents these results. The $\alpha$ regression results in row 1 show that the number of conference calls in period $t-1$ is positively and significantly related to probability of an information event in period $t$. This result is unexpected since PIN is an increasing function of $\alpha$.\textsuperscript{17} Thus, the negative associations between conference calls and information asymmetry documented in Panel A are not being

\textsuperscript{16} Unreported results show that the inferences from Table 3 are unchanged when Consensus is left out the regressions. The significance levels for the Analyst coefficients, while reduced, remain quite high.

\textsuperscript{17} Untabulated analyses indicate that the number of conference calls in period $t$ is also positively associated with $\alpha$ in periods $t+2$ ($t = 2.1$) and $t+3$ ($t = 0.32$) when the sample is restricted to “Regular” callers (see Section 4.6).
driven by a subsequent reduction in the frequency of information events.

The second regression reported in Panel B examines the net effect of disclosure frequency on the relative arrival rates of informed to uninformed traders, $\ln(\mu/\epsilon)$. Row 2 indicates that the Calls coefficient is significantly negative, indicating that more conference calls in one period are associated with relatively less trading by privately-informed investors during the next period. This finding is consistent with our arguments in Section 2 that more frequent conference calls attract more uninformed investors to the firm and that any increases in informed trading are not fully offsetting. The coefficient estimates from these two regressions combined with the mean values of $\alpha$ and $\ln(\mu/\epsilon)$ from Table 1 indicate that the reduction in $\mu/\epsilon$ is almost four times as large as the increase in $\alpha$. Whereas each call is associated with a $1/80^{th}$ increase in $\alpha$ ($0.49\%/39.51\%$), it is associated with a $1/20^{th}$ decrease in $\mu/\epsilon$ ($((e^{-0.66} - e^{-0.61})/e^{-0.61} = (e^{-0.05} - 1))$.

Thus, the results in Panel B indicate that more conference calls are associated with less information asymmetry via reductions in the relative amount of informed trading, and this effect more than overcomes the associated increase in the frequency of information events.

### 4.6. Endogeneity between conference calls and information asymmetry

The results reported above could be biased if the relationship between conference calls and information asymmetry is endogenous. That is, if firms hold more (fewer) conference calls in response to high (low) levels of asymmetry, then the association between calls and PINs documented above will be understated. To determine whether endogeneity is a significant factor, we examine a set of firms for which the decision to hold a conference call is less likely to be a function of the level of information asymmetry in the pre-call period.

We examine a set of firms who are “Regular” callers that routinely host conference calls in conjunction with their earnings announcements. Specifically, a firm is classified as a “Regular”
caller in a particular calendar quarter if it had at least two calls in conjunction with its quarterly earnings announcements (+/– 1 day) over the previous three calendar quarters.\footnote{We do not impose the restriction of three calls in three calendar quarters because the interval between quarterly earnings announcements is frequently different than three months.} Firms that regularly hold conference calls in conjunction with their earnings announcements are likely following an \textit{ex ante} disclosure policy. As such, the decision to host a call in the current period is unlikely to be affected by the concurrent level of information asymmetry. Following similar reasoning, we also examine firms that did not hold a single conference call during the previous three calendar quarters. These restrictions reduce our sample to 21,259 firm-quarters.

In order to assess the sensitivity of the results in Table 3, Panel A to the potentially endogenous relation between conference call activity and information asymmetry, we re-estimate the cross-sectional regression in Eq. (3) using our restricted sample. The untabulated results from this analysis show that the $\text{Calls}_{t-1}$ coefficient is significantly negative ($t = -13.4$). The magnitude of the coefficient ($-0.69$) is only slightly larger in magnitude than the coefficient ($-0.59$) based on the entire sample. In addition, untabulated tests show that the magnitudes of the $\text{Calls}_{t-1}$ coefficients in the $\alpha$ and $\ln(\mu/\epsilon)$ regressions are slightly larger (and equally significant) compared to their counterparts in Table 3, Panel B. These results suggest that any endogeneity between conference call activity and asymmetry only affects the results in Table 3 to a small extent, and as we discussed in Section 2, makes it more difficult to reject the null hypothesis.

