Insider Ownership and Financial Analysts' Information Environment: Evidence from Dual-Class Firms

Abstract

We examine the association of insider ownership with financial analysts' forecast accuracy and dispersion in a sample of U.S. dual class firms. Insider ownership exerts two effects: a positive incentive and a negative entrenchment effect. The lack of significant findings in prior research regarding the association between insider ownership and forecast accuracy may be attributable to the offsetting forces of these two effects. Using a comprehensive hand-collected sample of U.S. firms that maintain more than one class of common stock, we are able to disentangle incentive and entrenchment effects which are confounded in single-class firms. We find that disproportionate insider control is negatively associated with forecast accuracy and positively associated with forecast dispersion. Moreover, insider cash flow rights (insider voting rights) are positively (negatively) associated with forecast accuracy and negatively (positively) associated with forecast dispersion, consistent with incentive-alignment and entrenchment effects of ownership affecting financial analysts' forecasting environment in opposite directions.

I. Introduction

This study examines the association of insider ownership with financial analysts' forecast accuracy and forecast dispersion. Prior work on the effects of firms' ownership structure and forecast accuracy (Byard, Li, & Weintrop, 2006; Ljungqvist, Marston, Starks, Wei, & Yan, 2007; García-Meca & Sánchez-Ballesta, 2011) documents that institutional ownership significantly affects forecast accuracy, but does not establish a significant link between insider ownership and forecast accuracy. This is puzzling, because insider ownership has been recognized as an important aspect of firms' overall corporate governance structure (e.g., Harris & Raviv, 1988; La Porta, Lopez-de-Silanes, Shleifer, & Vishny, 2000; Singh & Davidson, 2003). Insider ownership can help align the interests of management and shareholders, but at the same time it provides management with a means of entrenchment (Morck, Shleifer, & Vishny, 1988; Duellman, Ahmed, & Abdel-Meguid, 2013). We posit that the non-significant findings in prior work may be due to these peculiar counteracting effects conveyed by insider ownership.

To disentangle the incentive and entrenchment effects of ownership, we therefore follow prior studies (Francis, Schipper, & Vincent, 2005; Masulis, Wang, & Xie, 2009; Gompers, Ishii, & Metrick, 2010, hereinafter: GIM) that focus on a set of companies with unique characteristics: firms with more than one class of common stock.[1] Unlike in firms with a single class of common stock, cash flow rights and voting rights owned by insiders tend not to be identical in dual-class firms. This unique feature of dual-class firms allows us to separately assess how incentive-alignment effects of insider ownership (from cash flow rights) and entrenchment effects of ownership (from voting rights) affect financial analysts' forecast accuracy and dispersion. Moreover, we are also able to examine how disproportionate insider control, measured as the excess of insider voting rights over their cash flow rights, commonly referred to as "wedge" (GIM, 2010), affects financial analysts' forecast properties.

We expect disproportionate insider control to be negatively associated with financial analysts' information environment for two interconnected reasons. First, managerial entrenchment may be associated with reduced credibility of the accounting information available to market participants (Francis et al., 2005). Less credible accounting information can generate uncertainty in analysts' information environment (Barniv & Cao, 2009) and therefore trigger less accurate analyst earnings forecasts. Second, concentrated ownership by insiders may stem from insiders' desire to operate in greater secrecy, which results in tighter control of the information flow from the corporation to external users (Fan & Wong, 2002).

[1] Traditionally dual-class firms have been a popular choice among firms with family-ownership roots (DeAngelo & DeAngelo, 1985; Field & Karpoff, 2002; Villalonga & Amit, 2009). A dual-class structure allows founders to benefit from going public and opening corporate ownership to outsiders, while maintaining control of the firm. Dual-class firms are also common among firms in which non-financial benefits of control are important, such as the ability to influence publication policies for firms operating in the media industry (Demsetz & Lehn, 1985; DeAngelo & DeAngelo, 1985). Recently, many high-profile technology firms have adopted a dual-class structure.

However, we note that a negative effect of disproportionate insider control on the quality of financial information is not unambiguous. For instance, Nguyen & Xu (2010) and Chen (2008) present evidence of a lower instance of earnings management in dual-class firms. Because entrenched insiders are less dependent on the good will of shareholders, they may have less reason to conceal or manage financial information. While this rationale suggests possibly higher quality information, for the same reason the quantity of information supplied may be lowered, because entrenched insiders may have lowered incentives to provide timely and/or comprehensive information to their outside shareholders. On the whole, we therefore believe that disproportionate insider control likely is negatively associated with financial analysts' information environment.

We begin our examination by developing a new hand-collected sample of dual-class firms from 2000 to 2012 largely following the approach of GIM (2010). That is, we first build a comprehensive list of potential dual-class firms identified from several sources, and then check all candidates to determine if the firm actually has a dual-class structure of shares. Through these procedures, we are confident to present a sample which approaches the population of public dual-class firms in the U.S. during this time period.[2] Our focus on dual-class firms potentially introduces a sample selection bias into our analysis, since the sample of firms we study is not randomly selected from the population of U.S. public firms. In addition, like most studies of companies' ownership structure or governance characteristics, endogeneity may affect the interpretability of our results. To address these concerns, we employ two-stage regressions that control for sample selection bias, following the Heckman (1979) methodology, and endogeneity, in addition to our single-stage regressions.

[2] In total, we identify 530 unique dual-class firms that contribute a total of 3,523 firm-years. Upon merging our dual-class sample with necessary data from Compustat, I/B/E/S, and Thomson Reuters the number of observations in our main analyses is 2,050 (356 unique firms).

Across all three approaches, we find a significant negative relationship of disproportionate insider control and analysts' forecast accuracy. Likewise, insider voting rights are negatively associated with forecast accuracy, while insider cash flow rights show a positive association. Consistent with our results for accuracy, we also find that disproportionate insider control is positively associated with the dispersion of earnings forecasts. Increased dispersion indicates greater uncertainty and lower consensus among analysts (Imhoff & Lobo, 1992; Barron & Stuerke, 1998; Barron, Kim, Lim, & Stevens, 1998) and a poorer information environment (Byard, Li, & Yu, 2011). The association of insider voting rights (cash flow rights) with forecast dispersion is correspondingly positive (negative). All of these results are robust across different forecast horizons, model specifications, and a number of additional robustness tests, for instance, controlling for firms' private benefits capacity and reverse causality. In addition, we find results consistent with our levels models for a changes (first differences) model which demonstrates that contemporaneous and lagged changes in disproportionate control are positively (negatively) associated with changes in analysts' forecast dispersion (accuracy).

Taken together, our findings add to prior work by establishing that insider ownership *does* have a bearing on financial analysts' forecast accuracy and forecast dispersion. However, the association is a product of counteracting elements of ownership. While incentive-alignment effects of ownership have a positive impact, entrenchment effects of ownership exercise a force in the opposite direction. For this reason, perhaps, prior work (Byard et al., 2006; García-Meca & Sánchez-Ballesta, 2011) was unable to discern an effect of insider ownership on financial analysts on average. Our results suggest that insider ownership influences the information environment of firms and provide further evidence of the importance of considering insider ownership in understanding firm and analyst behavior.

In the following Section II, we discuss the related literature and develop our research hypotheses. We describe our data and research design in Section III. Section IV presents empirical results. Concluding remarks appear in Section V.

II. Literature Review and Hypothesis Development

The question of the impact of insider ownership on the forecasting environment of financial analysts is largely open in the literature. Only a few studies to date examine the association between insider ownership and financial analysts' forecast accuracy and dispersion. Byard et al. (2006) do not discern a statistically significant effect of CEO or director ownership on the accuracy of financial analysts' forecasts in a sample of U.S. firms. Similarly, García-Meca & Sánchez-Ballesta (2011) do not find that insider ownership, defined as board of director ownership, significantly affects financial analysts' forecast errors using a sample of Spanish firms. In a related study, Haw, Ho, Hu, & Wu (2010) on average also do not detect a significant effect of the divergence between the controlling owner’s control and cash flow rights on forecast accuracy and dispersion in a sample of East Asian and Western European firms.[3]

Given the influence of insiders on the quality and quantity of firms’ mandatory and voluntary financial reporting, this lack of significant findings is surprising. However, the difficulty in demonstrating the influence of insider ownership on analysts’ forecasting environment may stem from the nature of ownership itself. Insider ownership exerts two distinct forces. On one hand, insider ownership mitigates the classic agency conflict between owners and

[3] Haw et al. (2010) examine the divergence of control and cash flow rights of the controlling owner regardless of the identity of the controlling owner. The controlling owner may be the founding family, as is common in Asian countries, a financial institution, another corporation, or the firm could be state-owned. The authors caution that this mix of identities of the controlling owner may have affected their results, because incentives to exploit excess control rights for personal gain, or to manipulate firms’ disclosure practices to this end, are likely not constant across different types of controlling owners.

managers of the firm by aligning interests of outside shareholders and those in charge of managing the firm. This incentive-alignment effect should result in increased informativeness of mandatory financial reporting as managers' incentives to manipulate earnings for private benefits are reduced (Warfield, Wild, & Wild, 1995; Wang, 2006; Duellman et al., 2013). A positive association of managerial ownership with the extent of voluntary disclosures, documented for instance by Nasir & Abdullah (2004) in Malaysia, also supports positive incentive-alignment effects. Because the incentive-alignment effects of ownership provide insiders with incentives for improved quality and quantity of financial information, insider ownership may therefore exert positive effects on the information environment of financial analysts.

