Do Career Concerns Affect the Delay of Bad News Disclosure?

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ABSTRACT

Theory argues that career concerns encourage managers to withhold bad news in the hopes that subsequent events will turn in their favor. However, empirical research finds that government regulation (i.e., Regulation Fair Disclosure, or "Reg FD") eliminates this problem, despite any link between Reg FD and career concerns. The purpose of this paper is to examine whether career concerns affect the delay of bad news disclosure. We conduct our analyses in two stages. First, we show that research design problems led prior literature to overstate the extent to which Reg FD reduces the relative delay of bad news disclosure. We find that managers delay the disclosure of bad news after Reg FD, providing implicit evidence that career concerns might be a factor in the decision to delay bad news disclosure. Second, we identify a compensation contract (i.e., ex-ante severance pay agreements) that firms use to explicitly reduce their CEO's career concerns. We find that if managers are promised a sufficiently large payment in the event of dismissal, they no longer delay the disclosure of bad news relative to good news. Overall, our findings provide strong support that managers delay bad news disclosure due to career concerns and suggest a mechanism through which firms can mitigate the delay.

Keywords: Bad news, information withholding, management forecast, career concerns

I. INTRODUCTION

Managers possess superior private information relative to the investment community and they face various incentives that affect when they release that private information (Healy and Palepu 2001; Verrecchia 2001). On the one hand, managers may accelerate the disclosure of bad news (or delay the disclosure of good news) to mitigate personal liability over litigation (Skinner 1994; Kasznik and Lev 1995; Baginski, Hassell, and Kimbrough 2002) or to lower the exercise price on their option grants (Yermack 1997; Aboody and Kasznik 2000). On the other hand, if managers face career concerns, they may withhold bad news and gamble that subsequent corporate events will allow them to "bury" the bad news. Kothari, Shu, and Wysocki (2009) present compelling evidence from management forecasts that, on average, managers delay the release of bad news relative to good news. They also find that this result is eliminated after Regulation Fair Disclosure (Reg FD), and conclude that the elimination is due to Reg FD constraining managers' ability to release private information (rather than career concerns).[1] This conclusion is warranted because Reg FD does not remove career concerns. However, it calls into question what role, if any, career concerns play in the asymmetric release of good and bad news.

The purpose of this study is to examine whether career concerns affect the delay of bad news disclosure. We conduct our analyses in two stages. First, we show that research design problems led prior literature to overstate the extent to which Reg FD induces timely disclosure of bad news. We find that managers still delay the disclosure of bad news after Reg FD. This finding provides implicit evidence that career concerns have the potential to encourage managers to delay bad news disclosure. Second, we identify a compensation contract (i.e., ex-ante severance pay agreements) that firms use to explicitly reduce their CEO's career concerns. We find that if

[1] Kothari et al. (2009) do not provide separate statistical tests on pre-Reg FD and post-Reg FD management forecast samples because they are interested in Reg FD as a control, but they do conclude that Reg FD reduces the bad news delay. However, inspection of their Table 5 suggests that the bad news delay is eliminated post-Reg FD (see Section 2.2 for further detail).

managers are promised a sufficiently large payment in the event of dismissal, they no longer delay the disclosure of bad news relative to good news. Overall, our findings provide strong support that managers delay bad news disclosure due to career concerns and suggest a mechanism through which firms can mitigate the delay.

In the first stage of our analysis, we transact on the findings of two recent methodological studies that examine the effects of Reg FD on manager forecasting behavior to learn whether the career concerns effect is, in fact, weak enough to be substantially offset by a regulatory change that increases the costs of early good news disclosure (which incentivizes its delay and in turn, reduces the asymmetric delay of bad news). These recent studies provide evidence that Reg FD changed the way managers provide forecasts in two fundamental ways. First, Ciconte, Kirk, and Tucker (2014) show that range forecasts increased from 46.9 percent of all forecasts before Reg FD to 80.5 percent of all forecasts after Reg FD. They also show that managers conservatively shift their range forecasts downward, and as a result, the upper bound of the range forecast represents investors' interpretation of managers' expectations better than the midpoint after Reg FD. The implication of this finding is that prior research that relies on the midpoint to interpret the news in management forecast releases is likely to incorrectly classify firms as providing bad news when in fact investors perceive these firms to be providing good news (see Exhibit 1). Such a misclassification will bias tests that infer bad news delay from larger absolute stock price reactions to bad news disclosures relative to good news disclosures by understating the negative reaction to bad news disclosures.

Second, Rogers and Van Buskirk (2013) show that the incidence of management forecasts that are released concurrently with an earnings announcement (i.e., "bundled" with an earnings announcement) increased from 25.3 percent before Reg FD to 70.0 percent after Reg FD. They also provide evidence that it is important to control for the information conveyed by the concurrent earnings announcement when interpreting investor reaction to the management forecast because the information in these two items is correlated (Rogers and Van Buskirk 2013). The implication of this finding is that prior research that fails to control for the news conveyed by the concurrent earnings surprise may not cleanly capture investor reaction to the information conveyed by the management forecast. Therefore, our first research question re-examines the effectiveness of regulation at mitigating the delay of bad news after adjusting our research design to reflect the manner in which managers provide forecasts post Reg FD.

In the second stage of our analysis, we consider whether an action specifically designed to remove career concerns, contracting with managers, is sufficient to reduce the extent to which they delay the disclosure of bad news relative to good news. Firms can provide managers with ex-ante severance contracts that promise payments in the event they are involuntarily dismissed from the firm, and recent studies find that such contracts reduce the career concerns of those managers (Rau and Xu 2013; Cadman, Campbell, and Klasa 2014). If career concerns explain why managers delay the disclosure of bad news, we should find that managers are more forthcoming when ex-ante severance agreements promise particularly large payments in the event of their dismissal.

We examine the stock price reaction to the release of management forecasts, and classify firms as releasing good news when their management forecast of earnings is above the most recent consensus analyst forecast and releasing bad news when their management forecast of earnings is below the most recent consensus analyst forecast. Following Kothari et al. (2009), we expect that if managers accumulate and withhold bad news up to a certain point, but leak and immediately reveal good news to investors, the magnitude of the negative stock price reaction to bad news management forecasts should be greater than the magnitude of the positive stock price reaction to good news management forecasts.[2]

We offer two main results. First, if we use the midpoint of management's range forecasts to classify the news conveyed by management's forecast and extend the number of post-Reg FD sample years through 2007, we find no evidence that managers delay the release of bad news relative to good news in the post-Reg FD time period (consistent with Kothari et al. (2009)'s finding of a significant decrease in the post-Reg FD period using a sample ending in 2002). However, if we use the upper bound of management's range forecasts to classify the news conveyed by management's forecast and control for the earnings surprise related to forecasts that are released concurrently with earnings announcements (i.e., "bundled" forecasts), we find that the *on average* result that managers delay the release of bad news relative to good news persists after Reg FD. These results suggest that research design choices in prior literature may miscalibrate investors' interpretation of managers' expectations, and leave open the possibility that career concerns affect the extent to which managers disclose bad news in a timely manner. Second, we find that when managers are explicitly promised ex-ante severance payments in the top quartile of the sample, there is no longer an asymmetric release of information. This result provides explicit evidence that career concerns affect the timing of bad news disclosure, and suggests that firms can contract with managers to encourage timely disclosure.

Additional analysis shows that our results are not sensitive to using a continuous measure for ex-ante severance payments, the inclusion of industry or firm fixed effects, controlling for the determinants of severance pay, and cannot be explained by the existence of a non-compete

[2] Kothari et al. (2009) also examine a sample of dividend changes. We focus on management forecasts for two reasons. First, dividend changes are less common and are disproportionately upward revisions (i.e., good news). Our severance pay data only covers the dates from 12/15/2006 to 12/31/2007, rendering the number of observations in that year with negative dividend changes insufficient for statistical tests. Second, Beyer, Cohen, Lys, and Walther (2010) explicitly call for more research on the effects of career concerns on management forecasts (page 306). We directly address that call for research with our tests.

agreement with management.[3]

We contribute to the literature examining the link between manager incentives and disclosure choice. In a theoretical paper, Hermalin and Weisbach (2007) link managers' career concerns to disclosure choice, and assume that owners assess the CEO's ability using all available information and replace the CEO if their assessment is too low. They conclude that optimal disclosure lacks transparency, and that this issue is particularly problematic with bad news. Survey evidence supports this notion, as Graham, Harvey, and Rajgopal (2005) find that a subset of CFOs admits that they delay bad news disclosures in the hope that the firm's status improves so that they never have to release the news. Kothari et al. (2009) present empirical evidence that, on average, managers delay the disclosure of bad news relative to good news and that this tendency is eliminated after Reg FD. We extend these findings in two key ways. First, we show that Kothari et al.'s (2009) primary finding of bad news delay is more generalizable than previously thought, as it continues to hold in today's Reg FD environment. In so doing, we highlight the need for future research in this area to use the upper bound of management's range forecasts as a proxy for investors' interpretation of managements' expectations rather than the midpoint (as suggested by Ciconte et al. 2014), as well as to control for the concurrent earnings surprise associated with "bundled" forecasts (as suggested by Rogers and Van Buskirk 2013). Second, we provide strong evidence confirming the prior theoretical, survey, and archival evidence suggesting that the reason managers delay the release of bad news relative to good news is due to their career concerns. We show that if firms contract with managers to provide a

[3] We also explore three alternative explanations for our finding that there is a larger stock price reaction to voluntarily disclosed bad news relative to good news (just as in Kothari et al. 2009). First, managers may disclose bad news promptly, while gradually releasing good news. Second, market participants may believe that the release of bad news is more credible than the release of good news. Finally, the stronger reaction to bad news could be driven by growth firms failing to meet analyst expectations (i.e., the "torpedo effect" from Skinner and Sloan 2002). We discuss the plausibility of these alternative explanations and, where possible, provide tests to disentangle them. Our results support the explanation that managers are withholding bad news, and they are less consistent with any of these alternative explanations.

sufficiently large ex-ante severance agreement, managers no longer withhold bad news relative to good news. Ex-ante severance agreements are a powerful proxy for career concerns because they are compensation contracts that are explicitly written to mitigate such concerns.