5. \textbf{Time-series of PINs around changes in disclosure policy}

While we control for several important factors in the cross-sectional tests described in Section 4, it is possible that we omitted one or more variables that are correlated cross-sectionally with both $\text{PIN}$ and $\text{Calls}$ and that any such omission(s) could have spuriously produced the negative
associations we document in Table 3. In this section, we present time-series tests that use each firm as its own control in order to mitigate any omitted correlated variable problem. We examine whether firms experience less information asymmetry after they hold an initial conference call. We consider a conference call to be an Initiation call for firm $i$ if the firm held no other calls during the prior 13 months. We use a period of 13 months to avoid misclassifying firms that regularly hold one call a year, perhaps in conjunction with their annual earnings announcements. Focusing on conference call initiations allows for a relatively powerful test if the impact of calls on asymmetry is a long-lasting one, as we expect.

Our call initiation sample extends from January 1999 through June 2002. Given our classification criteria of no calls within the last 13 months, this implies that the earliest Initiation calls occur in February 2000. Our sample consists of 2,092 Initiations. As our focus is on the long-term effects of disclosures, we exclude the one-week period surrounding the Initiation call (−3 to +3 calendar days). Week –1 to week –13 inclusive comprises the Pre-Initiation period –1, weeks –14 to –26 (−27 to –39) comprise period –2 (−3), and likewise for the Post-Initiation periods. The following timeline summarizes the timing of variable measurement for this analysis.

<table>
<thead>
<tr>
<th>Period #</th>
<th>Pre-Initiation</th>
<th>Initiation</th>
<th>Post-Initiation</th>
</tr>
</thead>
<tbody>
<tr>
<td>–3</td>
<td>–2</td>
<td>–1</td>
<td>0</td>
</tr>
<tr>
<td>Event Week</td>
<td>–39 … -27</td>
<td>–26 … -14</td>
<td>–13 … -1</td>
</tr>
<tr>
<td></td>
<td>0</td>
<td>1 … 13</td>
<td>14 … 26</td>
</tr>
<tr>
<td></td>
<td>27 … 39</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Accordingly, the data are aligned in event time according to the date of the Initiation. Our final sample consists of 8,333 firm-periods.

While this time-series research design minimizes the possibility of omitting cross-sectionally correlated variables, structural changes in the average level of information asymmetry over time due to market-wide factors can confound the inferences. For example, a systematic decline in
PINs over time can lead to a significantly lower Post-Initiation PINs compared with Pre-Initiation levels even though the decline is not due to the Initiation event. To prevent this problem from occurring, we specifically control for systematic time trends in our analyses. The regression equation is as follows, with \( i \) indicating the firm and \( \tau \) the time period:

\[
PIN_{i,\tau} = \gamma_0 + \gamma_1 PostInitiation_{i,\tau} + \gamma_2 Time_{i,\tau} + e_{i,\tau}
\] (1)

\( PostInitiation \) equals one if the period is after the Initiation week and zero otherwise. In the first specification reported in Table 4, \( Time \) is a continuous variable corresponding to the calendar date of the Initiation event. In the second specification, \( Time \) is a vector of indicator variables corresponding to the calendar quarter in which the Initiation event occurs.

The results of estimating Eq. (4) are presented in Table 4, Panel A. Models 1 and 2 include all call initiators with available data and PINs are estimated for the three Pre-Initiation and three Post-Initiation periods. The results for Model 1 show that PINs are significantly lower at the 0.1% level (\( t = –5.33 \)) for firms after they hold an initial conference call. The \( PostInitiation \) coefficient, \(-1.26\), indicates that PIN falls by over one percentage point on average after the initial conference call. We find this magnitude represents a moderate and economically plausible effect on the level of information asymmetry. Although this coefficient is about twice as large in magnitude as the cross-sectional coefficient of \(-0.59\) on \( Calls \) reported in Table 3, recall that \( PostInitiation \) is an indicator variable, whereas \( Calls \) is a count variable that has 1st and 99th percentile values of 0 and 3, respectively (see Table 1). The results in Model 2, which uses the quarterly indicator variables, are almost identical.