On the other hand, increasing levels of insider ownership may result in an entrenchment effect. High insider ownership reduces the threat of an ouster, and may enable corporate insiders to influence financial reporting and firm disclosure practices according to their self-interest (Lang, Lins, & Miller, 2004; Baik, Kang, & Morton, 2010). Insider ownership thus can create a Type II agency conflict between controlling and minority shareholders, where controlling shareholders may use their influence in the firm to extract perquisites at the cost of minority shareholders (Villalonga & Amit, 2009). Accordingly, users may place less credence in the accounting information produced by firms with high levels of insider ownership (Fan & Wong, 2002). The credibility of accounting information, however, has been shown to be positively associated with its informativeness (Teoh & Wong, 1993), and less credible accounting information can generate uncertainty in analysts' information environment (Barniv & Cao, 2009). Moreover, as insider ownership increases, insiders are also able to operate with greater discretion and to exercise tighter control over the flow of information (Fan & Wong, 2002). Specifically, concentrated ownership allows insiders to better manage who possesses information in the organization, and thereby to better guard against information leakage.[4] For these reasons, greater levels of insider ownership could decrease the informativeness of financial reporting, and the quality and quantity of firm voluntary disclosure.

In sum, theory and prior research therefore suggest that insider ownership may improve or diminish the quality of financial analysts' information environment, depending on whether incentive or entrenchment effects dominate. Yet typically these effects are confounded. The two separate forces—incentives and entrenchment—must be identified using only one variable: insider share ownership. We posit that prior studies potentially did not discern an association of insider ownership with financial analysts' forecast accuracy due to the countervailing effects of these two forces.

An analysis of dual-class companies offers a way around this problem. In a typical firm with a dual-class equity structure, one class of common stock has more votes per share than the other, while both classes generally have similar cash flow rights. The equity structure of dual-class firms therefore breaks the link between voting rights, which proxy for the entrenchment effect, and cash flow rights, which proxy for the incentive-alignment effect of ownership. By creating a material difference between the proportion of voting rights and cash-flow rights held, insiders are able to exert disproportionate control; insiders do not bear the economic consequences of their actions *pro rata* with their level of control over the firm.

Several studies examine how disproportionate insider control affects aspects of financial reporting quality. Francis et al. (2005) document that the informativeness of earnings, as

[4] We note that choosing higher levels of insider ownership to increase the ability to operate in greater secrecy must not necessarily be motivated by entrenchment, but can reflect a legitimate interest to more effectively guard sensitive information. Consistent with the latter explanation, Chemmanur, Paeglis, & Simonyan (2011) for instance provide evidence that firms may adopt strong anti-takeover measures, such as a dual-class structure of stock, to enable management’s long-term focus on value creation. However, regardless of whether tighter information control is used as a means of entrenchment, or to more strictly guard proprietary information, the information environment of financial analysts will likely be less rich.

indicated by the returns-earnings relationship, decreases as disproportionate insider control increases in a sample of U.S. dual-class firms. Fan & Wong (2002) present similar results and show that a separation of voting rights from cash flow rights, accomplished through cross-holdings and pyramid structures common in Asian economies, reduces earnings informativeness. Using samples of East Asian and Western European countries, the divergence of control from cash flow rights of the ultimate owner has also been shown to be positively associated with accruals based earnings management (Haw, Hu, Hwang, & Wu, 2004), classification shifting (Haw, Ho, & Li, 2011), and lowered earnings conservatism (Lim & Tan, 2009).

These findings in the literature strongly suggest that disproportionate insider control negatively affects the information environment of analysts. However, a few studies also suggest otherwise. For instance, because entrenched insiders are less beholden to shareholders, they have less reason to conceal or manage financial information. Pleasing investors by meeting earnings benchmarks matters less when insiders have enhanced influence over their compensation and job security. Consistent with this conjecture, Nguyen & Xu (2010) find that earnings management activities are associated positively with managerial cash flow rights, but negatively with managerial voting rights. Chen (2008) similarly finds lower earnings management in dual-class compared to single-class firms, but explains the finding with decreased capital-market pressure on entrenched managers who can focus on long-term value creation and hence have lowered incentives for short-term earnings manipulation.

While a lower concern about earnings may help earnings quality by reducing incentives for manipulation, the same lowered concern should in general reduce incentives to provide timely and/or comprehensive information to shareholders. That is, while the quality of some financial information may be higher, by the same logic the quantity of information supplied should be lower. Indeed, Tinaikar (2014) documents that the separation of voting and cash flow rights in dual-class firms is negatively associated with firms' voluntary compensation disclosures.

Overall, we therefore expect that disproportionate insider control is negatively associated with financial analysts' information environment. Further, we expect the incentive-alignment effects of ownership, which capture the sharing in the economic success of the firm through insider cash flow rights, to exercise a force towards higher quality information. Conversely, the entrenchment effect, which captures the extent to which insiders are isolated from the control by outside shareholders through insider voting rights, exerts a negative influence on financial analysts' information environment. When the information supplied to financial analysts is of inferior quality, analysts will tend to make less accurate forecasts (Lang & Lundholm, 1996; Hope, 2003). Accordingly, we hypothesize:

H1a: Disproportionate insider control is negatively associated with the accuracy of financial analysts' earnings forecasts.

H1b: Insider voting rights (insider cash flow rights) are negatively (positively) associated with the accuracy of financial analysts' earnings forecasts.

We similarly expect forecast dispersion to vary with the quality of the information environment. Forecast dispersion has been found to be associated with greater levels of information asymmetry (Lang & Lundholm, 1996; Barron & Stuerke, 1998; Irani & Karamanou, 2003) as well as a less-rich information environment (Byard et al., 2011). Further, Barron, Byard, Kile, & Riedl (2002) demonstrate that when financial information is less useful for predicting firm performance, the information contained in forecasts of individual analysts will tend to consist more of private knowledge than common knowledge. Accordingly, to the extent that the divergence of voting rights from cash flow rights reduces overall information quality, financial analysts will rely more on idiosyncratic information. Because dispersion reflects idiosyncratic error arising from reliance on private information (Barron et al., 1998), we hypothesize with respect to forecast dispersion:

H2a: Disproportionate insider control is positively associated with the dispersion of financial analysts' earnings forecasts.

H2b: Insider voting rights (insider cash flow rights) are positively (negatively) associated with the dispersion of financial analysts' earnings forecasts.

III. Data and Research Methodology

Sample

We construct a sample of U.S. dual-class firms from 2000 to 2012 following the approach of GIM (2010). That is, we first build a comprehensive list of potential dual-class firms from sources identified in GIM (2010).[5] We eliminate foreign and financial firms. For all remaining candidates, we access proxy statements and/or 10-Ks on the SEC's EDGAR database to verify the corporate structure and to determine insider ownership for each class of stock in every year.[6] We follow GIM (2010) in defining insider ownership comprehensively and include shares owned by family members, or trusts for the benefit of family members, as well as shares owned by parent or subsidiary corporations with board representation.[7] We also collect dividend data since classes of shares may not only differ with respect to voting, but also with respect to

[5] In addition, we include as candidates all firms reported as dual-class in the GIM (2010) sample, which spans 1995 to 2002. We thank Paul Gompers, Joy Ishii, and Andrew Metrick for generously sharing their list of dual-class companies underlying GIM (2010).

[6] Current SEC reporting requirements mandate firms to disclose the share ownership for each director as well as for all officers and directors as a group. We define corporate insiders as this disclosed group of officers and directors.

[7] The SEC-required ownership disclosures include options, warrants, deferred shares, and other "rights to stock" exercisable within 60 days of the date of disclosure. No distinction is made if such options are "in the money" and therefore likely to result in future ownership, or not. To maintain a clean measure of *actual* voting and cash flow rights owned by insiders, we screen the disclosures and compute ownership excluding all options and other rights.

their dividend rights. At this stage, we also collect governance characteristics included as controls: CEO-Chairman identity, board size, and number of independent directors. Through these procedures, we are confident to present a high-quality, comprehensive sample which likely approaches the population of U.S. non-financial dual-class firms in our sample period. Upon merging our dual-class data with Compustat, I/B/E/S, and Thomson Reuters, our final sample consists of 2,050 firm-years, representing 356 unique firms.[8]

Insider voting and cash flow rights

In companies with multiple classes of common stock, separate classes of shares may entitle their holders to a different number of votes per share and/or different dividend rights. Differences in the number of votes per share are the primary mechanism of creating a divergence between voting rights and cash flow rights in dual-class firms. A second control-enhancing mechanism is disproportionate board representation (Villalonga & Amit, 2009). We define insider voting rights (*VR*) as board election rights, i.e., the proportion of board seats insiders are able to command, regardless if this voting power stems from differences in the number of votes per share, from disproportionate board representation rights, or from both sources.[9] We compute insider cash flow rights (*CFR*) as fractional equity ownership, i.e., the percentage of shares held by corporate insiders to shares outstanding of all classes, weighted by dividend rights per class (Francis et al., 2005; GIM, 2010). Our primary measure of disproportionate insider control is the ratio of insider voting rights to insider cash flow rights (*WEDGE*).[10]

[8] We follow Horton, Serafeim, & Serafeim (2013) in choosing the consensus forecast that is calculated three months before the end of the reporting period as our base line. The sizes of our samples for alternate forecast horizons vary between 2,026 and 2,247 firm-years due to different number of forecasts available at different horizons.

[9] While infrequent, some companies differentiate voting rights according to the subject matter to be voted on. For instance, equal voting rights may exist for regular business, such as the confirmation of the auditor, but one class may have more votes per share when voting on takeover proposals or in the election of directors.

[10] The divergence of *VR* from *CFR* can be expressed as the ratio or difference of the two. The difference and ratio both have been used in prior work. We use the ratio of *VR* to *CFR* as our primary metric (Masulis et al., 2009) and examine the robustness of our results to using the difference as an alternate specification.