We also contribute to the executive compensation literature. Firms have recently come under considerable scrutiny for awarding large amounts of severance pay to CEOs who are dismissed from their positions. For instance, when Merrill Lynch disclosed that its CEO would receive a severance package of $161.5 million immediately after the firm lost a record $8.4 billion, U.S. Banking Committee Chairman Christopher Dodd complained, "Why do you give someone a contract that winds up paying him so handsomely when he fails?" (Vekshin and Katz 2007) These criticisms sparked research into the costs and benefits of severance pay. On the one hand, severance pay could be a manifestation of value-destructive agency problems in firms with entrenched managers who obtain pay beyond the value of their human capital (i.e., Bebchuk and Fried 2003, 2004). On the other hand, severance pay could be a form of efficient contracting that reduces managerial risk-aversion (i.e., career concerns) and increases firm value (i.e., Lambert and Larcker 1985, 1987; Knoeber 1986; Smith and Watts 1992; Ju, Leland, and Senbet 2014).

Several recent studies provide evidence that ex-ante severance contracts are more consistent with efficient contracting than a manifestation of agency problems (Rau and Xu 2013; Cadman et al. 2014; Chen, Cheng, Lo, and Wang 2015). Collectively, these studies find that firms provide ex-ante severance contracts to provide insurance for their managers' human capital, and that managers with these contracts invest in riskier projects that increase firm value. That is, these studies focus on why firms provide these contracts, and how these contracts influence the manager's investment decisions. We extend these findings by examining how these contracts influence the manager's disclosure decisions. Our evidence is consistent with ex-ante severance contracts acting as a form of efficient contracting whereby the manager's career concerns are mitigated and therefore s/he no longer delays the disclosure of bad news relative to good news. Our findings are inconsistent with severance pay as a form of rent extraction. If this were the case, we should find no association between severance pay and the degree to which managers delay the disclosure of bad news relative to good news.[4]

II. THEORY, PRIOR LITERATURE, AND HYPOTHESES

2.1. Prior research on the timing of bad news disclosures

Managers face a number of incentives that affect how forthcoming they are likely to be with bad news disclosures. On one hand, the risk of litigation induces managers to voluntarily disclose bad news prior to the mandatory earnings announcement (Skinner 1994; Baginski et al. 2002; Cao and Narayanamoorthy 2011) because timely disclosure (1) makes it more difficult for plaintiffs to successfully argue that management withheld information and (2) reduces the amount of damages (Skinner 1997). On the other hand, managers' career concerns could induce managers to delay the release of bad news (Kothari et al. 2009). Specifically, prior research shows that investors respond asymmetrically to bad versus good news disclosures, as the magnitude of the negative stock price reaction to bad news is greater than the magnitude of the positive stock price reaction to good news (Hutton et al. 2003; Rogers and Stocken 2005; Kothari et al. 2009; Ng et al. 2013). Kothari et al. (2009) argue that this asymmetry arises because managers gradually reveal good news to investors, but accumulate and withhold bad news until it is clear that subsequent corporate events will not reverse or offset the bad news.[5] If this is the

[4] In a contemporaneous study, Cheng, Cho, and Kim (2014) find that managers with longer vesting periods on their bonus and stock-based compensation are more forthcoming with bad news. Our paper differs from Cheng et al. (2014) in that they focus on stock-option compensation that provides managers with upside gain potential, while we focus on severance pay that provides managers with downside protection. Thus, we examine a compensation contract that explicitly links to career concerns as the reason that managers delay disclosing bad news (Hermalin and Weisbach 2007; Kothari et al. 2009).

[5] Other studies provide alternative explanations for the asymmetric market response to good versus bad news. Most notably, a stream of research argues that bad news is inherently more credible than good news and, as a result, the market reacts more

case, then when managers disclose good news through their management forecast it will likely have previously been disclosed to investors and, thus, garner a weaker market reaction compared to when they disclose bad news through their management forecasts. Consistent with this argument, the absolute market reaction to bad news is larger than the absolute market reaction to good news. Thus, *on average*, evidence suggests that managers delay bad news. It is important to note, however, that the asymmetry is weaker for firms that face high litigation risk, which is consistent with litigation risk inducing more timely disclosure (Skinner 1994; Baginski et al. 2002; Cao and Narayanamoorthy 2011).

Although Kothari et al. (2009) argue for career concerns as the reason that managers delay the release of bad news, there is little direct evidence. In their review of the recent literature, Beyer et al. (2010, 306) note "our understanding of how management's career concerns affect their disclosure strategies is still limited, a fact previously noted in the survey by Healy and Palepu (2001)." Nevertheless, some evidence supports career concerns as the motivation to delay bad news. In cross-sectional tests, Kothari et al. (2009) proxy for career concerns using distress risk – a measure of the companies' prospects rather than the executives' prospects – and find some, albeit weak, evidence of greater asymmetric disclosure timing when distress risk is high. A contemporaneous study by Ali et al. (2015) proxies for career concerns using (1) CEO tenure and (2) state indices for the enforceability of noncompetition agreements. Their results using these proxies also suggest that career concerns are associated with greater asymmetry. We rigorously consider the strength of these proxies in Sections 4.3.3 and 4.3.4.

strongly to the more credible bad news (Jennings 1987; Hutton et al. 2003; Ng et al. 2013). However, this research fails to find evidence that bad news forecasts are less biased than good news forecasts and in fact, some evidence shows that bad news forecasts are *more* biased than good news forecasts (Rogers and Stocken 2005). Furthermore, recent research warns against proliferating the "stylized inference" that bad news is inherently more credible than good news (Merkley, Bamber, and Christensen 2013). Nevertheless, similar to Kothari et al. (2009), in Section B.2 in Appendix B, we provide multiple empirical analyses to reduce the likelihood that our results are due to differences in the way investors assess the credibility of good and bad news disclosures. Lastly, while explanations other than the delay of bad news exist for the asymmetric market reaction, evidence from other methodologies and settings also supports the delay of bad news (e.g., Ge and Lennox 2011).

2.2. Prior research on management forecast changes pre- to post- Reg FD

Kothari et al. (2009) also examine the impact of regulation on the timeliness of bad news disclosures. They find evidence consistent with Reg FD eliminating the delay of bad news relative to good news. Specifically, in their Table 5, holding information asymmetry, litigation risk, and distress constant and setting RegFD equal to zero (i.e. pre-Reg FD) yields a total reaction to good news management forecasts of 0.029 and a total reaction to bad news forecasts of -0.106. In absolute value, the bad news reaction is clearly much larger supporting the conclusion of bad news delay. Post-Reg FD (i.e., RegFD = 1) yields total price reactions to good and bad news of 0.041 and -0.012, respectively. The absolute price reaction to good news is *larger* than the price reaction to bad news, supporting the conclusion that Reg FD eliminates the bad news delay effect and (if anything) encourages managers to delay the release of *good* news.[6]

Kothari et al. (2009) interpret this result to mean that managers are less able to asymmetrically release information after Reg FD because the regulation makes it more difficult to leak private information to investors. Consistent with this interpretation, King et al. (1990) argue that public disclosure of good news is more costly than private disclosure of good news to analysts. Thus, in the post-Reg FD period when private disclosure is disallowed, timely disclosure of good news is more costly.[7] As managers become less likely to gradually reveal good news, the asymmetry in the timeliness of good and bad news disclosures disappears.

Although the explanation is intuitive, Reg FD does not directly address career concerns, it merely disallows selective disclosure without immediate broad dissemination. The strong effect

[6] In fact, if one does not control for information asymmetry, litigation risk, and distress, but instead reevaluates the effect for conditions in which the delay of bad news is expected to be the greatest – information asymmetry is high (= 1), litigation risk is low (= 0), and distress is high (= 1) – the total price reactions to good and bad news post-Reg FD are 0.069 and -0.080, respectively. Although the absolute bad news price reaction is higher in this optimal setting for bad news delay, it is only marginally higher. Given the standard errors in the table implied by the coefficients and t-statistics, the difference is likely insignificantly different from zero (and importantly, this marginally and likely insignificant difference occurs under the set of conditions for which the bad news delay should be *largest*).

[7] If managers are already delaying the release of bad news, Reg FD should only encourage the delay of good news as well.

of Reg FD on disclosure timing suggests that managers are more concerned about the costs associated with broad public dissemination of good news (an unintended regulatory effect) than they are with their career concerns, which, in turn, casts doubt on the relative strength of the career concerns effect on disclosure timing. The cost of public disclosure of good news imposed by Reg FD having such a dampening effect on relatively early good news disclosure is a curious finding given that managers with career concerns would not want to delay good news disclosure. Further, inconsistent with Kothari et al. (2009), some evidence suggests that the delay of bad news relative to good news is *greater* after Reg FD (Roychowdhury and Sletten 2012). Roychowdhury and Sletten (2012) document an asymmetric market reaction to bad news versus good news *earnings announcements* that is consistent with managers leaking news prior to good news earnings announcements but withholding news prior to bad news earnings announcements. In contrast to the evidence in Kothari et al. (2009) on reactions to *management forecasts*, Roychowdhury and Sletten (2012) find that the asymmetry is stronger post-Reg FD (i.e., that managers' delay of bad news increases after Reg FD).