------------------------------------------------------------------

Insert Table 4 about here.  

------------------------------------------------------------------
Lang (1998) observes that the choice to hold a conference call often constitutes a commitment to a change in disclosure policy as once a firm holds a single conference call, it typically continues to hold calls on a regular basis. Economic theory (Verrecchia, 2001) suggests that Initiation events representing \textit{ex ante} commitments to higher disclosure levels will have a larger impact on the level of information asymmetry compared to one-time disclosures that do not represent a change in disclosure policy. While we cannot determine firms’ \textit{ex ante} intentions directly from our data, we can use their call behavior after the initial conference call as indicators of their intentions. Models 3 – 5 in Table 4 present the results when we separately analyze firms based on their post-Initiation call behavior.

In Models 3 and 4, we eliminate all firms that do not have a single conference call in the 39 weeks following the Initiation period. In Model 3, we further restrict the Initiation sample to those firms that had at least one call during any of the three Post-Initiation quarters. This restriction reduces the sample size to 6,378 firm-quarter observations. In Model 4, we impose a tighter restriction and only include firms that had at least one call in the first Post-Initiation period. We impose this restriction in order to increase the likelihood that the sample contains firms that have committed to a policy of regularly holding calls. Assuming that market participants require the subsequent call to confirm the change in disclosure policy, we exclude observations from the first Post-Initiation period so that the analysis is not contaminated by information not concurrently available to the market. This restriction reduces the sample size to 3,540 firm-quarters. In Model 5, we examine the remaining firms that did not have any conference calls in the 39 weeks subsequent to the initial call. This sample contains 1,955 firm-quarters, which is roughly 25% of the total Initiation sample. We do not expect that holding a single call will lead to a significant long-term decrease in asymmetry because without signifying
increased disclosure in the future, one call is unlikely to alter the behavior or either uninformed or (potentially) informed investors.

The results show that for both Models 3 and 4, PINs are significantly lower in the post-Initiation periods ($t = -2.97$ and $-2.56$, respectively). The magnitudes of the two $PostInitiation$ coefficients are similar to each other ($-0.83$ and $-0.99$, respectively) and are about the same magnitude as in Models 1 and 2. The magnitudes of the coefficients suggest that after adopting a policy of regularly holding conference calls, firms generally experience a roughly five percent decrease in the average level of information asymmetry. As expected, the $PostInitiation$ coefficient ($-0.32$) in Model 5 for firms that do not hold additional conference calls is not significantly different from zero ($t = -0.67$). Together, these results suggest that only firms that have demonstrated a commitment to a higher level of disclosure experience significant decreases in the level of information asymmetry after the initial call. When the initial call appears to represent a one-time $ex post$ disclosure decision, perhaps in response to an unusual event such as an acquisition or change in CEO, we find no evidence that the single call leads to a significant, long-term effect on the level of information asymmetry.

Similar to the analysis presented in Table 3, Panel B, we examine how the PIN parameters change in the periods surrounding the initial conference call. Specifically, we examine the change in the frequency of private information events, $a$, and the relative amount of informed trading, $ln(\mu/\varepsilon)$ around the Initiation period. The regression specification follows Eq. (4), except that the dependent variables are based on the PIN parameters rather than the PIN itself. The results are presented in Table 4, Panel B and are based on the full Initiation sample of 8,333 firm-periods.

The results from these analyses are consistent with those from the cross-sectional
regressions in Panel B of Table 3. In particular, the $\alpha$ regression results indicate that the probability of an information event increases after the initial conference call as the $PostInitiation$ coefficient is significantly positive ($t = 4.63$), contrary to our expectations. The $\ln(\mu/\epsilon)$ regression results show that the $PostInitiation$ coefficient is significantly negative ($t = -6.8$), indicating that after the initial conference call, relatively more uninformed investors trade in the firm. More uninformed trading reduces the risk of trading with a privately informed investor. This negative effect on PIN dominates the positive effect on PIN via the increase in $\alpha$ and explains the negative association documented in Table 4, Panel A.