Research design

We measure our two dependent variables, analysts' forecast accuracy (*ACC*) and forecast dispersion (*DISP*), similar to Duru & Reeb (2002) and Byard et al. (2006). Specifically, we compute *ACC* as the absolute value of the difference between the mean I/B/E/S consensus forecast of annual earnings and the actual annual earnings reported by I/B/E/S, scaled by stock price on the day prior to the measurement of the I/B/E/S consensus forecast.[11] For ease of interpretation we multiply this scaled forecast error by -1, so that higher values of *ACC* indicate higher accuracy. *DISP* is the standard deviation of analysts' forecasts scaled by stock price on the day prior to the measurement of the I/B/E/S consensus forecast. Following Byard et al. (2011), we require that a minimum of two analysts are included in the computation of the consensus forecast to ensure that dispersion is meaningful.[12]

In addition to our variable(s) of interest, we include controls for firm, forecast, and corporate governance characteristics in our primary model:

$$\begin{aligned} ACC\ (DISP)_{it} = {} & \beta_0 + \beta_1 WEDGE_{it} + \beta_2 lnSIZE_{it} + \beta_3 LOSS_{it} + \beta_4 EPS\_VOL_{it} + \\ & \beta_5 HORIZON_{it} + \beta_6 FOLLOW_{it} + \beta_7 CEO\_CHAIR_{it} + \beta_8 BOARD\_SIZE_{it} + \\ & \beta_9 IND\_DIR_{it} + \beta_{10} INST\_OWN_{it} + \beta_k YEAR_{it} + \beta_j INDUSTRY_{it} + \varepsilon \qquad (1) \end{aligned}$$

A negative (positive) coefficient on the independent variable of interest, *WEDGE*, the extent of divergence of voting rights owned by insiders from cash flow rights owned by insiders, will provide evidence in favor of Hypothesis 1a (2a). We further investigate Hypotheses 1b and 2b by replacing *WEDGE* with the proportion of voting rights (*VR*) and cash flow rights (*CFR*)

[11] The consensus forecast here refers to the mean forecast. Our inferences are unchanged using the median forecast. Following Horton et al. (2013), we use a three months forecast horizon for our base line results. As reported below, our results are similar when we use alternative forecast horizons, i.e., the consensus forecasts made most recently, six months, or nine months before the end of the reporting period.
[12] As a test of robustness, we alternatively require a minimum of three or five forecasts to compute dispersion. While these choices reduce the sample size, results remain qualitatively unchanged.

owned by insiders for in model (1). We expect a negative (positive) coefficient on *VR* (*CFR*) when *ACC* is the dependent variable and the opposite signs when *DISP* is the dependent variable.

Our first set of controls reflects firm characteristics. *lnSIZE* is the natural log of market capitalization, *LOSS* represents an indicator variable equaling one if net income for the firm is negative, and zero otherwise, and *EPS_VOL* is the standard deviation of earning per share over the prior five-year period (or as many years as available). Second, we control for forecast characteristics. *HORIZON* is the number of calendar days between the I/B/E/S consensus forecast date and fiscal year end, and *FOLLOW* is the number of analysts following a firm, measured by the number of individual financial analyst forecasts included in the computation of the consensus forecast. As a third set of controls, we include the set of corporate governance characteristics used by Byard et al. (2006): *CEO_CHAIR* is an indicator variable equal to one if the CEO is also the chair of the board, and zero otherwise; *BOARD_SIZE* is the number of directors on the board; *IND_DIR* is the percentage of independent directors; and *INST_OWN* is the proportion of stock held by institutional investors.[13] Data for *CEO_CHAIR*, *BOARD_SIZE*, and *IND_DIR*, have been collected from firms' annual proxy statements; institutional ownership data (*INST_OWN*) is sourced from Thomson Reuters. Finally, we control for year and industry effects by including year and industry (Fama & French, 1997) indicator variables in all models.

In addition to examining single stage results from model (1), we conduct two additional analyses which control for potential sample selection bias or endogeneity. It is possible that dual-class firms are different from single-class firms with respect to characteristics other than dual-class status, and that such differences affect the association between insiders' disproportionate

[13] We do not include the proportion of independent directors serving on the audit committee (Byard et al., 2006). Following passage of the Sarbanes Oxley Act in 2002, all members of the audit committee must be independent directors. Accordingly, the variable would be invariant for the majority of our sample period, which ranges from 2000 to 2012.

control rights and the forecasting environment of financial analysts. Accordingly, analyses of a sample of dual-class firms could be affected by self-selection bias and, as a consequence, our findings may not extend to single-class firms. We therefore control for possible sample selection effects by using the Heckman (1979) methodology.

To this end, we estimate the following probit model, based on GIM (2010), which models the decision to establish a dual-class ownership structure at the time of a firm's going public based on various proxies for the availability of private benefits of control:

$$\begin{aligned}\Pr(DUAL) = \; & \beta_0 + \beta_1 \mathrm{NAME}_i + \beta_2 \mathrm{MEDIA}_i + \beta_3 \mathrm{SALESRANK}_i + \beta_4 \mathrm{PROFITRANK}_i + \\ & \beta_5 \%\mathrm{FIRMS}_i + \beta_6 \%\mathrm{SALES}_i + \beta_7 \mathrm{SALES/REGIONSALES}_i + \\ & \beta_k \mathrm{LISTINGYEAR}_i + \beta_j \mathrm{INDUSTRY}_i + \varepsilon \end{aligned} \tag{2}$$

The dependent variable *DUAL* is equal to one for firms electing a dual-class structure and zero otherwise. The independent variables are an indicator if at IPO the firm name contains a person's name; an indicator if the firm operated in the media industry at the time of its IPO;[14] the percentile rank of the firm's sales and profits in the year of its IPO relative to firms with the same IPO year; the percentage of all firms and all sales in the same Core Based Statistical Area (CBSA) in the year prior to the firm's IPO; the percentage of the firm's sales relative to the sales of all firms in the same CBSA in the year of its IPO; and indicator variables for the CRSP listing year and forty-eight Fama & French (1997) industries.[15] Results indicate that *NAME*, *MEDIA*, and *SALESRANK* (*SALES/REGIONSALES)* are significantly positively (negatively) associated

[14] Smart & Zutter (2003) and GIM (2010) suggest that dual-class shares are more common among firms operating in media industries. About twenty percent of firms in our dual-class sample are media firms compared to about three percent of such firms in the population.

[15] We include all variables identified in GIM (2010) in model (2), except for the state anti-takeover law variable (Gompers, Ishii, & Metrick, 2003), because necessary data to construct the index are not available after 2006. All variables are defined as in GIM (2010). We thank Paul Gompers, Joy Ishii, and Andrew Metrick for providing the name indicator variable for the universe of firms in GIM (2010). For firms not included in their study, we follow their procedure to identify the presence of a family name in the firm name. If a firm is not headquartered in an identifiable CBSA, we use the county as the firm's geographic area.

with the probability of choosing a dual-class structure.[16] Following Amoako-Adu, Baulkaran, & Smith (2014) and McGuire, Wang, & Wilson (2014), we use the coefficient estimates from model (2) to construct an inverse Mills ratio (*INV_MILL*, Heckman, 1979), which we include as a control for sample selection bias in model (1).

Endogeneity is a concern in any study of firms' corporate governance or ownership characteristics. We therefore perform a two-stage least squares (2SLS) analysis and regress *WEDGE*, *VR*, and *CFR*, and on a set of firm-level instrumental variables identified in Khalil, Magnan, & Cohen. (2008) and Haw et al. (2010):

$$WEDGE\ (VR, CFR)_{it} = \beta_0 + \beta_1 FAMILY_{it} + \beta_2 FIN\_NEED_{it} + \beta_3 \Delta SALES_{it} + \beta_4 lnAT_{it} + \beta_5 ROA_{it} + \beta_6 INTANG_{it} + \beta_7 MEDIA_{it} + \beta_8 FIRMAGE_{it} + \varepsilon \qquad (3)$$

FAMILY indicates a family firm defined as a firm where the founder and/or his or her descendants own 25% or more of the firm's voting rights (Andres, 2008). We research firms' founders from SEC filings and publicly available sources. *FIN_NEED* is a firm's need for financing, proxied by ROE / (1-ROE) as in Khalil et al. (2008); *ΔSALES* is percentage sales growth; *lnAT* is the natural log of total assets; *ROA* is income before extraordinary items divided by total assets; *INTANG* is the percentage of intangible assets, computed as 1 – [(net PPE + inventories) / total assets]; *MEDIA* is indicator variable that takes a value of one if the firm operates in the media industry, and zero otherwise; *FIRMAGE* is the age of the firm, based on its founding year which we collected from publicly available information. In our third set of regressions, we replace *WEDGE*, *VR*, and *CFR* and with their predicted values from the first-stage regression model (3) as an explicit control for endogeneity.

[16] For brevity, we do not tabulate results for model (2). Results are available upon request.

IV. Empirical Analysis

Descriptive Statistics

Table 1 reports descriptive statistics concerning our baseline sample of dual-class firms, which encompasses 356 unique firms contributing a total of 2,050 firm-years. The clear majority of observations (n = 1,910) have two classes of shares. In the remaining 140 cases more than two classes of common stock exist (6.8% of the sample). In 82.2% of cases, differences exist in the number of votes per share in the election of the board of directors. In addition, voting schemes whereby each class of stock elects a certain number of directors are common, as 595 observations (29.0% of sample) maintain this arrangement. The proportion of firms with unequal votes across classes of shares and disproportionate board election rights exceeds 100% due to firms which adopt both of these elements.

While voting rights differ substantially, all shares are generally entitled to receive equal dividends. In only 10.8% of observations, do dividend rights differ across classes of stock. Also, in only 32 observations (1.6% of sample) we find that shares that are identical with respect to voting and cash flow rights, but differ in other characteristics only, for instance, differences may pertain to redemption rights or voting rights other than in the election of the board. This indicates that firm overwhelmingly chose dual-class structures to create a discrepancy between voting and cash-flow rights across classes of shares.