Given the conflicting findings and the fact that Reg FD has no clear link to career concerns, it is important to examine other, potentially confounding, changes around Reg FD. Specifically, two changes in forecasting occurred that could have implications for researchers when they measure the news conveyed by management forecasts. First, after Reg FD and the corporate scandals of Enron and WorldCom that occurred during the same time period, managers are far more likely to issue forecasts that cover a broad range of possible earnings per share outcomes rather than a single point estimate. Range forecasts increased from 46.9 percent to 80.5 percent of quarterly EPS forecasts in the First Call CIG database from the 1996-2001 period to the 2002-2010 period (Ciconte et al. 2014). Ciconte et al. (2014) argue that this change in forecasting behavior also affected how investors (and researchers) should interpret managers' true expectation within that range. They argue that, because Reg FD makes private communication illegal, managers must use public channels like management forecasts if they wish to walk analysts down to a beatable forecast estimate. Therefore, after Reg FD managers are more likely to issue range forecasts with an intentional pessimistic bias, while their true expectation of earnings is towards the top of the range or even above the range.

Consistent with these arguments, the evidence in Ciconte et al. (2014) suggests that between 1996 and 2001, the midpoint is the best proxy for managers' expectations; however, between 2002 and 2010, the upper bound (i.e., the highest number provided in a range forecast) is the best proxy for managers' expectations. Specifically, in the earlier time period, the distribution of actual earnings realizations is centered slightly above the *midpoint* of range management forecasts: actual earnings are below the midpoint for 38.5% of forecasts, at it for 11.6%, and above it for 49.9%. In the post-Reg FD period, the distribution of actual earnings is centered slightly above the *upper bound* of range management forecasts: actual earnings are below the upper bound for 38.4% of forecasts, at it for 15.9%, and above it for 45.7% of forecasts.[8]

Although actual earnings are distributed around the upper bound of range forecasts in more recent years, Ciconte et al. (2014) find that analysts do not unravel the pessimistic bias, which results in beatable analyst expectations. Importantly, however, they find that investors *do* unravel the bias and react to management forecasts as though the upper bound is management's true expectation. Thus, Ciconte et al. (2014) argue that the upper bound of range forecasts best represents investors' interpretation of managers' expectations.

[8] In addition, Ciconte et al. (2014) create a continuous measure ranging from -1 to 1 that is calibrated to equal -1, 0, and 1 when actual earnings fall at the lower bound, midpoint, and upper bound, respectively. In the earlier time period, 1996-2001, the mean value is -0.37 and the median is 0.00, while in the later time period, 2002-2010, the mean value is 1.20 and the median is 1.00. Therefore, the evidence is consistent with interpreting the midpoint as managers' true expectation of earnings in the earlier period, and the upper bound as managers' true expectation of earnings in the later period.

A second change in management forecasts since the passage of Reg FD is the increased popularity of management forecasts that are released together with an earnings announcement (i.e., "bundled" forecasts). Bundled forecasts increased from 25.3 percent to 70.0 percent of the forecasts in the First Call CIG database pre- to post-Reg FD (Rogers and Van Buskirk 2013). The market reaction around these bundled forecasts is confounded by the market reaction to the earnings surprise in the concurrent earnings announcement; therefore, we account for the increased prevalence of bundled forecasts and the potential effects on the empirical design by controlling for the earnings surprise from any concurrent earnings announcement.[9]

In sum, differences in the asymmetric timeliness of disclosure across the pre- and post–Reg FD periods cannot be unequivocally attributed to the effects of regulation. We consider important concurrent changes in management forecasting, namely the increased use of both pessimistic range forecasts and bundled forecasts, and re-examine whether the impact of Reg FD offsets the impact of career concerns on disclosure timing.

2.3. Prior research on ex-ante severance pay contracts

If career concerns motivate the asymmetric release of good versus bad news disclosures, then an action specifically designed to remove career concerns, contracting with managers, should mitigate the asymmetry. Several recent academic studies investigate whether compensation contracts that provide CEOs with large payments in the event they are involuntarily dismissed (i.e., ex-ante severance pay contracts) are value-creating or value-destructive contracts from the shareholder's perspective.

[9] Rogers and Van Buskirk (2013) also propose a refined measure of management forecast news for bundled forecasts. Specifically, they provide a methodology for calculating a conditional expectation to which the management forecast is compared in determining forecast news. The conditional expectation is the pre-earnings announcement analyst forecast plus the predicted analyst revision (i.e., the expectation of how analysts will respond to the current earnings announcement), where the predicted revision is a function of the current earnings surprise and the coefficient relating analyst revisions to current earnings surprises for firms that did not issue bundled forecasts. As discussed in footnote 11, our results are unchanged if we use this approach.

Two recent studies examine the consequences of providing managers with a higher *amount* of contracted severance pay. Cadman et al. (2014) find that the contracted severance payment amount is positively associated with CEO risk-taking and the extent to which a CEO invests in projects that have a positive net-present-value. In addition, Brown (2014) finds that firms that provide their managers with higher severance pay exhibit a lower frequency of earnings management. Together, these studies suggest that severance pay contracts are not a form of value-destructive rent extraction from shareholders, but instead represent a form of efficient contracting that reduces the CEO's career concerns and encourages value-enhancing risk taking.

Finally, Cadman et al. (2014) note that, beginning with the 2006 fiscal year, the ExecuComp database reports information on the ex-ante contracted amount of severance pay a dismissed CEO would receive. However, they identify a number of coding errors for this variable, and use a hand-collected sample to show that the correlation between what ExecuComp reports for contracted severance pay and the true value for this pay is 0.76. In light of these findings, we follow the hand-collection procedures detailed in Cadman et al. (2014) to quantify the amount of ex-ante severance pay a CEO is promised in the event s/he is involuntarily terminated.[10]

2.4. Effect of Reg FD on bad news delay after adjusting for changes in disclosure behavior

Kothari et al.'s (2009) method depends on the measurement of the sign of management forecast news and isolation of the price reaction to the management forecast. Consistent with management forecast research prior to that point, Kothari et al. use the midpoint of a range

[10] Section II.B of Cadman et al. (2014) fully describes this hand-collection process. To summarize, we collect from firms' disclosures (DEF-14, DEF-14A, or Item 11 of Form 10-K) the incremental amounts the CEO would receive if he is either terminated "without cause" or resigns for "good reason." The term "without cause" is defined in the contract, and "cause" generally includes conditions such as willful misconduct, breach of fiduciary duties, and fraud. However, "cause" does not typically include CEO incompetence or poor firm performance. The term "good reason" is also defined in the contract, and is generally interpreted to be the case where the firm demotes a CEO by substantially reducing his responsibilities and/or compensation rather than firing him, and as a result of the firm's actions, the CEO resigns for good reason. Because our variable for severance pay is designed to capture the ex-ante, incremental separation pay owed to dismissed CEOs, we remove amounts from the separation payment total that are "vested," or already earned by the CEO regardless of whether the CEO remains with the firm. We make this adjustment because these components of severance pay should not affect a CEO's career concerns.

management forecast to measure the sign of the management forecast news. However, Ciconte et al. (2014) show that investors respond to the news conveyed by management range forecasts using the upper bound of the range (i.e., the highest number in the range) rather than the midpoint. Therefore, using the midpoint to determine the news conveyed by forecasts can incorrectly classify firms as providing bad news when in fact investors perceive the firms to be providing good news. Such a misclassification will bias tests that compare stock price reactions to bad and good news disclosures by understating the reaction to bad news, making it appear that managers are providing bad news in a more timely fashion than they really are. This issue is particularly important after Reg FD because there was a substantial push towards managers issuing pessimistic range forecasts (Ciconte et al. 2014). [11]

Furthermore, prior literature shows that when management forecasts are issued concurrently with earnings announcements, it is important to incorporate the news conveyed by the concurrent earnings announcement if one is interested in examining investor reactions to the news conveyed by the management forecast (Rogers and Van Buskirk 2013). Given that the earnings surprise in a concurrent earnings announcement and management's forecast revision are likely to be correlated, it would be impossible to separate investors' response to future earnings or current earnings without controlling for the current earnings surprise.

In sum, prior research that finds no evidence of asymmetric disclosure timeliness after Reg FD but fails to incorporate important changes in the way managers forecast in the post-FD time period. We first re-examine whether the effect of Reg FD was sufficiently strong to offset the effect of career concerns after adjusting our methodology. Our first hypothesis follows:

[11] The potential for this bias to have an effect on inference is implied by the findings of Roychowdhury and Sletten (2012), who document an asymmetric market reaction to bad news versus good news *earnings announcements* and find that the asymmetry is *stronger* in the post-Reg FD period. By examining earnings announcements and by measuring news based on the sign of market returns, they are not exposed to the measurement issue identified by Ciconte et al. (2014).

H1: Using the upper bound of management range forecasts as the market expectation and controlling for the earnings surprise in a concurrent earnings announcement, disclosure of bad news is delayed (or less timely) relative to good news after Reg FD.