Some caution should be taken in interpreting the results in this section since the decision to hold an initial conference call might be endogenously related to the prior level of information asymmetry. For example, unusual events (mergers, top management changes, etc.) might cause both a higher level of asymmetry in the pre-Initiation period and lead the firm to hold the initial call. In such cases, the level of uncertainty would fall mechanically after the uncertainty about the event resolves itself in the post-Initiation periods, thereby inducing our results. However, the types of one-time events that are consistent with such a mechanical decline in asymmetry are also likely to lead firms to hold just a single conference call, rather than to implement a continuing policy of increased disclosure resulting in subsequent calls. Our finding that the post-Initiation decrease in asymmetry is only significant for firms that hold additional calls provides confidence that it is the firm’s change in disclosure policy that is leading to the decrease in PIN, rather than a mechanical decline due to the resolution of event-related uncertainty.

6. **Summary and conclusions**

In this paper, we investigate the long-run association between conference calls and the level of information asymmetry between investors. This research question is important because prior
literature demonstrates that the level of information asymmetry is positively associated with the cost of equity capital (Easley et al., 2002). We use the Probability of Informed Trade (PIN) as our measure of information asymmetry, which is based on the EKO market microstructure model (Easley et al., 1997).

We find strong evidence supporting our hypothesis, which is inconsistent with the suggestion in Leuz and Verrecchia (2000) that the disclosure environment in the U.S. is uniformly too rich to allow researchers to detect measurable economic benefits of higher disclosure levels. We find that there is a highly significant and negative association between conference call frequency and the subsequent level of information asymmetry in the cross-section after controlling for potentially confounding factors. We also examine whether information asymmetry is lower for firms after they hold an initial conference call. We find that firms whose initial calls are followed by subsequent calls experience a significant reduction in asymmetry after the initial call. At the same time, firms that do not hold additional calls in subsequent quarters do not experience a decrease in asymmetry. These results suggest that only ex ante commitments to increased disclosure result in sustained decreases in asymmetry, consistent with economic theory. Further analyses suggest that our results are robust to the potentially endogenous relation between information asymmetry and conference call activity.

In addition, we examine the underlying sources of the negative relation between conference calls and information asymmetry. We find that more frequent conference calls are associated with less trading by privately informed investors relative to the amount of trading by uninformed investors. These findings are consistent with the theoretical results in Fishman and Hagerty (1989) and Merton (1987) that more informative disclosures by firms increase the amount of uninformed trading.
Finally, our results should be interpreted with some caution because they are based on a partial analysis that does not take account of other avenues of voluntary disclosure. If conference calls are complements for other disclosures used by firms, the reduction in information asymmetry that we document may be driven by those other disclosures rather than the conference calls themselves. To clarify this matter requires an examination of the association between a more comprehensive measure of voluntary disclosure activity and information asymmetry, and whether such additional disclosure activities complement or substitute for conference calls.
References


Figure 1 – Game tree for EKO model

Notes:
\( \alpha \) is the probability of a private information event.
\( \delta \) is the probability that the private information event contains bad news.
\( \mu \) is the daily rate of informed trade arrival.
\( \epsilon_b (\epsilon_s) \) is the daily rate of uninformed buy (sell) trade arrival.
Nodes to left of the dotted line occur once per day while trade occurs continuously throughout the day to the right of the line.
Table 1 – Descriptive statistics for cross-sectional analysis

PIN is the Probability of Informed Trade in calendar quarter $t$ according to the EKO microstructure model. $\alpha$ is the probability of a private information event. $\ln(\mu/\varepsilon)$ is the natural logarithm of the ratio of the daily rate of informed trading on information event days to the daily rate of uninformed trades. $Calls$ is the number of conference calls held during calendar quarter $t-1$. $Size$ is the natural logarithm of market value at the beginning of quarter $t$. $Inside$ is the ratio of shares and derivatives owned by insiders to the number of shares outstanding. $InstOwn$ is the proportion of outstanding shares owned by institutional shareholders. $Analysts$ is the number of analysts making earnings forecasts during quarter $t$. $Consensus = \ln(1+1/STD)$ where $STD$ is the standard deviation of EPS forecasts, averaged over the quarter. $n = 34,035$.