With respect to the voting and ownership structure, insiders on average own 79.0% of the superior, but only 12% of the inferior, shares outstanding.[17] The median superior class of stock

[17] For ease of exposition, we exclude firms with more than two classes of common stock from the descriptive statistics of the voting and ownership structure.

has ten votes per share, whereas the median inferior share only commands one vote.[18] Among firms that maintain disproportionate board representation schemes, the median (mean) proportion of directors that are elected by the superior class is 66.7% (62.9%), the remaining number of directors being elected by the inferior class. "Benign" dual-class structures, where insiders predominantly hold the inferior class of shares, are rare. In only 74 observations (3.6% of sample) do we find that insiders own fewer voting rights than cash flow rights.

Finally, with respect to our variables of interest, insiders of the firm control median (mean) voting rights in the election of the board of directors of 59.5% (55.9%) compared to cash flow rights of 20.8% (26.6%). The ratio of voting rights to cash flow, our primary measure of the degree of separation between the two, shows that insiders command a median (mean) of 2.35 (2.69) times the number of voting rights as they do cash flow rights. The median (mean) difference between the proportion of voting and cash flow rights owned by insiders, our alternate measure of the wedge between the two, is 29.3% (28.0%).

[Insert Table 1 about here]

Descriptive statistics of our control variables are summarized in Table 2. Median (mean) *lnSIZE* of our sample firms is 20.42 (20.47), which corresponds to a median (mean) market capitalization of 740 (3,877) million dollars. Loss years represent 18.5% of our sample, which spans two financial crises. Descriptive statistics of the board characteristics variables reveal that the CEO serves as chairman of the board in 51.6% of our sample. *BOARD_SIZE* is fairly uniform across our sample with a mean and median of nine directors and a standard deviation of

[18] The mean number of votes per share for the superior class is affected by eleven total observations from three firms, Lin TV Corporation, Virgin Mobile US, and Prodigy Communications. These firms only have one or two superior shares outstanding, but each share entitles its bear to millions of votes. For instance, Prodigy Communications reports 70.245 million shares of Class A common stock and only one share of Class B common stock outstanding in its 2001 proxy statement. However, its only share of Class B common stock entitles its owner to 51.843 million votes per share. Removing, these observations does not affect our results.

2.33. On average, 62% of the board is comprised of independent directors, very similar to the 64.3% mean independent directors reported by Byard et al. (2006). Lastly, median (mean) institutional ownership amounts to 72.5% (68.8%).

[Insert Table 2 about here]

Regression Results

Column A of Table 3 displays the results of our single stage regression of forecast accuracy on disproportionate insider control and controls, model (1). The coefficient on *WEDGE* is negative and significant (t = -3.39, p < 0.01, two-tailed). This finding supports our hypothesis H1a, indicating that the accuracy of financial analysts' earnings forecasts decrease as the extent of separation of voting rights from cash flow rights increases. With respect to our controls, firm size is positively and significantly associated with forecast accuracy, whereas *LOSS*, *EPS_VOL*, *FOLLOW*, and *DISP* all show a significant negative association, consistent with prior research. While most of the firm and forecast characteristics controls are significant, we find little support for our governance characteristics controls except for *INST_OWN*, which displays a significant positive association with accuracy.[19]

We repeat our analysis replacing *WEDGE* with its two individual components, *VR* and *CFR*. Results of a regression of forecast accuracy on insider voting and cash flow rights and controls are reported in Table 3, Column B. The coefficient on *VR* is negative and significant (t = -2.45, p = 0.015, two-tailed), while the coefficient on *CFR* is positive and significant (t = 3.55, p < 0.01, two-tailed). These results confirm that forecast accuracy increases in insider cash flow

[19] The positive association of institutional ownership *INST_OWN* with forecast accuracy *ACC* is consistent with the findings by Ljungqvist et al. (2007) and Kerl & Ohlert (2015), but opposite to Byard et al. (2006). Differences in the data source may explain the divergent results. Kerl & Ohlert (2015) rely on FactSet/LionShares for their institutional ownership data. Byard et al. (2006) use RiskMetrics data. Ljungqvist et al. (2007) retrieve institutional ownership data from Thomson Reuters, which is also our source. A positive association may be more plausible because institutional investors positively affect reporting quality by acting as effective monitors (Velury & Jenkins, 2006).

rights, and decreases in insider voting rights, supporting our hypothesis H1b. Results for our control variables remain qualitatively unchanged from those previously reported.

In columns C and D of Table 3 we report results specifically controlling for sample selection bias by employing the Heckman (1979) methodology, and adding *INV_MILL*, the inverse Mills ratio derived from an estimation of model (2), to model (1). Results are not affected by using this alternate approach; *WEDGE*, *VR* and *CFR* all remain statistically significant at similar levels.

To control for endogeneity, we also replicate our analyses using a 2SLS approach. In the first-stage regression, we regress the potentially endogenous variables *WEDGE*, *VR*, and *CFR* on the set of instrumental variables identified in Khalil et al. (2008) and Haw et al. (2010), model (3). The explanatory power of the first stage regression is reasonable, with an R^2 ranging from 9.4% to 17.6% and comparable to the R^2 of approximately 14.3% achieved by Khalil et al. (2008). With the exception of *ROA* and *ΔSALES*, all variables are significant. The first-stage regression results are summarized in Table 4.

[Insert Table 4 about here]

Next, we re-estimate model (1) replacing *WEDGE*, *VR*, and *CFR* with their predicted values from model (3). Results, presented in Columns E and F of Table 3, indicate that all variables remain significant at comparably (strong) levels. The coefficients of *WEDGE*, *VR*, and *CFR* are greater in the second stage of the 2SLS regression compared to the results from our single-stage approach. We interpret these increases as evidence for a reduction in single-equation bias from using a 2SLS approach. (Beaver, McAnally, & Stinson, 1997).

Regressions results of model (1) using financial analysts' forecast dispersion, *DISP*, as the dependent variable are in reported in Table 5. Column A shows that the coefficient on

WEDGE is positive and significant (t = 3.54, p < 0.01, two-tailed) indicating that the dispersion of financial analysts' earnings forecasts increases in disproportionate insider ownership. This finding supports our hypothesis H2a. In general, results for the control variables in the dispersion regression complement those for accuracy, such that characteristics which increase accuracy reduce dispersion.

In Column B of Table 5, we report results including *VR* and *CFR* in place of *WEDGE*. *VR* is positively related to *DISP* (t = 2.04, p = 0.042, two-tailed), while the association of *CFR* with *DISP* is negative (t = -3.23, p < 0.01, two-tailed). We thus find support also for H2b as insider voting rights (cash flow rights) are significantly positively (negatively) related to forecast dispersion. Results for the control variables are not materially different from the results in Column A.

[Insert Table 5 about here]

In columns C and D of Table 5 we show results for our regressions of analysts' forecast dispersion on *WEDGE*, *VR*, and *CFR* when controlling for sample selection bias. Similarly, columns E and F of Table 5 show results of our 2SLS approach to adjust for endogeneity. In either case results are consistent with our single stage regression results and display comparable levels of statistical significance.

Taken together, results in Tables 3 and 5 confirm that insider voting and cash flow rights exert divergent effects on the accuracy and dispersion of financial analysts' earnings forecasts. Specifically, we find a negative (positive) association of the degree of separation of insider voting from cash flow rights with forecast accuracy (dispersion). This finding is in line with Francis et al. (2005) and Fan & Wong (2002), who report that the informativeness of mandatory financial reporting decreases as the separation of voting rights from cash flow rights increases. A possible alternate effect, suggested by Nguyen & Xu (2010) or Chen (2008) who find that earnings management activities decrease in firms characterized by disproportionate control and that the forecasting environment hence could be improved, appears unlikely given our findings

Also supporting an entrenchment explanation, we find that insider voting rights decrease forecast accuracy, and increase forecast dispersion, while, confirming an incentive story, insider cash flow rights increase accuracy, and decrease dispersion. These results are consistent with insiders reducing the amount and/or reliability of information provided as their entrenchment from voting rights increases. A reduced amount or reliability of public information decrease financial analysts' ability to issue accurate forecasts, and increases the need to substitute private for common information (Barron et al., 2002) with a resulting increase in forecast dispersion. Conversely, a richer information environment associated with insider cash flow rights enhances forecast accuracy and decreases dispersion.

Taken as a whole, our results appear to conflict with Haw et al. (2010), who pose similar hypotheses concerning the association of control-cash flow divergence and analysts' forecast properties, but do not find a significant effect on average. Our strong divergent results using a contemporary sample of U.S. dual-class firms may be due to the fact that Haw et al. (2010) document that the association between control-cash flow divergence and forecast properties is contingent on the effectiveness of a country's legal regime.[20] More importantly, we focus on insider ownership, whereas Haw et al. (2010) examine control-cash flow divergence of the controlling owner regardless of whether this owner is an insider or outsider, a family, the state, a

[20] Haw et al.'s (2010) sample comprises firms from nine East Asian and thirteen Western European countries in the 1990 to 1996 period. We note as a limitation to our study that our results may not generalize to other jurisdictions. However, because the U.S. is widely seen as offering a high quality legal system with strong enforcement, we posit that the negative effects of insider control-cash flow divergence on analysts' forecast properties likely are only more pronounced in countries with weaker legal regimes. Indeed, Haw et al. (2010) find negative effects of control-cash flow divergence of the controlling owner only in a set of countries with weak legal institutions.

financial institution, or other corporation. It is likely that our narrower focus on insider ownership enables us to find stronger results, because motives and incentives to influence the information environment are likely not constant across different types of ultimate owners.