2.5. Effect of severance pay contracts on the delay of bad news disclosures

Kothari et al. (2009) argue that career concerns motivate managers to delay the disclosure of bad news. Prior research finds that managers with high levels of ex-ante severance pay take more risk and thus behave as if their career concerns are lessened (Rau and Xu 2013; Cadman et al. 2014). Therefore, if career concerns motivate managers to delay bad news, we predict that large severance payments will mitigate that delay. Our second hypothesis follows:

H2: Bad news delay is decreasing in the amount of ex-ante severance pay the CEO would receive upon dismissal.

III. RESEARCH DESIGN

To investigate the delay of bad news relative to good news, we follow Kothari et al. (2009) and use the following three models:

$$CAR = \alpha + \beta_0 BadNews + \varepsilon \tag{1}$$

$$CAR = \alpha + \beta_0 BadNews + \beta_1 MFNews + \beta_2 MFNews \times BadNews + \varepsilon \tag{2}$$

$$\begin{aligned} CAR = \alpha &+ \beta_0 BadNews + \beta_1 MFNews + \beta_2 MFNews \times BadNews + Additional\ Controls \\ &+ Additional\ Controls \times BadNews + \varepsilon \end{aligned} \tag{3}$$

where *CAR* equals the five-day cumulative abnormal return (relative to the value-weighted market return) centered around the management forecast announcement date; *MFNews* equals the management earnings forecast minus the most recent analyst forecast consensus, scaled by the analyst forecast consensus; and *BadNews* is an indicator variable that equals 1 if *MFNews* is less than zero, and zero otherwise. To re-examine the effect of Reg FD (i.e., to test H1), we first use the midpoint of range forecasts as investors' interpretation of managers' expectations and do not control for any earnings surprise (as in Kothari et al. 2009). Then, we adjust for management forecast changes by using the upper bound as investors' interpretation of managers' expectations for range forecasts (Ciconte et al. 2014) and controlling for the earnings surprise from any concurrent earnings announcements (Rogers and Van Buskirk 2013).[12]

The *Additional Controls* are control variables identified by prior literature that capture opportunities and incentives that managers face that could encourage them to withhold bad news. Specifically, we control for the firm's litigation risk (Kasznik and Lev 1995; Skinner 1994, 1997; Baginski et al. 2002), information asymmetry between insiders and outside investors (Frankel, McNichols, and Wilson 1995; Lang and Lundholm 2000; Aboody and Kasznik 2000), distress risk (Weisbach 1988; Gilson 1989; DeAngelo 1998; Kothari et al. 2009), and insider and managerial ownership levels (Kothari et al. 2009). Each of these control variables is interacted with *BadNews* to capture any potential nonlinearity. We define these *Additional Controls* fully in Appendix A. Finally, following Kothari et al. (2009), in all the models in our paper, we control for outliers by truncating at the top and bottom 1 percent of *MFNews* and use standard errors that control for heteroscedasticity.

The intercept in each of the above equations, α, captures the market reaction to good news forecasts and the coefficient on *BadNews*, β_0, captures the *incremental* market reaction to bad news management forecasts. Therefore, we compare the magnitude of the total market reaction to good news (i.e., $|\alpha|$) to the magnitude of the total market reaction to bad news (i.e., $|\alpha + \beta_0|$), and test whether the magnitude of the reaction to bad news is statistically larger than the reaction to good news using an *F*-test. Following Kothari et al. (2009), we interpret a significant

[12] Because we are interested in replicating Kothari et al. (2009), we incorporate Rogers and Van Buskirk (2013) by controlling for the earnings surprise in the concurrent earnings announcement whenever the forecast is "bundled." Kothari et al. (2009) scale the news in management forecasts by the analyst forecast consensus, and exclude from the sample any management forecasts with news less than 1% of the most recent analyst forecast consensus. On the other hand, Rogers and Van Buskirk (2013) scale the news in management forecasts by stock price, and examine good (bad) news forecasts larger (less) than 0.01% (-0.01%) of this stock price. For these reasons, we cannot replicate Kothari et al. (2009) and explicitly follow the classification scheme from Rogers and Van Buskirk (2013). Instead, we control for the earnings surprise in the spirit of Rogers and Van Buskirk (2013). However, if we follow Rogers and Van Buskirk (2013) explicitly (i.e., scale the news in management forecasts by stock price and remove forecasts with an absolute magnitude of news less than 0.01% of stock prices), our inferences are unchanged.

difference to mean that managers delay the release of bad news while gradually releasing good news, leading to the stronger reaction for bad news. If H1 holds, we should find that $|\alpha + \beta_0| > |\alpha|$ after adjusting for the methodological implications of increased pessimistic range forecasts and increased bundled forecasts in the post-Reg FD period.

In Equation (1), a significantly greater absolute market response to bad news relative to good news could be due to a greater amount of *total* news released in bad news forecasts and/or a greater amount of information *per unit* of bad news. While both explanations are consistent with managers withholding and delaying bad news disclosures, we include a second model that controls for the amount of news. In Equation (2), the coefficient on $MFNews \times BadNews$, β_2, represents the strength of the market response per unit of bad news. Finally, in Equation (3), we examine the asymmetric market response to bad versus good news forecasts after controlling for the amount of news as well as managers' opportunities and incentives to delay bad news.

3.2. Research design for H2

H2 predicts that the asymmetry in the market's reaction to good versus bad news is decreasing in the amount of ex-ante severance payment the CEO would receive upon dismissal. Thus, we extend Models (1), (2), and (3) by including main and interaction effects of severance:

$$CAR = \alpha + \beta_0 BadNews + \beta_1 SevPay + \beta_2 SevPay \times BadNews + \varepsilon \quad (4)$$

$$CAR = \alpha + \beta_0 BadNews + \beta_1 SevPay + \beta_2 SevPay \times BadNews + \beta_3 MFNews + \beta_4 MFNews \times BadNews + \varepsilon \quad (5)$$

$$CAR = \alpha + \beta_0 BadNews + \beta_1 SevPay + \beta_2 SevPay \times BadNews + \beta_3 MFNews + \beta_4 MFNews \times BadNews + \beta_5 MFNews \times SevPay + \beta_6 MFNews \times BadNews \times SevPay + \varepsilon \quad (6)$$

$$CAR = \alpha + \beta_0 BadNews + \beta_1 SevPay + \beta_2 SevPay \times BadNews + \beta_3 MFNews + \beta_4 MFNews \times BadNews + \beta_5 MFNews \times SevPay + \beta_6 MFNews \times BadNews \times SevPay + Additional\ Controls + Additional\ Controls \times BadNews + \varepsilon \quad (7)$$

where *SevPay* is an indicator variable that equals to 1 if the ratio of CEO contracted severance pay to CEO cash compensation is in the top quartile and 0 otherwise.[13] To test H2, we compare the absolute magnitude of the market reaction to bad news forecasts made by high severance pay CEOs (i.e., $|\alpha + \beta_0 + \beta_1 + \beta_2|$) to the magnitude of the market reaction to good news forecasts made by high severance pay CEOs (i.e., $|\alpha + \beta_1|$), and perform an *F*-test to examine whether they are statistically different. We also calculate the difference between the absolute magnitude of the market reaction to bad news forecasts made by CEOs without high severance pay (i.e., $|\alpha + \beta_0|$) and the magnitude of the market reaction to good news forecasts made by CEOs without high severance pay (i.e., $|\alpha|$). If H2 holds, we should find that the differential market reaction for high-severance pay CEOs is smaller than the differential market reaction for low-severance pay CEOs. Furthermore, an insignificant *F*-test for the high-severance CEOs would indicate that severance payment contracts fully mitigate the delay of bad news relative to good news.

IV. RESULTS

We begin with a replication of Kothari et al. (2009). We follow their sample selection process that spans from 1995 to 2002, and Table 1 Panel A presents descriptive statistics for key variables. Kothari et al. (2009) only examine quarterly management earnings forecasts, but we extend our sample to annual forecasts because, as shown in the panel headings in Table 1, the incidence of quarterly vs. annual management forecasts has changed substantially from pre-Reg FD (far more quarterly forecasts) to post-Reg FD (far more annual forecasts). The prevalence of annual forecasts is more pronounced when the sample is extended to include more post-Reg FD years than in Kothari et al. (2009).

Panel A of Table 2 presents average management forecast news and price reactions to that news for various subsamples. Panel B of Table 2 provides an exact replication of Table 4 from

[13] In Section 4.3.1, we extend the calculation of *SevPay* to include various definitions of CEO contracted severance pay. Our results are robust to all of these variations.

Kothari et al. (2009), where we use quarterly management forecasts, the midpoint of range forecasts for investors' interpretation of the news, and do not control for the news associated with any simultaneously released earnings. Our number of observations ties closely to Kothari et al. (2009) (4,112 versus 4,016, respectively), and our results are strikingly similar. For example, the results for Model (2) show that the market reaction to good news management forecasts is a positive 1.3 percent (i.e., α = 0.013), whereas the market reaction to bad news management forecasts is a negative 6.3 percent (i.e., $\alpha + \beta_0 = 0.013 - 0.076 = -0.063$). The 6.3 percent negative reaction to bad news matches exactly with the economic significance in column (2) of Table 4 in Kothari et al. (2009) ($\alpha + \beta_0 = -0.063$).

As a reminder, the interpretation of these results is that if managers delay the disclosure of bad news, the market will be more surprised and the reaction will be larger to bad news disclosures relative to good news disclosures. An *F*-test for the difference in these two market reactions is statistically significant at *p*-value less than 0.001. Thus, using Kothari et al.'s (2009) sample period and specifications, Table 2 shows that we are able to exactly replicate the result that, *on average*, managers appear to delay the disclosure of bad news relative to good news.