Panel A – Descriptive statistics for cross-sectional regression variables

<table>
<thead>
<tr>
<th>Variable</th>
<th>Mean</th>
<th>Std</th>
<th>1%</th>
<th>Median</th>
<th>99%</th>
</tr>
</thead>
<tbody>
<tr>
<td>PIN (%)</td>
<td>18.24</td>
<td>7.68</td>
<td>5.41</td>
<td>17.16</td>
<td>38.80</td>
</tr>
<tr>
<td>$\alpha$</td>
<td>39.51</td>
<td>11.79</td>
<td>12.74</td>
<td>39.67</td>
<td>66.29</td>
</tr>
<tr>
<td>$\ln(\mu/\varepsilon)$</td>
<td>-0.61</td>
<td>0.69</td>
<td>-2.14</td>
<td>-0.60</td>
<td>0.90</td>
</tr>
<tr>
<td>Calls</td>
<td>0.86</td>
<td>0.70</td>
<td>0</td>
<td>1</td>
<td>3</td>
</tr>
<tr>
<td>Size</td>
<td>12.79</td>
<td>1.81</td>
<td>9.14</td>
<td>12.71</td>
<td>17.29</td>
</tr>
<tr>
<td>Inside</td>
<td>0.30</td>
<td>0.28</td>
<td>0.00</td>
<td>0.20</td>
<td>1.00</td>
</tr>
<tr>
<td>InstOwn</td>
<td>0.43</td>
<td>0.26</td>
<td>0.01</td>
<td>0.43</td>
<td>0.93</td>
</tr>
<tr>
<td>Analyst</td>
<td>2.08</td>
<td>3.44</td>
<td>0</td>
<td>0</td>
<td>16</td>
</tr>
<tr>
<td>Consensus</td>
<td>2.64</td>
<td>3.53</td>
<td>0</td>
<td>0</td>
<td>12.05</td>
</tr>
</tbody>
</table>

Panel B – PINs and conference calls by calendar quarter

<table>
<thead>
<tr>
<th>Quarter</th>
<th>Firms</th>
<th>PIN (%)</th>
<th>Calls</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Mean</td>
<td>Std</td>
<td>Total</td>
</tr>
<tr>
<td>Q1 1999</td>
<td>2,682</td>
<td>N/A</td>
<td>1,180</td>
</tr>
<tr>
<td>Q2 1999</td>
<td>2,726</td>
<td>18.16</td>
<td>1,519</td>
</tr>
<tr>
<td>Q3 1999</td>
<td>2,875</td>
<td>18.32</td>
<td>2,132</td>
</tr>
<tr>
<td>Q4 1999</td>
<td>3,047</td>
<td>18.57</td>
<td>2,448</td>
</tr>
<tr>
<td>Q1 2000</td>
<td>3,011</td>
<td>16.63</td>
<td>2,677</td>
</tr>
<tr>
<td>Q2 2000</td>
<td>3,083</td>
<td>17.85</td>
<td>2,753</td>
</tr>
<tr>
<td>Q3 2000</td>
<td>3,109</td>
<td>18.23</td>
<td>2,849</td>
</tr>
<tr>
<td>Q4 2000</td>
<td>3,214</td>
<td>18.43</td>
<td>3,192</td>
</tr>
<tr>
<td>Q1 2001</td>
<td>3,347</td>
<td>18.61</td>
<td>3,594</td>
</tr>
<tr>
<td>Q2 2001</td>
<td>3,203</td>
<td>19.13</td>
<td>3,585</td>
</tr>
<tr>
<td>Q3 2001</td>
<td>3,237</td>
<td>18.61</td>
<td>3,363</td>
</tr>
<tr>
<td>Q4 2001</td>
<td>3,183</td>
<td>18.56</td>
<td>3,540</td>
</tr>
</tbody>
</table>
Table 2 – Spearman rank correlations

Correlations are for the pooled sample of 34,035 firm-quarter observations. All correlations are significant at the 5% level or lower. See Table 1 for variable definitions.