Additional analyses

Alternate forecast horizons

We follow Horton et al. (2013) and Byard et al. (2006) and examine the robustness of our results to different forecast horizons. Our base line results, reported in Tables 3 and 5, are for forecasts issued on average three months prior to the end of the reporting period. We choose one shorter horizon, the most recent forecast issued before the end of the reporting period, as well as two longer forecasting periods: forecasts issued on average six and nine months before the end of the reporting period. Table 6 (Table 7) show regression results of *ACC* (*DISP*) on *VR* and *CFR* across these three alternate forecasts horizons. Overall, results are not affected. For forecast accuracy, the coefficients are negative on *VR*, and positive on *CFR*, and remain statistically significant across all alternate horizons (Table 6). In the dispersion regressions, *VR* dips slightly below two-tailed significance at the most recent horizon. All other results remain similar across alternate horizons with positive (negative) coefficients on *VR* (*CFR*) and comparable significance levels (Table 7).

[Insert Tables 6 and 7 about here]

For brevity, we do not tabulate results of these analyses for *WEDGE*. All results for *WEDGE* are similar to those reported in Tables 3 and 5 when using alternate forecast horizons. In the accuracy regressions, the effect of *WEDGE* is significant at the $p < .01$ level across all three alternate horizons. In the dispersion regressions, *WEDGE* is significant at the $p < .05$ level for the most recent horizon, and at the $p < .01$ level for the six and nine month horizons.

Changes model

Our theoretical framework assumes a particular direction of causality, where an increasing divergence of insider voting rights from their cash flow rights leads to lower forecast accuracy and increased forecast dispersion. We acknowledge that our empirical analyses presented thus far mainly demonstrate association. To shed additional light on the relationship between disproportionate insider control and financial analysts' forecasting environment, we therefore also estimate the following changes specification of model (1):

$$\begin{aligned}\Delta ACC\ (\Delta DISP)_{it} = \quad & \beta_0 + \beta_1 \Delta WEDGE_{it} + \beta_2 \Delta lnSIZE_{it} + \beta_3 \Delta LOSS_{it} + \beta_4 \Delta EPS\_VOL_{it} + \\ & \beta_5 \Delta HORIZON_{it} + \beta_6 \Delta FOLLOW_{it} + \beta_7 \Delta CEO\_CHAIR_{it} + \\ & \beta_8 \Delta BOARD\_SIZE_{it} + \beta_9 \Delta IND\_DIR_{it} + \beta_{10} \Delta INST\_OWN_{it} + \varepsilon \qquad (4)\end{aligned}$$

In model (4), we regress changes in financial analysts' forecast accuracy (and dispersion) on contemporaneous changes in insiders' disproportionate control ($\Delta WEDGE$) and controls. We also examine the association of the prior year change in disproportionate control, $\Delta WEDGE_{t-1}$, with the contemporaneous change in *ACC* and *DISP*. Results are presented in Table 8.

[Insert Table 8 about here]

The analysis reveals that, consistent with the results for our levels model, changes in disproportionate insider control, modelled either as the contemporaneous or lagged change, are significantly associated with corresponding changes in financial analysts' forecast accuracy and dispersion. Our finding that the lagged change in disproportionate control is more strongly associated with accuracy and dispersion than the contemporaneous change suggests that changes in disproportionate control may affect changes in firms' information practices with delay.

Additional endogeneity tests

Endogeneity is a concern in any study of firms' corporate governance or ownership characteristics. One concern with regard to endogeneity is the possibility of omitted variables that may jointly impact financial analysts' forecasting environments and *WEDGE*, *VR*, and *CFR*. Our primary approach to address the endogeneity problem is by means of the two-stage least squares analysis presented above. All results proved robust to this approach. However, to further address the possibility that insiders' incentives to select a particular ownership structure drive not only the choice of ownership structure, but also affect the forecasting environment of financial analysts', we conduct two additional tests.

As suggested by Masulis et al. (2009), one important factor likely affecting insiders' incentives to choose a specific ownership structure is the availability of private benefits of control. Insiders of firms with a higher capacity for private benefits may be more likely to establish a dual-class structure which enables the extraction of perquisites. At the same time, insiders interested in extracting private benefits of control potentially have increased incentives to reduce publicly available information when more opportunities to extract private benefits exist. Following Masulis et al. (2009), we control for existent opportunities for private benefits of control by including the variables predicting dual-class status based on a firm's private benefits capacity from model (2) directly in model (1). Moreover, related to the idea that a firm's private benefits' capacity may be affecting both the selection of its ownership structure as well as analysts' forecasting environment, we also add an explicit control for firms' compensation structure, CEO total compensation (Execucomp variable *TDC1*), to model (1). Our inferences are unchanged when including these two additional controls for firms' private benefits capacity to our model.

Finally, to address concerns of reverse causality, we follow Masulis et al. (2009) and replace the annual values of *WEDGE* in our levels regressions with its value in the year of a firm's first appearance in our sample, as well as with its prior year (lagged) value. Results are not materially affected by these alternate specifications.

Other tests

We also test a number of alternative definitions of our variables. We use the difference between *VR* and *CFR*, instead of the ratio of the two, as an alternate specification of *WEDGE* (GIM, 2010). We also replicate all regressions requiring a minimum of three or five observations for the computation of dispersion. Instead of the mean, we also use the median consensus earnings forecasts in our forecast accuracy models. With respect to our control variables, we alternatively use the natural logarithm of the book value of assets as a control for firm size; define *HORIZON* relative to the earnings announcement date, instead of relative to the fiscal period end date; use the log of *FOLLOW* and *BOARD_SIZE*, instead of their count values; and use NAICS industry controls instead of Fama & French (1997) industries. To ensure that results are not driven by, or sensitive to, the inclusion of a small number of observations with uncommon dual-class structures where difference in between classes of shares do not pertain to voting or cash flow rights (32 observations), or where insiders primarily hold the inferior class of shares (73 observations), we rerun all analyses excluding these subsets of firms. Results are not affected by these choices.

V. Summary and Conclusion

Using a comprehensive hand-collected sample of U.S. dual-class firms, we assess the distinct effects of insider voting and cash flow rights on analyst forecast accuracy and dispersion.

Some studies establish that the quality (Francis et al., 2005) and credibility (Fan & Wong, 2002) of financial reporting decreases as the wedge between insider voting and cash flow right widens. However, other studies provide evidence of improved financial reporting quality in the presence of disproportionate insider control (Ngyuen & Xu, 2010; Chen, 2008). We find that disproportionate insider control is negatively (positively) associated with the accuracy (dispersion) financial analysts' forecasts. Moreover, insider cash flow rights (insider voting rights) are positively (negatively) associated with financial analyst forecast accuracy, and negatively (positively) associated with forecast dispersion. These results are robust to controls for sample-selection bias and endogeneity, and not affected by choosing alternative forecast horizons or estimation techniques.

In contrast to prior studies, our results demonstrate that insider ownership does affect the information environment of financial analysts, and that the opposing incentive-alignment and entrenchment effects cause it to do so in a non-monotonic fashion. Forecast accuracy (dispersion) increases (decreases) in insider cash flow rights likely due to an incentive-alignment effect, which positively affects the quality of financial reporting, while at the same time forecast accuracy (dispersion) decreases (increases) in insider voting rights due to an entrenchment effect, which has been shown to negatively impact financial reporting. By delineating these two distinct effects of insider ownership on forecast accuracy and dispersion, we augment, and offer an explanation for, prior research that did not discern a significant association of insider ownership with financial analysts' earnings forecast accuracy (Byard et al., 2006; Haw et al., 2010; García-Meca & Sánchez-Ballesta, 2011).

We use a sample of U.S. dual-class firms primarily as a laboratory that enables us to separately assess the incentive and entrenchment effects of insider ownership, which are confounded in single-class firms. However, our examination of forecast accuracy and dispersion in dual-class firms is also of interest as a study of this particular subset of firms in its own right. A spate of high profile IPOs of technology firms with a dual-class structure of common stock has recently thrust this equity structure to the fore.[21] Dual-class firms, which comprised about 6% of U.S. public firms in the 1995 to 2002 period (GIM, 2010), now account for 8.7% of companies included in the Russell 3000 index (Equilar, 2015). In 2013, 13.6% of all U.S. firms conducting an initial public offering (IPO) adopted a dual-class structure of stock (Equilar, 2015). This increasing prominence of dual-class firms calls for a better understanding of their characteristics, especially with respect to these firms' information behavior and their financial reporting quality. In the light of our findings, particular care appears to be justified in the generation of earnings forecasts for dual-class firms. Likewise, the decreased reliability of earnings forecasts as disproportionate insider control increases suggests particular caution from the investing public relying on financial analysts' earnings forecasts for these firms.

[21] Google started the trend with its IPO in 2004, followed by other upstarts, such as LinkedIn, Zynga, Groupon, Yelp, and Facebook.

References

Amoako-Adu, B., Baulkaran, V., & B. Smith. (2014). Analysis of dividend policy of dual and single class U.S corporations. *Journal of Economics & Business, 72,* 1-29

Andres, C. (2008). Large shareholders and firm performance–an empirical examination of founding-family ownership. *Journal of Corporate Finance, 14*(4), 431-445.

Baik, B., Kang, J.-K., & R. Morton. (2010). Why are analysts less likely to follow firms with high managerial ownership? *Journal of Accounting, Auditing & Finance, 25*(2), 171-200.

Barniv, R., & Cao, J. (2009). Does information uncertainty affect investors' responses to analysts' forecast revisions? An investigation of accounting restatements. *Journal of Accounting & Public Policy, 28*(4), 328-348.