4.1. Effects of regulation (Reg FD) on the asymmetric release of good and bad news

Kothari et al. (2009) find that after Reg FD the asymmetric reaction to good and bad news is significantly reduced, and they conclude that regulation leveled the playing field and reduced the extent to which managers delay the release of bad news. However, as previously noted, prior research suggests that after Reg FD, managers tend to provide more range forecasts and to bundle guidance on future earnings with a concurrent earnings announcement (Rogers and Van Buskirk 2013; Ciconte et al. 2014). Table 1, Panels B, C, and D support these claims. Specifically, Panel D shows that, after Reg FD, 79.7% (88.3%) of quarterly (annual) management forecasts are range forecasts.[14] These percentages are significantly higher than those for pre-Reg FD time periods (i.e., 60.3% and 50.4%, respectively), as shown in Panel B. In other words, the evidence supports prior findings that range forecasts are more prevalent after Reg FD, so research design choices that interpret the news conveyed by range forecasts are more important during this time period.

Additionally, Panel D shows that, after Reg FD, 3.5% (69.4%) of quarterly (annual) management forecasts are "bundled" with earnings releases. These percentages compare to pre-Reg FD time period percentages of 4.3% (28.6%), as shown in Panel B.[15] In other words, the evidence supports prior findings that "bundled" forecasts are more prevalent after Reg FD (particularly with annual forecasts), so research design choices that interpret the news conveyed by bundled forecasts are more important during this time period.

With these descriptive results as motivation, Tables 3 and 4 assess how Kothari et al.'s (2009) results hold after Reg FD when making research design adjustments to account for these differences. Table 3 shows results using quarterly management forecasts, while Table 4 provides results using annual management forecasts. Because the inferences are similar between the two, for brevity we only discuss the results in Table 4.

Panel A of Table 4 compares the results using Kothari et al.'s (2009) research design before and after Reg FD (i.e., using the midpoint of range forecasts as investors' interpretation of managers' expectations and failing to control for the earnings surprise in "bundled" forecasts). Specifically, in the first column, we find that prior to Reg FD the reaction to bad news is significantly greater than the reaction to good news ($|0.012–0.053| > |0.012|$; p-value < 0.001).

[14] Panel D in Table 1 uses the post-Reg FD time period of 2001 to 2007. We elect to end the sample in 2007 in this table because that is the last year for which we have severance contract data for our tests of H2. Panel C shows that the inferences are unchanged if we only examine the post-Reg FD years in Kothari et al.'s (2009) sample (i.e., 2001 and 2002).
[15] Following Kothari et al. (2009), we require each management forecast to have an analyst forecast announced within a month (per I/B/E/S detail file) prior to the management forecast date. As a result, the quarterly guidance samples in Table 1 exhibit a lower portion of quarterly earnings guidance bundled with concurrent earnings announcements.

However, in the second column, after Regulation FD, there is no longer a statistically different reaction between good and bad news (*p*-value = 0.99). One interpretation of this result is that the regulatory changes from Reg FD succeeded in leveling the playing field, and managers no longer delay the release of bad news.

In Panel B of Table 4, however, we alter the research design so that investors' expectation of the news in management forecasts is the upper bound of the range rather than the midpoint, and add a control for the earnings surprise from any simultaneously released earnings news. As can be seen in the second column, after Reg FD we continue to find that investor reaction to bad news is statistically greater than to good news (|0.012–0.033| > 0.012; *p*-value < 0.001). It is important to note that the adjusted R^2 in column 2 of Panel B (15.68%) is not statistically lower than the adjusted R^2 in column 2 of Panel A of (14.33%), suggesting that altering the research design to adjust for these changes in managerial forecasting behavior did not induce noise into the empirical tests.[16]

Finally, Figure 1 provides a graph of the coefficients on good and bad news over time using a Fama-MacBeth regression. Panel A provides results using Kothari et al. (2009)'s research design, while Panel B uses the upper bound for investors' interpretation of managers' expectations and controls for any earnings surprise. As presented in Panel A, the magnitude of the reaction to bad news ("Abs(BadNews)") is significantly larger than the magnitude of the reaction to good news ("GoodNews") up until 2001 when Reg FD took effect. After 2001, these magnitudes are largely indistinguishable, suggesting that Reg FD may have curtailed managers from delaying the release of bad news. However, Panel B shows that when using the upper bound as investors' interpretation of managers' expectations and adjusting for "bundled" forecasts, the magnitude of the reaction to bad news ("Abs(BadNews)") continues to be about

[16] The adjusted R^2 does not decline for either of the two adjustments considered separately (results not tabulated).

two times the magnitude of the reaction to good news ("GoodNews") during the years after Reg FD. In other words, after adjusting the research design for important differences in how managers provide forecasts after Reg FD, we find that *on average* managers continue to delay the release of bad news.

Taken together, the results in Figure 1 and Tables 1 through 4 lead to two primary conclusions. First, without adjustment for post-Reg FD changes in management forecasting behavior (i.e., how managers use range forecasts to convey beliefs and time their forecasts to accompany earnings announcements), the post-Reg FD results in Kothari et al. (2009) hold in extended samples of both quarterly and annual management forecasts. Second, after adjusting for changes in the way managers provide their forecasts after Reg FD, Kothari et al.'s (2009) more general (i.e., pre-Reg FD) finding of bad news delay persists after Reg FD. Thus, we find clear evidence for our first hypothesis (H1), which predicts that even after Reg FD, *on average*, managers continue to delay the release of bad news relative to good news. This finding provides implicit evidence that career concerns may still be a determining factor in managers' decision to delay bad news disclosure.

4.2. Effect of contracting on the asymmetric release of good and bad news

After providing implicit evidence that career concerns motivate managers to asymmetrically time disclosures, we next turn to the explicit effect that contracting on career concerns might have on managers' delay of bad news disclosures (H2). Specifically, firms can use compensation contracts known as ex-ante severance pay agreements to explicitly reduce their CEOs' career concerns. If career concerns motivate managers to withhold bad news relative to good news, these contracts should alleviate this phenomenon.

4.2.1. Sample selection and descriptive statistics for H2

In order to examine the effects of these contracts, we hand-collect data on severance pay. The severance data is only available annually in 2006 and 2007, and consequently this constraint leads to a different sample than was used in our tests of H1. Table 5 Panel A presents the sample selection process for H2. Our initial sample consists of all 10,138 annual management earnings forecasts of firms in the intersection of First Call, CRSP, Compustat, and I/B/E/S for fiscal years ended between December 15, 2006 and December 31, 2007. We use annual rather than quarterly earnings forecasts because the compensation data is only reported on an annual basis.[17] We select this time period because it covers the first full year over which firms were required to quantify ex-ante severance pay amounts.[18]

We eliminate observations that are open-ended or qualitative forecasts, that are missing necessary severance pay information, or that lack an analyst EPS forecast within 30 trading days prior to the management forecast. Additionally, we exclude observations with analyst consensus forecast less than five cents, management forecast news less than one percent of the analyst consensus forecast, and the top and bottom one percent of management forecast news (as in Kothari et al. 2009). The final sample consists of 2,906 management forecasts.

Table 5 Panel B presents an industry breakdown of our H2 sample into two-digit SIC industry sectors. Manufacturing firms issue the largest percentage of sample forecasts (42.12%), followed by telecommunication, transportation, and utilities (13.97%), services (13.83%), and retailing (13.97%). The industry breakdown of sample forecasts is in line with the industry breakdown of Compustat firms, with the exception of retailing (13.28%), which makes up 5% of Compustat firms.[19]

[17] In untabulated results, our inferences are unchanged if we use quarterly management forecasts for tests of H2.
[18] SEC Release No. 33-8732a, which required expanded disclosure of ex-ante severance pay, became effective beginning with fiscal years that ended on December 15, 2006 (SEC 2006).
[19] To eliminate any concerns regarding industry composition, we re-estimate our multivariate regression analyses for H2 after including industry fixed effects, and the results are unchanged. See Section 4.3.2.

Table 6 reports descriptive statistics of the sample used to test H2. When we calculate management forecast news using the upper bound of range forecasts, median and mean *MFNews* are positive (i.e., 0.025 and 0.035 respectively), and the majority of forecasts are good news (i.e., median *BadNews* = 0, mean *BadNews* = 0.271). In contrast, in untabulated analysis, we calculate management forecast news using the midpoint of range forecasts, and the median forecast news, *MFNews*, is negative (i.e., -0.011) and the average forecast news is slightly positive (i.e., 0.005). Further, the majority of forecasts are classified as bad news forecasts (i.e., median *BadNews* = 1, mean *BadNews* = 0.528). Thus, if the true expectation is the upper bound (as shown in Ciconte et al. (2014)), use of the midpoint results in a bias towards bad news.