<table>
<thead>
<tr>
<th></th>
<th>PIN</th>
<th>α</th>
<th>ln(μ/ε)</th>
<th>Calls</th>
<th>Size</th>
<th>Inside</th>
<th>InstOwn</th>
<th>Analyst</th>
<th>Consensus</th>
</tr>
</thead>
<tbody>
<tr>
<td>PIN</td>
<td>1</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>α</td>
<td>–0.20</td>
<td>1</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>ln(μ/ε)</td>
<td>0.86</td>
<td>0.63</td>
<td>1</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Calls</td>
<td>–0.25</td>
<td>0.22</td>
<td>–0.31</td>
<td>1</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Size</td>
<td>–0.76</td>
<td>0.45</td>
<td>–0.82</td>
<td>0.26</td>
<td>1</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Inside</td>
<td>0.35</td>
<td>–0.15</td>
<td>0.34</td>
<td>–0.06</td>
<td>–0.40</td>
<td>1</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>InstOwn</td>
<td>–0.44</td>
<td>0.28</td>
<td>–0.48</td>
<td>0.27</td>
<td>0.58</td>
<td>–0.35</td>
<td>1</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Analyst</td>
<td>–0.08</td>
<td>0.07</td>
<td>–0.10</td>
<td>0.06</td>
<td>0.12</td>
<td>–0.04</td>
<td>0.16</td>
<td>1</td>
<td></td>
</tr>
<tr>
<td>Consensus</td>
<td>–0.09</td>
<td>0.08</td>
<td>–0.11</td>
<td>0.04</td>
<td>0.16</td>
<td>–0.04</td>
<td>0.18</td>
<td>0.89</td>
<td>1</td>
</tr>
</tbody>
</table>
Table 3 – Cross-sectional associations with conference call frequency

This table examines the cross-sectional association between PINs (PIN parameters in Panel B) in period $t$ and the number of conference calls in period $t-1$. Periods are calendar quarters. Variable definitions are in Table 1. All regressions include stock exchange indicators and industry indicators based on Fama and French (1997). Pooled regression with data from all quarters includes calendar quarter indicators. Samples exclude outliers having studentized residuals with absolute value $> 3$. Huber-White $t$-statistics are presented below the coefficients. *, **, *** denote significance at the 5%, 1%, and 0.1% levels, respectively (one-tailed test).