Barron, O., Byard, D., Kile, C., & Riedl, E. (2002). High-technology intangibles on analysts' forecasts. *Journal of Accounting Research, 40*(2), 289-312.

Barron, O., Kim, O., Lim, S., & Stevens, D. (1998). Using analysts' forecast to measure properties of analysts' information environment. *The Accounting Review, 73*(4), 421-433.

Barron, O., & Stuerke, P. (1998). Dispersion in analysts' earnings forecasts as a measure of uncertainty. *Journal of Accounting, Auditing, & Finance, 13*(3), 245-270.

Beaver, W., McAnally, M., & Stinson, C. (1997). The information content of earnings and prices: A simultaneous equation approach. *Journal of Accounting and Economics 23*(1), 53-81.

Byard, D., Li, Y., & Weintrop, J. (2006). Corporate governance and the quality of financial analysts' information. *Journal of Accounting & Public Policy, 25*(5), 609-625.

Byard, D., Li, Y., & Yu, Y. (2011). The effect of mandatory IFRS adoption on financial analysts' information environment. *Journal of Accounting Research, 49*(1), 69-96.

Chemmanur, T., Paeglis, I., & Simonyan, K. (2011). Management quality and antitakeover provisions. *Journal of Law & Economics, 54*(3), 651-692.

Chen, F. (2008). *Capital market pressures and earnings management: Evidence from U.S. dual-class firms* (Doctoral dissertation, Columbia University, New York, NY, USA). Retrieved from http://gradworks.umi.com/33/05/3305207.html

DeAngelo, H., & DeAngelo, L. (1985). Managerial ownership of voting rights. A study of public corporations with dual-classes of common stock. *Journal of Financial Economics, 14*(1), 33-69.

Demsetz, H., & Lehn, K. (1985). The structure of corporate ownership: causes and consequences. *Journal of Political Economy, 93*(6), 1155-77.

Duellman, S., Ahmed, A., & Abdel-Meguid, A. (2013). An empirical analysis of the effects of monitoring intensity on the relation between equity incentives and earnings management. *Journal of Accounting & Public Policy, 32*(6), 495-517.

Duru, A., & Reeb, D. (2002). International diversification and analysts' forecast accuracy and bias. *The Accounting Review, 77*(2), 415-433.

Equilar. 2015. Executive compensation and governance considerations in the initial public offering. Retrieved from http://www.equilar.com/publications/articles/executive-compensation -and-governance-considerations-in-the-initial-public-offering.

Fama, E., & French, K. (1997). Industry costs of equity. *Journal of Financial Economics, 43*(2), 153-193.

Fan, J. P., & Wong, T. J. (2002). Corporate ownership structure and the informativeness of accounting earnings in East Asia. *Journal of Accounting & Economics, 33*(3), 401-425.

Field, L., & Karpoff, J. (2002). Takeover defenses of IPO firms. *The Journal of Finance, 57*(5), 1857-1889.

Francis, J., Schipper, K., & Vincent, L. (2005). Earnings and dividend informativeness when cash flow rights are separated from voting rights. *Journal of Accounting & Economics, 39*(2), 329-360.

García-Meca, E., & Sánchez-Ballesta, J. (2011). Ownership structure and forecast accuracy in Spain. *Journal of International Accounting, Auditing & Taxation, 20*, 73-82.

Gompers, P. A., Ishii, J. L., & Metrick, A. (2003). Corporate governance and equity prices. *Quarterly Journal of Economics, 118*(1), 107-155.

Gompers, P. A., Ishii, J., & Metrick, A. (2010). Extreme governance: an analysis of dual-class firms in the United States. *Review of Financial Studies, 23*(3), 1051-1088.

Gow, I., Ormazabal, G., & Taylor, D. (2010). Correcting for cross-sectional and time-series dependence in accounting research. *The Accounting Review, 85*(2), 483-512.

Harris, M., & Raviv, A. (1988). Corporate governance: voting rights and majority rules. *Journal of Financial Economics, 20*, 203-235.

Haw, I. M., Ho, S. S. M., & Li, A. Y. (2011). Corporate governance and earnings management by classification shifting. *Contemporary Accounting Research, 28*(2), 517-553.

Haw, I. M., Ho, S. S. M., Hu, B., & Wu, W. (2010). Analysts' forecast properties, concentrated ownership and legal institutions. *Journal of Accounting, Auditing & Finance, 25*(2), 235-259.

Haw, I. M., Hu, B., Hwang, L. S., & Wu, W. (2004). Ultimate ownership, income management and legal and extra-legal institutions. *Journal of Accounting Research, 42*(2), 423-462.

Heckman, J. J. (1979). Sample selection bias as a specification error. *Econometrica, 47*(1), 153-162.

Hope, O. (2003). Disclosure practices, enforcement of accounting standards, and analysts' forecast accuracy: an international study. *Journal of Accounting Research, 41*(2), 235-272.

Horton, J., Serafeim, G., & Serafeim, I. (2013). Does mandatory IFRS adoption improve the information environment? *Contemporary Accounting Research, 30*(1), 388-423.

Imhoff, E., & Lobo, G. (1992). The effect of ex ante earnings uncertainty on earnings response coefficients. *The Accounting Review*, 67(2), 427-439.

Irani, A., & Karamanou, I. (2003). Regulation Fair Disclosure, analyst following, and analyst forecast dispersion. *Accounting Horizons, 17*(1), 15-29.

Kerl, A., & Ohlert, M. (2015). Star-analysts' forecast accuracy and the role of corporate governance. *Journal of Financial Research*, 38(1), 93-120.

Khalil, S., Magnan, M. L., & Cohen, J. R. (2008). Dual-class shares and audit pricing: evidence from the Canadian markets. *Auditing: A Journal of Practice & Theory, 27*(2), 199-216.

Lang, M. H., Lins, K. V., & Miller, D. P. (2004). Concentrated control, analyst following, and valuation: do analysts matter most when investors are protected least? *Journal of Accounting Research, 42*(3), 589-623.

Lang, M., & Lundholm, R. (1996). Corporate disclosure policy and analyst behavior. *The Accounting Review, 71*(4), 467-492.

La Porta, R., Lopez-de-Silanes, F., Shleifer, A., & Vishny, R. (2000). Investor protection and corporate governance. *Journal of Financial Economics, 58*, 3-27.

Lim, C. Y., & Tan, M. S. (2009). Control divergence, timeliness in loss recognition, and the role of auditor specialization: evidence from around the world. *Journal of Accounting, Auditing & Finance*, *24*(2), 295-332.
Ljungqvist, A., Marston, F., Starks, L.T., Wei, K.D., & Yan, H. (2007). Conflicts of interest in sell-side research and the moderating role of institutional investors. *Journal of Financial Economics, 85*(2), 420-456.
Masulis, R., Wang, C., & Xie, F. (2009). Agency problems at dual class companies. *Journal of Finance, 64*(4), 1697-1727.
McGuire, S., Wang, D., & Wilson, R. (2014). Dual class ownership and tax avoidance. *The Accounting Review, 89*, 1487-1516.
Morck, R., Shleifer, A., & Vishny, R. (1988). Management ownership and market valuation. *Journal of Financial Economics, 20*, 293-315.
Nasir, N., & Abdullah, S. (2004). Voluntary disclosure and corporate governance among financially distressed firms in Malaysia. *Financial Reporting, Regulation and Governance, 3*, 95-139.
Nguyen, V. T., & Xu, L. (2010). The impact of dual class structure on earnings management activities. *Journal of Business Finance & Accounting, 37*(3-4), 456-485.
Singh, M., & Davidson III, W. (2003). Agency costs, ownership structure and corporate governance mechanisms. *Journal of Banking & Finance, 27*(5), 793-816.
Smart, S. B., & Zutter, C. J. (2003). Control as a motivation for underpricing: a comparison of dual and single-class IPOs. *Journal of Financial Economics*, *69*(1), 85-110.
Teoh, S. W., & Wong, T. J. (1993). Perceived auditor quality and the earnings response coefficient. *The Accounting Review*, *68*(2), 346-366.
Tinaikar, S. (2014). Voluntary disclosure and ownership structure: an analysis of dual-class firms. *Journal of Management & Governance, 18*(2), 373-417.
Velury, U., & Jenkins, D. (2006). Institutional ownership and the quality of earnings. *Journal of Business Research*, *59*(9), 1043-51.
Villalonga, B., & Amit, R. (2009). How are U.S. family firms controlled? *The Review of Financial Studies*, *22*(8), 3047-3091.
Wang, D. (2006). Founding family ownership and earnings quality. *Journal of Accounting Research, 44*(3), 619-656.
Warfield, T., Wild, J., & Wild, K. (1995). Managerial ownership, accounting choices and informativeness of earnings. *Journal of Accounting & Economics, 20*(1), 61-91.