Panel B of Table 6 presents a correlation matrix. The correlations indicate that severance pay (in both amount, *SevAmt*, and ratio to cash compensation, *SevAmt / CashComp*) is positively correlated with the issuance of bad news management forecasts (*BadNews*). In addition, the ratio of severance payment to cash compensation (*SevAmt / CashComp*) is positively correlated with the cumulative abnormal returns around management forecasts (*CAR*). This univariate evidence is consistent with H2.[20,21]

4.2.2. Univariate and multivariate tests of H2

H2 predicts that managers with sufficiently large severance contracts will have fewer career concerns and, thus, should be less likely to delay the release of bad news. In the context of our model, we should find a smaller differential reaction to the release of bad news relative to good news for the sub-sample with high severance pay. We test this assertion in Table 7, which presents results from estimating Equations (4), (5), and (6) using the upper bound to classify

[20] We note that none of the correlations among variables that are simultaneously included in our regressions give us concerns with respect to multi-collinearity. In addition, variance inflation factors in all of our regressions are less than 5, so multi-collinearity is not a concern (Kutner, Nachtsheim, and Neter 2004).
[21] We present Pearson correlations. As previously mentioned, we control for the effects of outliers by truncating our variables at the top and bottom 1 percent (Kothari et al. 2009). In untabulated results, we examine Spearman correlations. None of the signs flip between Pearson and Spearman correlations, providing further evidence that our results are not driven by outliers.

firms as releasing good news or bad news in our post-Reg FD time period.[22]

The univariate analysis in Panel A of Table 7 shows that the amount of news released in both good news and bad news forecasts is greater for low severance pay firms relative to high severance pay firms, but the differences are only statistically significant at the median. In addition, the market reaction to bad news forecasts is weaker for high severance pay firms (mean = -0.021, median = -0.013) compared to low severance pay firms (mean = -0.044, median = -0.033). Thus, the univariate evidence in Panel A of Table 7 is consistent with severance pay mitigating the delay of bad news (i.e., H2).

Panel B of Table 7 provides a multivariate test of H2 using Equations (4) through (6). We find that when CEOs' ex-ante severance payments are in the top quartile of the sample (i.e., nine times the CEOs' annual cash compensation), there is no longer a differential market reaction to bad news disclosures relative to good news disclosures. For example, when estimating Model (6) we find that the market reaction to good news for CEOs with high severance pay is a positive 1.2 percent (i.e., $\alpha + \beta_1 = 0.012 - 0.000 = 0.012$) and the market reaction to bad news for CEOs with high severance pay is a negative 1.7 percent (i.e., $\alpha + \beta_0 + \beta_1 + \beta_2 = 0.012 - 0.047 - 0.000 + 0.018 = -0.017$). The difference in magnitude between these two is not statistically significant (p-value = 0.41). Furthermore, Model (6) in Panel B of Table 7 shows that the market reaction to good news for CEOs *without* high severance pay is a positive 1.2 percent (i.e., $\alpha = 0.012$), and the market reaction to bad news for CEOs *without* high severance pay is a negative 3.5 percent (i.e., $\alpha + \beta_0 = 0.012 - 0.047 = -0.035$). This difference is significant (p-value < 0.001).

[22] In untabulated results, we estimate Equations (4), (5), and (6) using the midpoint of range forecasts (rather than the upper bound) to calculate the news conveyed by management forecasts. Consistent with the upper bound results, we find that the coefficients on *SevPay×BadNews* are all positive when using the midpoint, again suggesting that high severance pay induces more timely disclosure of bad news. However, the significance levels are weaker. As previously discussed, when using the midpoint to calculate news, we cannot detect the asymmetric market response in our post Reg FD sample period and therefore cannot use these estimations to draw conclusions about the ability of severance pay to mitigate or eliminate the extent to which managers delay the release of bad news.

In Table 8, we present results of H2 estimating Equation (7) which includes the *Additional Controls* that capture opportunities and incentives that could encourage managers to withhold bad news. As can be seen, our results continue to hold throughout. Specifically, the *F*-tests in columns (1) through (4) suggest that the asymmetric market response to bad versus good news holds for low severance pay firms (p-value < 0.001), but does not hold for high severance pay firms (p-value ≥ 0.38). We also find several results similar to Kothari et al. (2009) with respect to the *Additional Controls* variables. Specifically, litigation risk appears to only marginally mitigate managers' propensity to withhold bad news (0.009, p-value < 0.20). Further, firms with greater information asymmetry appear to have a greater ability to delay bad news (–0.029, p-value < 0.01), firms with greater financial distress appear to marginally have a greater tendency to delay bad news (–0.062, p-value < 0.10), and firms with greater inside ownership appear to have a greater tendency to delay bad news (–0.14, p-value < 0.05). Overall, the results in Tables 6 through 8 are consistent with H2 and suggest that when managers receive a sufficiently large ex-ante severance pay contract, they no longer delay the release of bad news disclosures.

4.3. Additional analysis

4.3.1. Sensitivity of results to various definitions of ex-ante severance pay

In our previous tests, our measure for severance pay (*SevPay*) was an indicator variable equal to one if the CEO is in the top quartile of the sample for severance pay, and zero otherwise. These results suggest that as long as a CEO is promised severance payments that are sufficiently large, career concerns will be reduced to the point that the CEO no longer delays the disclosure of bad news. In untabulated results, we separately redefine the *SevPay* indicator variable to be equal to one if the CEO is in the top quintile, tercile, or median, respectively, and zero otherwise. In all cases, our results are qualitatively similar, continuing to show that when managers receive severance payments, they no longer delay the release of bad news.

We also replace the *SevPay* indicator variable with a continuous measure of ex-ante contracted severance pay. In untabulated results, we continue to find that the presence of severance pay reduces the extent to which managers delay the release of bad news. However, we no longer find that this asymmetric release of information completely goes away. Taken together, these results suggest that managers no longer delay the release of bad news only when they are promised sufficiently large amounts of severance pay.

4.3.2. Sensitivity to the inclusion of industry and firm fixed effects

Throughout our study, we remain consistent with the research design from prior literature and do not include fixed effects.[23] However, to control for the extent that severance payments are correlated within industry, we include industry fixed effects based on two-digit Standard Industrial Classification (SIC) code industry classifications in untabulated tests of H2. We continue to find that the bad news reaction is greater than the good news reaction when severance pay is low (by 3.2 percent, *p*-value = 0.04), while the difference in the reaction to bad news vs. good news is statistically indistinguishable when severance pay is high (i.e., *p*-value = 0.17). This result suggests that differences among industries are not driving our results.

Because we have a (brief) time-series of severance pay data for some firms in our sample, we also include firm fixed effects in our tests of H2. We continue to find that the bad news reaction is greater than the good news reaction when severance pay is low (by 4.9 percent, *p*-value = 0.01), while the difference in the reaction to bad news and good news is statistically indistinguishable when severance pay is high (i.e., *p*-value = 0.46). This suggests that our H2 results are not sensitive to time-invariant correlated omitted variables related to severance pay.

[23] This is particularly important in our tests of H1 because we are carefully comparing our results to the prior literature by only adjusting the reference point of the news in management forecasts (i.e., comparing the midpoint to the upper bound; incorporating the earnings surprise in an accompanying earnings announcement).

4.3.3. Sensitivity to the inclusion of the determinants of ex-ante severance pay

Cadman et al. (2014) find that firms provide larger levels of severance pay to CEOs that: (1) have a non-competition clause (i.e., they cannot work for a competitor of the firm for a specified number of years upon dismissal), (2) are hired from outside of the firm, (3) are early in their tenure as CEO, and (4) have low levels of equity ownership in the firm. To distinguish whether our results are due to the formal act of contracting, or simply due to characteristics that are associated with having a severance contract, in Table 8 columns (5) and (6) we control for these determinants of severance pay contracts. We continue to find that the bad news reaction is greater than the good news reaction when severance pay is low (by 2.1 and 2.0 percent, respectively; p-value = 0.01), while the difference in the reaction to bad news and good news is not statistically different when severance pay is high (i.e., p-value ≥ 0.88). This result suggests that our H2 results are due to the act of writing a contract with the CEO rather than simply reflecting the fact that CEOs with certain characteristics are more likely to have a contract.

4.3.4 Sensitivity to non-competition agreements

As previously discussed, a contemporaneous working paper by Ali et al. (2015) examines the extent to which managers delay bad news disclosures by using non-competition agreements as a proxy for career concerns. They predict that if managers have an enforceable non-compete agreement then they would have greater concern about their careers if they are fired because they could not work for a competitor for a certain number of years (usually two or three). They do not have firm-level data on whether a manager actually has a non-compete agreement; rather, they assume that if a firm is headquartered in a state where non-competition agreements are more enforceable, then the CEO has one. In contrast, we have hand-collected data for our managers and can precisely identify whether the CEO has a non-compete agreement. In this section, we use that data to further consider which contract is driving the change in the timing of bad news disclosure (i.e., non-competition agreements or severance agreements) or if it is both.

It is not clear that non-compete agreements would be a powerful proxy for career concerns for at least three reasons. First, these agreements can be legally challenged by managers and this would obviously reduce their power as a proxy for career concerns. Absent a compelling reason such as protection of trade secrets or key contracts, courts do not consider it a fair labor practice to reduce competition in the labor markets. Garmaise (2011) notes that labor laws vary by state and non-compete agreements are enforced in a relatively small number of states. Second, because of the ability to challenge their legality, non-compete agreements typically only cover a small number of the firm's direct rivals at its level in the supply chain (MacElree 2015). This again reduces the effectiveness by which these agreements would increase a CEO's career concerns. Finally, upper management's skills can be transferred across industries, and prior research shows that firms typically hire CEOs with prior CEO experience (Cadman et al. 2014).