Panel A – Using PINs as the dependent variable

<table>
<thead>
<tr>
<th></th>
<th>Calls</th>
<th>Size</th>
<th>Inside</th>
<th>InstOwn</th>
<th>Analysts</th>
<th>Consensus</th>
<th>Adj. R² (%)</th>
<th>n</th>
</tr>
</thead>
<tbody>
<tr>
<td>Predicted sign</td>
<td>–</td>
<td>–</td>
<td>+</td>
<td>+</td>
<td>–</td>
<td>–</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Pooled sample</td>
<td>–0.59***</td>
<td>–2.91***</td>
<td>1.48***</td>
<td>0.28*</td>
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<tr>
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<td>–0.30*</td>
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<td>–0.07*</td>
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<td>–2.14</td>
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<tr>
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<td>–0.54***</td>
<td>–2.78***</td>
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<td>–0.07*</td>
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<tr>
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<td>–0.79***</td>
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<tr>
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<td>–0.56***</td>
<td>–2.48***</td>
<td>1.47***</td>
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<td>–0.14***</td>
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<tr>
<td>Q2 2000</td>
<td>–0.72***</td>
<td>–3.03***</td>
<td>1.60***</td>
<td>0.84*</td>
<td>–0.07*</td>
<td>0.04</td>
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<td>1.27</td>
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<td>–0.58***</td>
<td>–3.00***</td>
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<td>5.32</td>
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<tr>
<td>Q4 2000</td>
<td>–0.91***</td>
<td>–3.00***</td>
<td>1.66***</td>
<td>1.13**</td>
<td>–0.23***</td>
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<td>2.52</td>
<td>–6.46</td>
<td>5.29</td>
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<tr>
<td>Q1 2001</td>
<td>–0.77***</td>
<td>–2.90***</td>
<td>1.41***</td>
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<td>–0.24***</td>
<td>0.15</td>
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<td>3.71</td>
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<td>–6.77</td>
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<tr>
<td>Q2 2001</td>
<td>–0.45***</td>
<td>–3.34***</td>
<td>1.11**</td>
<td>0.39</td>
<td>–0.23***</td>
<td>0.19</td>
<td>62.2</td>
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<tr>
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<td>–3.15***</td>
<td>1.06***</td>
<td>–1.13</td>
<td>–0.21***</td>
<td>0.18</td>
<td>65.3</td>
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<td>–4.92</td>
<td>–45.67</td>
<td>3.04</td>
<td>–2.79</td>
<td>–6.74</td>
<td>5.03</td>
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<tr>
<td>Q4 2001</td>
<td>–0.58***</td>
<td>–3.14***</td>
<td>0.97**</td>
<td>–1.18</td>
<td>–0.29***</td>
<td>0.25</td>
<td>63.7</td>
<td>3,139</td>
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<td>–3.94</td>
<td>–44.80</td>
<td>2.69</td>
<td>–2.58</td>
<td>–9.43</td>
<td>6.48</td>
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<td></td>
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<tr>
<td>Mean of quarterly coefficients</td>
<td>–0.62***</td>
<td>–2.93***</td>
<td>1.48***</td>
<td>0.47</td>
<td>–0.16***</td>
<td>0.12***</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Fama-MacBeth $t$-statistic</td>
<td>–12.57</td>
<td>–39.62</td>
<td>12.21</td>
<td>1.52</td>
<td>–6.86</td>
<td>5.77</td>
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</table>
Table 3 – Cross-sectional associations with conference call frequency – continued

Panel B – Using the PIN parameters as the dependent variables
Results based on a pooled sample with data from all calendar quarters.

<table>
<thead>
<tr>
<th>Predicted sign</th>
<th>Calls</th>
<th>Size</th>
<th>Inside</th>
<th>InstOwn</th>
<th>Analysts</th>
<th>Consensus</th>
<th>Adj. R² (%)</th>
<th>n</th>
</tr>
</thead>
<tbody>
<tr>
<td>α - Probability (in %) of a private information event</td>
<td>0.49</td>
<td>3.42</td>
<td>-0.64</td>
<td>1.88***</td>
<td>0.12</td>
<td>-0.09***</td>
<td>31.4</td>
<td>33,953</td>
</tr>
<tr>
<td>ln(μ/ε) – Relative amount of informed trading</td>
<td>-0.05***</td>
<td>-0.32***</td>
<td>0.14***</td>
<td>0.01</td>
<td>-0.02***</td>
<td>0.02</td>
<td>75.0</td>
<td>33,769</td>
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</tbody>
</table>
Table 4 – Time-series tests around conference call initiations
This table shows results of analyzing whether PINs (PIN parameters in Panel B) decrease subsequent to a firm’s initial conference call. An initiation event is determined as a conference call that was preceded by no call activity during the prior 13 months. PINs are estimated over 13 week periods, with three pre-initiation and three post-initiation periods. The initiation period of seven calendar days centered on the initiation event is excluded from the analysis. PostInitiation = 1 if the period is after the initiation event. Time is a continuous variable for the calendar date of the initial conference call. Data spans 14 calendar quarters from January, 1999 through June, 2002. Samples exclude outliers having studentized residuals with absolute value > 3. Huber-White t-statistics reported below coefficients. *, **, *** denote respectively significant coefficients at the 5%, 1%, and 0.1% levels (one-tailed test).