Table 1

Descriptive Statistics of Dual-Class Firm Sample

	N	Mean	Std. Dev.	Q1	Median	Q3
Equity structure						
More than two classes of shares	2,050	0.068	0.252	0.000	0.000	0.000
Unequal number of votes per class	2,050	0.822	0.382	1.000	1.000	1.000
Disproportionate board representation	2,050	0.290	0.454	0.000	0.000	1.000
Unequal dividends per class	2,050	0.108	0.311	0.000	0.000	0.000
No difference in voting or dividend rights	2,050	0.016	0.124	0.000	0.000	0.000
Voting and ownership structure						
Votes per share superior class*	1,910	27,462	1,186,257	1.000	10.000	10.000
Percent of superior shares owned by insiders*	1,910	0.790	0.305	0.681	0.954	1.000
Votes per share, inferior class*	1,910	0.850	0.351	1.000	1.000	1.000
Percent of inferior shares owned by insiders*	1,910	0.120	0.188	0.012	0.040	0.138
Proportionate board, elected by superior class**	555	0.629	0.172	0.547	0.667	0.727
Proportionate board, elected by inferior class**	555	0.367	0.170	0.273	0.333	0.434
Negative divergence: insider CFR exceed insider VR	2,050	0.036	0.185	0.000	0.000	0.000
Variables of interest						
VR	2,050	0.559	0.267	0.361	0.595	0.755
CFR	2,050	0.266	0.207	0.105	0.208	0.390
WEDGE (ratio)	2,050	2.690	1.409	1.508	2.349	3.680
WEDGE (difference)	2,050	0.280	0.164	0.150	0.293	0.432

The sample represents 2,050 firm-years (356 unique firms) used in the regressions of forecast accuracy and dispersion at the three months forecast horizon (baseline results; Tables 3 and 5). For parsimony of exposition, we eliminate observations with more than two classes of shares from certain descriptive statistics such that inferior (superior) refers to the inferior (superior) class of two classes of stock. Variable definitions: *VR* is the proportion of voting rights owned by insiders; *CFR* is the proportion of dividend rights owned by insiders. *WEDGE* is the divergence of insider voting from cash flow rights computed as the ratio of, or difference between, *VR* and *CFR*. * indicates descriptive statistics excludes 140 observations (6.8% of sample) with more than two classes of stock. ** indicates descriptive statistics for sub-set of firms with disproportionate board representation rights (n = 596) excludes 40 observations (6.7%) with more than two classes of stock.

Table 2

Descriptive Statistics of Select Variables

	Mean	Std. Dev.	Q1	Median	Q3
Dependent variables					
ACC	-0.060	0.427	-0.014	-0.004	-0.001
DISP	0.017	0.167	0.001	0.002	0.006
Firm characteristics controls					
lnSIZE	20.474	1.662	19.302	20.423	21.481
LOSS	0.185	0.388	0.000	0.000	0.000
EPS_VOL	0.237	3.614	0.012	0.023	0.059
Forecast characteristics controls					
FOLLOW	8.010	6.191	3.000	6.000	11.000
HORIZON	74.493	2.080	73.000	74.000	76.000
Governance characteristics controls					
CEO_CHAIR	0.516	0.500	0.000	1.000	1.000
BOARD_SIZE	8.953	2.327	7.000	9.000	10.000
IND_DIR	0.619	0.162	0.500	0.625	0.727
INST_OWN	0.688	0.221	0.558	0.725	0.860

The sample represents 2,050 firm-years (356) unique firms). Variable definitions: *ACC* is analyst forecast accuracy, calculated as (|Mean Forecast – Actual EPS| / Stock Price) multiplied by -1; *DISP* is the standard deviation of analyst forecasts scaled by stock price; *lnSIZE* is the natural logarithm of market value; *LOSS* is an indicator variable if net income is negative, and zero otherwise; *EPS_VOL* is the standard deviation of EPS over the prior five years, scaled by the stock price; *FOLLOW* is the number of analysts' forecasts used in computing the consensus forecast; *HORIZON* is the number of days between the I/B/E/S consensus forecast date and fiscal year end; *CEO_CHAIR* is an indicator variable equal to one if the CEO is also the chair of the board, and zero otherwise; *BOARD_SIZE* is the number of directors on the board; *IND_DIR* is the percentage of independent directors; *INST_OWN* is the proportion of stock held by institutional investors.

Table 3

The Association of Insider Voting and Cash Flow Rights with Forecast Accuracy

	Column A			Column B			Column C			Column D			Column E			Column F		
	Coeff. est.	*t* value		Coeff. est.	*t* value		Coeff. est.	*t* value		Coeff. est.	*t* value		Coeff. est.	*t* value		Coeff. est.	*t* value	
WEDGE	-5.00	-3.39	a				-4.13	-3.14	a				-23.16	-3.92	a			
VR				-26.49	-2.45	b				-27.95	-2.18	b				-109.56	-3.47	a
CFR				52.73	3.55	a				41.37	3.38	a				226.63	3.41	a
Firm characteristics controls																		
lnSIZE	7.86	4.70	a	7.65	4.52	a	8.06	4.77	a	8.05	4.70	a	10.55	5.31	a	10.30	5.00	a
LOSS	-59.71	-9.19	a	-59.64	-9.20	a	-57.72	-8.84	a	-57.67	-8.81	a	-63.54	-8.88	a	-63.31	-8.84	a
EPS_VOL	-95.96	-5.39	a	-95.59	-5.50	a	-86.14	-5.19	a	-86.41	-5.23	a	-90.45	-5.46	a	-90.42	-5.38	a
Forecast characteristics controls																		
HORIZON	-0.28	-0.46		-0.25	-0.41		-0.09	-0.15		-0.08	-0.13		-0.02	-0.04		0.00	0.00	
FOLLOW	-0.77	-2.59	a	-0.72	-2.37	b	-0.80	-2.69	a	-0.78	-2.59	b	-0.94	-3.04	a	-0.94	-3.01	a
Governance characteristics controls																		
CEO_CHAIR	-0.31	-0.10		-0.26	-0.09		-2.24	-0.77		-2.24	-0.78		-1.17	-0.41		-1.42	-0.50	
BOARD_SIZE	-0.69	-0.98		-0.69	-0.96		-0.45	-0.64		-0.47	-0.66		-0.16	-0.22		-0.17	-0.24	
IND_DIR	0.74	0.70		12.64	1.08		13.34	1.24		18.88	1.58		10.32	0.97		11.48	1.06	
INST_OWN	7.39	3.68	a	35.82	4.01	a	31.33	3.68	a	34.33	3.91	a	28.24	3.35	a	28.04	3.32	a
Heckman selection control																		
INV_MILL							3.51	0.86		3.72	0.93		-3.85	-0.91		-2.27	-0.55	
Model *F* statistic		10.65	a		10.01	a		10.29	a		9.89	a		13.27	a		12.41	a
Adjusted R^2		0.418			0.420			0.400			0.400			0.415			0.413	
N		2,050			2,050			1,993			1,993			1,915			1,915	

The dependent variable is *ACC*, forecast accuracy, calculated as (|Mean Forecast – Actual EPS| / Stock Price) multiplied by -1; *WEDGE* is the divergence of insider voting rights from cash flow rights computed as the ratio of VR to CFR; *VR* is the proportion of voting rights owned by insiders; *CFR* is the proportion of dividend rights owned by insiders. Other variables are defined in Table 2. All models include an intercept, as well as year and Fama & French (1997) indicator variables. Test statistics and significance levels are calculated based on standard errors (Rogers) clustered at the firm level. Letters a and b denote significance at the 1% and 5% levels, respectively (two-sided tests). Coefficient estimates are multiplied by 1,000 for purposes of exposition.

Table 4

2SLS: First Stage Regression Results

Dependent Variables:	Column A *WEDGE*			Column B *VR*			Column C *CFR*		
	Coeff. est.	*t* value		Coeff. est.	*t* value		Coeff. est.	*t* value	
FAMILY	0.464	6.71	a	0.215	17.15	a	0.065	6.42	a
FIN_NEED	0.180	2.03	b	0.032	2.02	b	-0.010	-0.77	
ΔSALES	-0.142	-1.38		-0.011	-0.06		0.005	0.31	
lnAT	0.150	7.06	a	-0.014	-3.69	a	-0.020	-6.54	a
ROA	0.281	1.43		-0.027	-0.77		-0.032	-1.12	
INTANG	0.624	4.30	a	-0.073	-2.77	a	-0.089	-4.19	a
MEDIA	-0.163	-1.94	a	0.121	7.95	a	0.084	6.89	a
FIRMAGE	0.267	5.81	a	-0.078	-9.31	a	-0.064	-9.48	a
Model *F* statistic		26.52	a		53.33	a		27.07	a
Adjusted R^2		0.094			0.176			0.096	
N		1,962			1,962			1,962	

The dependent variables are *WEDGE*, the divergence of insider voting rights from cash flow rights computed as the ratio of VR to CFR, *VR* the proportion of voting rights owned by insiders, and *CFR* the proportion of dividend rights owned by insiders. *FAMILY* is a dummy variable that takes a value of 1 if the founder and/or the founder's familial descendants own 25% or more of outstanding voting rights; *FIN_NEED* is ROE/(1-ROE); *ΔSALES* is percentage sales growth; *lnAT* is natural log of total assets; *ROA* is income before extraordinary items divided by total assets; *INTANG* is 1 – (Net PPE + Inventories)/Total Assets; *MEDIA* is a dummy variable that takes a value of 1 if a firm is a member of the media industry; *FIRMAGE* is age of the firm. Test statistics and significance levels are calculated based on standard errors (Rogers) clustered at the firm level. Letters a and b denote significance at the 1% and 5% levels, respectively (two-sided tests).