Nevertheless, we use our hand-collected data on non-compete agreements to more fully examine what role (if any) they play in the delay of bad news disclosure. We begin by attempting to replicate the result from Ali et al. (2015) showing that non-compete agreements increase career concerns and lead to an increase in the delay of bad news disclosure. We first alter our sample to include firms included by Ali et al. but excluded in our sample (i.e., including non-Execucomp / non-S&P 1500 firms). Column 1 in Panel A of Table 9 shows that, consistent with Ali et al., if a firm is headquartered in a state that has the most extensive non-compete enforceability (i.e., index values 7 through 9 from Garmaise 2011) it is marginally more likely to delay the disclosure of bad news (*t*-statistic = 1.79). However, column 2 in Panel A of Table 9 shows that if we remove non-Execucomp firms, this marginal result is no longer significant.

In Panel B of Table 9, we examine whether non-compete agreements have an effect on our sample after we interact our hand-collected non-compete data that shows whether the executive actually has such a contract with the enforceability of these contracts. Throughout Panel B, we continually find no evidence that non-compete agreements affect the extent to which managers delay bad news disclosure. Importantly, when we include severance pay in these regressions, we continue to find that severance pay eliminates the asymmetric disclosure of news. Overall, severance pay is a stronger proxy for career concerns, and our evidence shows that the severance pay contract eliminates the delay of bad news disclosure while non-competition contracts do not.

4.4. Discussion on causality

Our tests are designed exactly as those in Kothari et al. (2009). Consequently, we examine the short-window market reaction to the release of information, and therefore, our tests are a natural experiment that examines how investors react to good and bad news disclosures. However, our cross-sectional tests related to severance pay do not definitively imply that managers will no longer delay the release of bad news if firms provide them with severance pay.[24] It is possible that some other factor leads firms to provide CEOs with sufficiently large severance payments, and this factor also leads CEOs to be more forthcoming with bad news disclosure. That said, our tests are strongly motivated by theory (Hermalin and Weisbach 2007) which predicts that a reduction in career concerns will reduce the extent to which managers delay the release of bad news, and ex-ante severance pay contracts are explicitly designed to reduce managers' career concerns (Rau and Xu 2013; Cadman et al. 2014). Furthermore, as discussed in Section 4.3.2, our results are not sensitive to the inclusion of firm fixed effects, suggesting that time-invariant correlated omitted variables are not driving our results. If anything, the difference in the market reaction to bad news relative to good news is stronger after including firm fixed

[24] Kothari et al. (2009) face this same limitation in their cross-sectional tests related to career concerns and litigation risk.

effects. Furthermore, our results are not sensitive to the inclusion of the determinants of severance pay, suggesting that the results are due to the act of writing the severance contract with the CEO rather than certain CEO characteristics that are associated with having a contract. Finally, regardless of causality, our tests suggest that ex-ante severance pay contracts are associated with more timely disclosure of bad news and, thus, serve as a form of efficient contracting for the firm. If severance pay contracts were a form of rent extraction, we would have found no relation between severance pay and the disclosure of bad news.

4.5. Alternative explanations

Kothari et al. (2009) note several alternative explanations to the interpretation of our results as indicating that managers delay the release of bad news, and that this delay can be mitigated if the manager receives a compensation contract that sufficiently reduces his/her career concerns. In Appendix B, we discuss the plausibility of these alternative explanations and, where possible, provide tests to disentangle them. Our results support the explanation that managers are withholding bad news, and are less consistent with any alternative explanations.

V. CONCLUSION

We re-examine the extent to which managers delay the disclosure of bad news and offer two key insights. First, we show that failing to alter the research design to reflect the fact that (1) investors expect management's range forecasts to come in at the upper bound rather than the midpoint and (2) some forecasts are accompanied by earnings announcements – understates the extent to which management delays bad news relative to good news in post-Reg FD periods. Making these design modifications provides implicit evidence that the career concerns may have an effect on managers' decision to delay bad news disclosure. Second, we identify a compensation contract that firms can use to explicitly reduce their CEO's career concerns – an ex-ante severance agreement. We find that if managers are promised a sufficiently large payment in the event of dismissal, they no longer delay the disclosure of bad news relative to good news. Overall, our findings provide strong support that managers delay bad news disclosure due to career concerns and suggest a mechanism through which firms can mitigate the delay.

Future research that investigates the extent to which managers delay the disclosure of bad news should use the upper bound of managers' range forecasts to determine the news conveyed by the forecast and should control for earnings surprises that accompany any "bundled" forecasts. In addition, future studies may wish to examine the capital market costs associated with managers' decision to delay the disclosure of bad news, and the extent to which firms can contract with severance payments to reduce these costs. For example, studies could examine whether the delay of bad news leads to higher costs of debt and equity capital, less efficient investment, and whether contracting influences firms' financing and investing decisions through the capital market outcomes associated with bad news delay.

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APPENDIX A

The definition of variables used in empirical tests

Variables	Description
CAR	Cumulative abnormal returns above value-weighted market returns for 5 trading days around the management earnings forecast (MEF) announcement date (i.e., [-2, +2]); if a MEF was issued after 4PM, the next day is regarded as the announcement date.
MFNews	News in MEF relative to the most recent analyst forecast consensus issued within 30 days prior to the MEF, scaled by the analyst forecast consensus. Following Kothari et al. (2009), the analyst forecast consensus should be larger than 5 cents, and the news should be larger than 1% of the analyst forecast consensus.
BadNews	Coded 1, if *MFNews* is smaller than zero; otherwise coded 0
SevAmt	Contracted severance payment amount to the firm's CEO as reported in the 10-K report in the fiscal year
CashComp	Sum of salary and bonus paid to the CEO in the fiscal year as reported in the ExecuComp database
SevPay	Coded 1, if the ratio of the contracted severance payment amount (i.e., *SevAmt*) to the CEO's cash compensation (i.e., *CashComp*) is in the top quartile (i.e., 9 times); otherwise coded 0
LitRisk	Coded 1, if the firm's estimated litigation risk, based on the Rogers and Stocken (2005) model, is above the median of the sample; otherwise coded 0
HiAsym	Coded 1, if the firm is above the sample median value of a single information asymmetry factor; otherwise coded 0 (Kothari et al. 2009). The information asymmetry factor is derived from a principal component analysis based on the information asymmetry proxies: market-to-book ratio, stock volatility, high-tech firms, financial leverage, and regulatory status (Kothari et al. 2009).
DistressRisk	Coded 1, if the firm's Z-score (Zmijewski 1984) financial distress rank is in the top quartile of the sample; otherwise coded 0
HiInsideOwn	Coded 1, if the share-ownership of the firm's insiders is above the median value of the sample in the fiscal year; otherwise coded 0
OutsideCEO	Coded 1, if the CEO was appointed from outside; if internally promoted, coded 0
CEOown	The stock ownership by the CEO as of the fiscal year end
CEOtenure	The CEO's tenure with the firm as of the fiscal year end
Noncompete	Coded 1, if the CEO signed a non-compete agreement with the firm; otherwise coded 0

APPENDIX B

Alternative explanations tested in KSW (2009)

The purpose of this appendix is to exhaustively test our results for the same alternative explanations that were tested by Kothari, Shu, and Wysocki (KSW) (2009). Because our results are not sensitive to these explanations and because the analysis is largely identical to that in KSW (2009), we present these tests in an appendix rather than the text. As a reminder, we interpret our results to indicate that managers delay the release of bad news, and that this delay can be mitigated if the manager receives a compensation contract that sufficiently reduces his/her career concerns. In this section, we consider several alternative explanations for the statistical associations that we document.

B.1. Managers disclose bad news promptly, but good news gradually

An alternative explanation for our finding that the market reaction to bad news is stronger than the market reaction to good news is that managers disclose bad news promptly and release good news gradually. Consequently, investors would react more to bad news. However, in a multi-period, efficient markets setting, it is not clear why investors would not recognize that good news disclosures are consistently under-disclosed (i.e., the tip of the iceberg) and thus would not react to the discounted present value of the cash flow consequences of both the disclosed and inferred good news (Kothari et al. 2009). In such a situation, one would expect that, on average, the market reaction to good news is equal to that of bad news.

Despite this conceptual issue with the "promptly-disclose-bad-news" explanation, we nevertheless present three empirical tests designed to examine whether our results are likely to be explained by managers disclosing bad news promptly while gradually releasing good news. First, we examine the market reaction over the three months leading up to the management forecast after separating firms into good and bad news buckets. Figure 2 shows the return patterns of good news and bad news firms over the three-month window leading up to the management forecast. Panel A shows the mean cumulative abnormal return, and Panel B shows the fraction of the total news revealed over time. As can be seen, the returns appear to be incorporated throughout the three-month window for good news firms, while the returns for bad news firms are initially close to zero, and then drop sharply toward the end, with more than 80 percent of the news revealed in the last two weeks. Collectively, the evidence in Figure 2 is consistent with our hypothesis that managers delay the release of bad news relative to good news, and is inconsistent with the "promptly-disclose-bad-news" explanation. In untabulated results, we performed the same analysis from Figure 2 on the subset of firms with managers that are promised high severance pay (i.e., *SevPay* = 1), and consistent with our tests of H2 we find that these patterns no longer hold.

Second, if good news forecasts only reveal a fraction of the total news, we would expect to find that forecast errors are larger (i.e., the difference between ex-post actual earnings and management's forecast) relative to bad news forecasts. For our sample of 2,906 management forecasts, the mean absolute forecast error for good and bad news is 0.84% and 1.04%, respectively, and the differences are statistically insignificant (i.e., p-value = 0.30). These results are similar in partitions of both low and high severance pay. Once again, the data is inconsistent with the explanation that good news is released gradually, but bad news is released promptly and completely.