Panel A: Using PINs (in %) as the dependent variable

<table>
<thead>
<tr>
<th>Model</th>
<th>Sample Includes</th>
<th>Constant</th>
<th>Post-Initiation</th>
<th>Time</th>
<th>Adj. R² (%)</th>
<th>n</th>
</tr>
</thead>
<tbody>
<tr>
<td>Predicted sign‡</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>1.</td>
<td>All firm-periods with data</td>
<td>Coefficient $20.1$</td>
<td>$-1.26^{***}$</td>
<td>$0.29$</td>
<td>$0.63$</td>
<td>$8,333$</td>
</tr>
<tr>
<td></td>
<td></td>
<td>$t$-statistic $60.78$</td>
<td>$-5.33$</td>
<td>$6.91$</td>
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</tr>
<tr>
<td>2.</td>
<td>All firm-periods with data†</td>
<td>Coefficient $13.5$</td>
<td>$-1.24^{***}$</td>
<td>quarterly indicators</td>
<td>$1.75$</td>
<td>$8,333$</td>
</tr>
<tr>
<td></td>
<td></td>
<td>$t$-statistic $2.06$</td>
<td>$-5.11$</td>
<td></td>
<td></td>
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</tr>
<tr>
<td>3.</td>
<td>Firms with calls in post-initiation periods</td>
<td>Coefficient $20.1$</td>
<td>$-0.83^{**}$</td>
<td>$0.19$</td>
<td>$0.24$</td>
<td>$6,378$</td>
</tr>
<tr>
<td></td>
<td></td>
<td>$t$-statistic $51.80$</td>
<td>$-2.97$</td>
<td>$3.72$</td>
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</tr>
<tr>
<td>4.</td>
<td>Include only firms that have calls in period +1††</td>
<td>Coefficient $19.9$</td>
<td>$-0.99^{**}$</td>
<td>$0.15$</td>
<td>$0.20$</td>
<td>$3,540$</td>
</tr>
<tr>
<td></td>
<td></td>
<td>$t$-statistic $39.15$</td>
<td>$-2.56$</td>
<td>$2.21$</td>
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</tr>
<tr>
<td>5</td>
<td>Firms with zero calls in post-initiation periods</td>
<td>Coefficient $23.1$</td>
<td>$-0.32$</td>
<td>$0.13$</td>
<td>$0.14$</td>
<td>$1,955$</td>
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<td>$t$-statistic $34.21$</td>
<td>$-0.67$</td>
<td>$1.67$</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

‡ Predicted sign does not apply to Model 5, which has a predicted coefficient of zero.
† Time is replaced by calendar quarter indicator variables that correspond to the initial call date.
†† Post-initiation period excludes period +1 to ensure that call information is available to market.

Panel B: Using the PIN parameters as the dependent variable

<table>
<thead>
<tr>
<th>Model</th>
<th>Dependent variable</th>
<th>Constant</th>
<th>Post-Initiation</th>
<th>Time</th>
<th>Adj. R² (%)</th>
<th>n</th>
</tr>
</thead>
<tbody>
<tr>
<td>Predicted sign</td>
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<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>1.</td>
<td>$\alpha$ – Probability (in %) of a private information event</td>
<td>Coefficient $45.90$</td>
<td>$1.63$</td>
<td>$0.34$</td>
<td>$1.17$</td>
<td>$8,333$</td>
</tr>
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<td>$t$-statistic $92.93$</td>
<td>$4.63$</td>
<td>$5.41$</td>
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</tr>
<tr>
<td>2.</td>
<td>$\ln(\mu/\delta)$ – Relative amount of informed trading</td>
<td>Coefficient $-0.63$</td>
<td>$-0.12^{***}$</td>
<td>$0.01$</td>
<td>$0.55$</td>
<td>$8,333$</td>
</tr>
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