Table 5

The Association of Insider Voting and Cash flow Rights with Forecast Dispersion

	Column A		Column B		Column C		Column D		Column E		Column F	
	Coeff. est.	*t* value	Coeff. est.	*t* value	Coeff. est.	*t* value	Coeff. est.	*t* value	Coeff. est.	*t* value	Coeff. est.	*t* value
WEDGE	1.54	3.54 [a]			1.25	3.58 [a]			7.61	4.29 [a]		
VR			7.28	2.04 [b]			4.12	1.74 [c]			40.14	3.95 [a]
CFR			-15.44	-3.23 [a]			-11.44	-3.25 [a]			-83.35	-4.35 [a]
Firm characteristics controls												
lnSIZE	-2.70	-5.79 [a]	-2.64	-5.60 [a]	-2.81	-5.99 [a]	2.83	-5.97 [a]	-3.57	-6.63 [a]	-3.60	-6.53 [a]
LOSS	15.18	9.47 [a]	15.17	9.46 [a]	14.18	9.33 [a]	14.16	9.30 [a]	14.53	9.22 [a]	14.60	9.20 [a]
EPS_VOL	27.39	4.49 [a]	27.31	4.60 [a]	23.64	4.55 [a]	23.73	4.59 [a]	26.02	5.00 [a]	26.06	5.00 [a]
Forecast characteristics controls												
HORIZON	-0.23	-1.43	-0.24	-1.68 [c]	-0.22	-1.39	-0.22	-1.39	-0.20	-1.18	-0.20	-1.23
FOLLOW	0.24	2.70 [a]	0.22	2.47 [b]	0.25	2.94 [a]	0.25	2.85 [a]	0.28	3.54 [a]	0.29	3.51 [a]
Governance characteristics controls												
CEO_CHAIR	0.27	0.30	0.25	0.28	0.87	1.09	0.87	1.10	0.52	0.69	0.61	0.81
BOARD_SIZE	0.73	3.83 [a]	0.73	3.81 [a]	0.63	3.51 [a]	0.64	3.54 [a]	0.42	2.49 [b]	0.41	2.37 [b]
IND_DIR	-0.42	-0.15	-2.18	-0.70	-2.55	-0.96	-4.49	-1.59	-0.40	-0.16	-0.67	-0.26
INST_OWN	-10.42	-3.91 [a]	-11.51	-4.31 [a]	-9.52	-3.64 [a]	-10.28	-3.95 [a]	-7.88	-3.10 [a]	-7.87	-3.05 [a]
Heckman selection control												
INV_MILL					1.65	1.29	1.54	1.21	3.54	2.97 [a]	3.17	2.66 [a]
Model *F* statistic		84.97 [a]		38.88 [a]		28.94 [a]		29.87 [a]		28.96 [a]		31.30 [a]
Adjusted R^2		0.437		0.438		0.418		0.418		0.443		0.443
N		2,050		2,050		1,993		1,993		1,915		1,915

The dependent variable is *DISP*, the standard deviation of analyst forecasts scaled by stock price; *WEDGE* is the divergence of insider voting rights from cash flow rights computed as the ratio of VR to CFR; *VR* is the proportion of voting rights owned by insiders; *CFR* is the proportion of dividend rights owned by insiders. Other variables are defined in Table 2. All models include an intercept, as well as year and Fama & French (1997) indicator variables. Test statistics and significance levels are calculated based on standard errors (Rogers) clustered at the firm level. Letters a, b, and c denote significance at the 1%, 5%, and 10% levels, respectively (two-sided tests). Coefficient estimates are multiplied by 1,000 for purposes of exposition.

Table 6

The Association of Insider Voting and Cash Flow Rights with Forecast Accuracy: Alternative Forecast Horizons

	Column A			Column B			Column C		
Horizon:	Most recent			6 months			9 months		
	Coeff. est.	*t* value		Coeff. est.	*t* value		Coeff. est.	*t* value	
VR	-24.53	-2.33	b	-18.05	-1.76	c	-26.69	-2.12	b
CFR	54.25	3.52	a	51.31	3.50	a	61.72	3.42	a
Firm characteristics controls									
lnSIZE	11.48	5.75	a	10.30	5.52	a	12.15	5.44	a
LOSS	-54.45	-8.06	a	-60.80	-9.43	a	-67.66	-9.63	a
EPS_VOL	-97.19	-6.20	a	-125.25	-6.56	a	-148.79	-6.22	a
Forecast characteristics controls									
HORIZON	-0.18	-3.52	a	-0.65	-0.89		-1.35	-1.68	c
FOLLOW	-1.37	-3.98	a	-0.95	-3.05	a	-1.00	-3.38	a
Governance characteristics controls									
CEO_CHAIR	1.25	0.35		-0.60	-0.20		-2.50	-0.74	
BOARD_SIZE	-0.65	-0.82		-1.16	-1.67	c	-0.55	-0.70	
IND_DIR	18.15	1.38		9.86	0.88		16.94	1.38	
INST_OWN	26.84	2.75	a	29.86	3.18	a	38.91	3.61	a
Model *F* statistic		8.57	a		21.18	a		17.70	a
Adjusted R^2		0.392			0.486			0.510	
N		2,247			2,063			2,026	

The dependent variable is *ACC*, forecast accuracy, calculated as (|Mean Forecast – Actual EPS| / Stock Price) multiplied by -1; *VR* is the proportion of voting rights owned by insiders; *CFR* is the proportion of dividend rights owned by insiders. Other variables are defined in Table 2. All models include an intercept, as well as year and Fama & French (1997) indicator variables. Test statistics and significance levels are calculated based on standard errors (Rogers) clustered at the firm level. Letters a, b, and c denote significance at the 1%, 5%, and 10% levels, respectively (two-sided tests). Coefficient estimates are multiplied by 1,000 for purposes of exposition.

Table 7

The Association of Insider Voting and Cash Flow Rights with Forecast Dispersion: Alternative Forecast Horizons

	Column A			Column B			Column C		
Horizon:	Most recent			6 months			9 months		
	Coeff. est.	*t* value		Coeff. est.	*t* value		Coeff. est.	*t* value	
VR	6.66	1.61		8.00	2.24	b	7.76	2.00	b
CFR	-15.83	-2.89	a	-14.14	-2.99	a	-14.34	-2.79	a
Firm characteristics controls									
lnSIZE	-3.69	-6.10	a	-2.74	-5.13	a	-3.84	-5.90	a
LOSS	15.96	7.69	a	13.03	8.20	a	12.78	7.20	a
EPS_VOL	32.55	5.76	a	35.60	5.78	a	39.40	5.23	a
Forecast characteristics controls									
HORIZON	0.07	3.28	a	0.06	0.35		0.07	0.32	
FOLLOW	0.41	3.72	a	0.25	3.06	a	0.42	3.87	a
Governance characteristics controls									
CEO_CHAIR	-1.68	-1.43		0.27	0.31		-0.84	-0.86	
BOARD_SIZE	0.96	3.42	a	0.64	3.10	a	0.89	3.62	a
IND_DIR	-1.01	0.67		3.35	1.17		5.04	1.56	
INST_OWN	-8.97	-2.82	a	-10.05	-4.24	a	-10.44	-3.68	a
Model *F* statistic		93.38	a		25.37	a		14.19	a
Adjusted R^2		0.399			0.487			0.456	
N		2,247			2,063			2,026	

The dependent variable is *DISP*, the standard deviation of analyst forecasts scaled by stock price; *VR* is the proportion of voting rights owned by insiders; *CFR* is the proportion of dividend rights owned by insiders. Other variables are defined in Table 2. All models include an intercept, as well as year and Fama & French (1997) indicator variables. Test statistics and significance levels are calculated based on standard errors (Rogers) clustered at the firm level. Letters a and b denote significance at the 1% and, 5% levels, respectively (two-sided tests). Coefficient estimates are multiplied by 1,000 for purposes of exposition.

Table 8

The Association of Changes in Disproportionate Insider Control with Changes in Forecast Accuracy and Dispersion

Dependent Variable	Column A ΔACC			Column B ΔACC			Column C ΔDISP			Column D ΔDISP		
	Coeff. est.	*t* value		Coeff. est.	*t* value		Coeff. est.	*t* value		Coeff. est.	*t* value	
ΔWEDGE	-8.91	-2.12	b				2.01	1.41	d			
$\Delta WEDGE_{t-1}$				-16.78	-3.11	a				2.87	2.08	b
Firm characteristics controls												
ΔlnSIZE	45.35	6.92	a	46.44	6.53	a	-11.52	-6.84	a	-10.31	-7.30	a
ΔLOSS	-46.85	-4.35	a	-40.27	-4.12	a	4.01	2.52	b	3.35	2.50	b
ΔEPS_VOL	30.25	1.44		36.84	1.09		9.46	1.36		8.28	0.99	
Forecast characteristics controls												
ΔHORIZON	-0.04	-0.08		-0.29	-0.48		0.07	0.46		0.07	0.52	
ΔFOLLOW	0.59	0.81		0.60	0.84		-0.31	-1.99	b	-0.29	-1.99	b
Governance characteristics controls												
ΔCEO_CHAIR	4.76	0.72		11.45	1.89	c	0.14	0.11		-0.06	-0.06	
ΔBOARD_SIZE	3.96	2.56	b	4.99	2.58	b	0.25	0.66		0.24	0.66	
ΔIND_DIR	-56.14	-3.14	a	-56.63	-3.00	a	19.66	2.97	a	18.10	3.14	a
ΔINST_OWN	40.22	2.26	b	43.76	2.25	b	-7.59	-2.03	b	-4.97	-1.39	
Model *F* statistic		24.48	a		2.38	a		2.90	a		4.17	a
Adjusted R^2		0.260			0.247			0.220			0.215	
N		1,694			1,406			1,694			1,406	

The dependent variables are Δ*ACC*, the change in forecast accuracy, calculated as (|Mean Forecast – Actual EPS| / Stock Price) multiplied by -1 and Δ*DISP*, the change in the standard deviation of analyst forecasts scaled by stock price; *ΔWEDGE* is the change in the divergence of voting rights from cash flow rights computed as the change in the ratio of VR to CFR; $\Delta WEDGE_{t-1}$ is the prior year's change in *WEDGE*. All other variables are the changes in variables defined in Table 2. All models include an intercept, as well as year and Fama & French (1997) indicator variables. Test statistics and significance levels are calculated based on standard errors (Rogers) clustered at the firm level. Letters a, b, and c denote significance at the 1%, 5%, and 10% levels, respectively (two-sided tests); d denotes significance at the 10% level (one sided-test) for variables for which hypotheses are made. Coefficient estimates are multiplied by 1,000 for purposes of exposition.