Finally, high litigation risk firms have the strongest incentives to release bad news early. Thus, if firms promptly release bad news and gradually release good news, we should find the effect to be most pronounced for these firms. However, as reported in Table 7, we find the opposite. Specifically, we find that the asymmetry is *less* pronounced (not more pronounced) for high litigation risk firms, which is again inconsistent with the alternative explanation. Collectively, the conceptual explanations and empirical results in this section provide reasonable assurance that our tests capture the extent to which managers delay the release of bad news disclosures, and are not instead a result of managers promptly releasing bad news and gradually releasing good news.

B.2. Differential credibility of good and bad news management forecasts

As previously discussed, a potential alternative explanation is that bad news is more credible than good news (Jennings 1987; Hutton et al. 2003; Ng, Tuna, and Verdi 2013). If good news management forecasts are less credible due to over-optimism, we should find that signed forecast errors (i.e., ex post actual earnings less management forecast) are more negative (i.e., biased) for good news forecasts relative to bad news forecasts. Contrary to this prediction, for our sample we find that the mean forecast error for good news and bad news is -0.38 percent and -0.37 percent, respectively, and the difference between these two is statistically insignificant (i.e., *p*-value = 0.90). These results are similar in partitions of both low and high severance pay. In addition, these results are similar to those reported in other samples (Rogers and Stocken 2005; Kothari et al. 2009), and are inconsistent with forecasts of good news being less credible than forecasts of bad news.

Second, if investors perceive good news to be less credible than bad news, it must be due to greater uncertainty in the good news forecast relative to the bad news forecast. This greater uncertainty is a form of idiosyncratic risk. Because investors can diversify away idiosyncratic risk, the pricing implications of this differential credibility should be minimal (Kothari et al. 2009) and should not explain our findings. Furthermore, if there is a difference in uncertainty between good and bad news forecasts, we should find that good news forecast errors have a higher standard deviation relative to bad news forecast errors. In our sample, the standard deviation of signed (absolute) forecast errors for good news forecasts is 1.46% (0.35%), while the standard deviation of signed (absolute) forecast errors for bad news forecasts is 1.61% (0.36%). These standard deviations are not statistically significantly different between the two groups (*p*-values of 0.30 and 0.21, respectively). These results are similar in partitions of both low and high severance pay. In other words, we find no evidence that good news forecasts are in fact more uncertain than bad news forecasts, which is inconsistent with the credibility explanation.

Finally, if bad news forecasts are more credible than good news forecasts, we should not only find that the market reacts more strongly to bad news in short windows around the forecast announcement date (as we do in Tables 5 and 6), but we should also find this same pattern over the three month window leading up to the announcement date. Figure 2 shows that the three month cumulative price reactions are similar in magnitude for good and bad news forecasts. This result is inconsistent with bad news forecasts being more credible than good news forecasts. Collectively, the conceptual explanations and empirical results in this section provide reasonable assurance that our tests capture the extent to which managers delay the release of bad news disclosures, and are not instead a result of differences in the credibility of good and bad news.

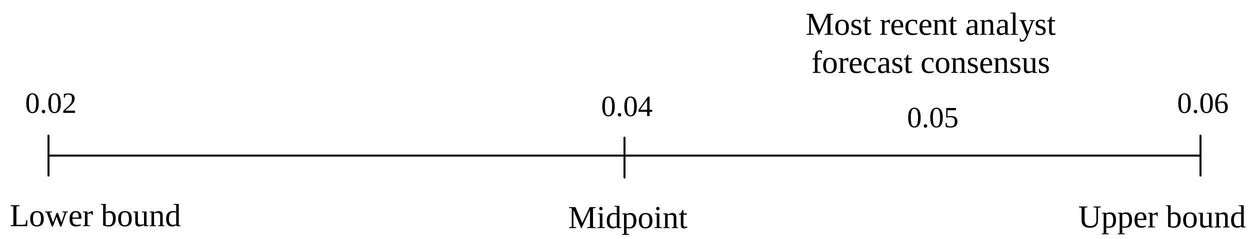
B.3. Torpedo effect

Prior research documents that the market reacts more strongly when growth firms fail to meet analyst expectations (i.e., the "torpedo effect" documented in Skinner and Sloan 2002). To ensure our results are not capturing differences in firm growth, in untabulated results we control for firm growth opportunities in all of our regressions using firms' market-to-book assets ratio. Our inferences are unchanged, suggesting that our results are not limited to the subsample of firms with high growth opportunities.

EXHIBIT 1

Example of misclassification of management forecast news

The most recent analyst forecast consensus of earnings per share is \$0.05, and then management forecasts earnings per share to be between \$0.02 and \$0.06.



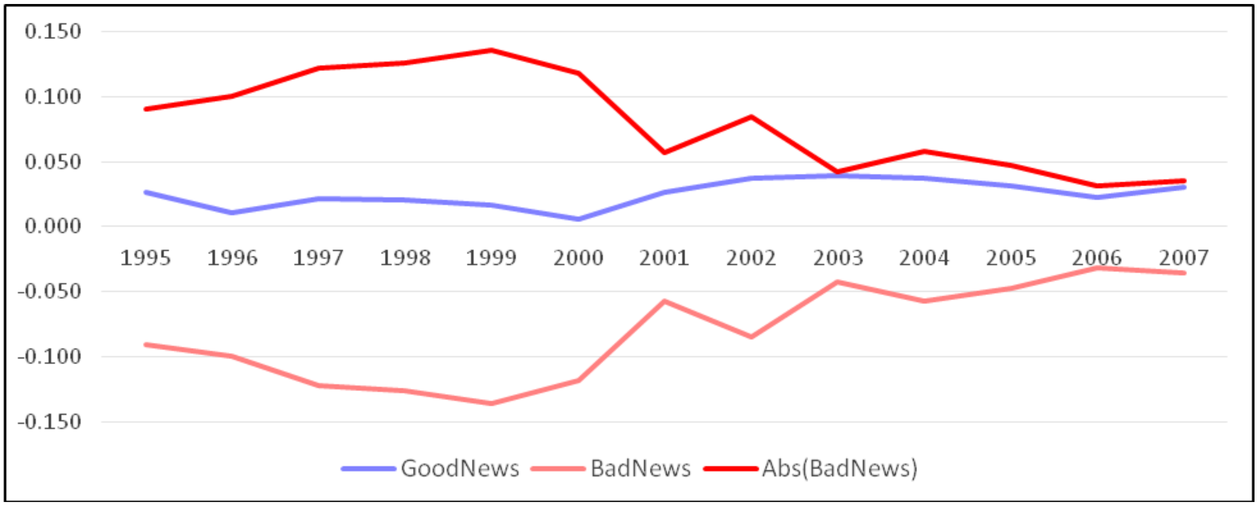
Management range forecast

	A specific example	**General conditions**
Criteria for misclassification	0.04 < 0.05 ≤ 0.06	MF midpoint < AF consensus ≤ MF upper bound
Classification of news using midpoint	0.04 < 0.05	Bad news
Investors' true interpretation of news using upper bound	0.06 > 0.05	Good news

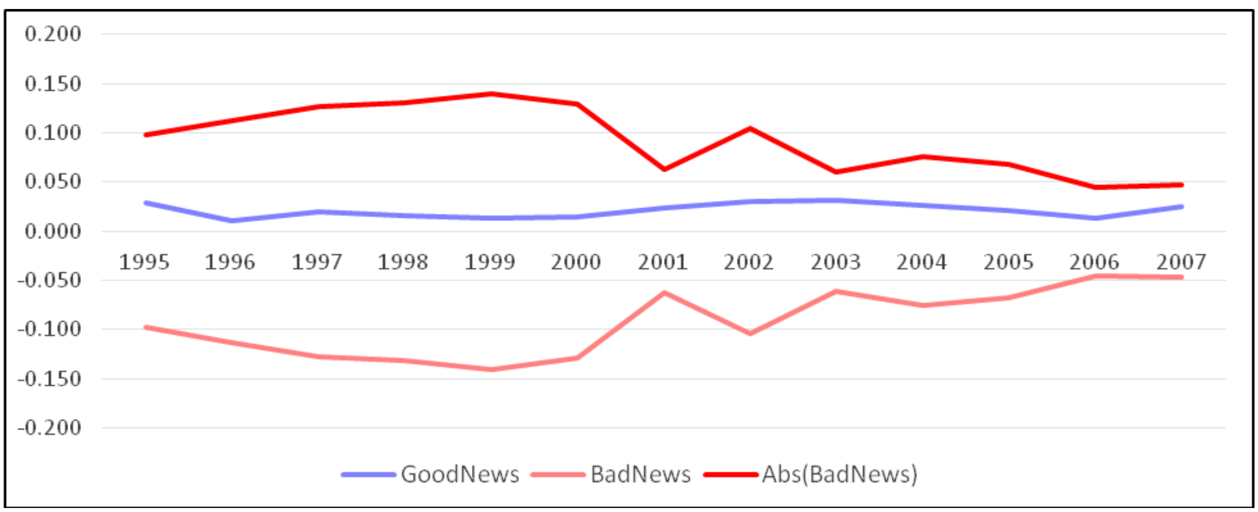
FIGURE 1

Graph of Fama-MacBeth regression coefficients for good vs. bad news over time

Panel A: Regression coefficients using the mid-point as investors' interpretation of managers' expectations and failing to control for bundled forecasts (as in Kothari, Shu, Wysocki 2009)



Panel B: Regression coefficients using the upper-bound as investors' interpretation of managers' expectations (based on Ciconte et al. 2014) and adjusting for bundled forecasts (based on Rogers and Van Buskirk 2013)

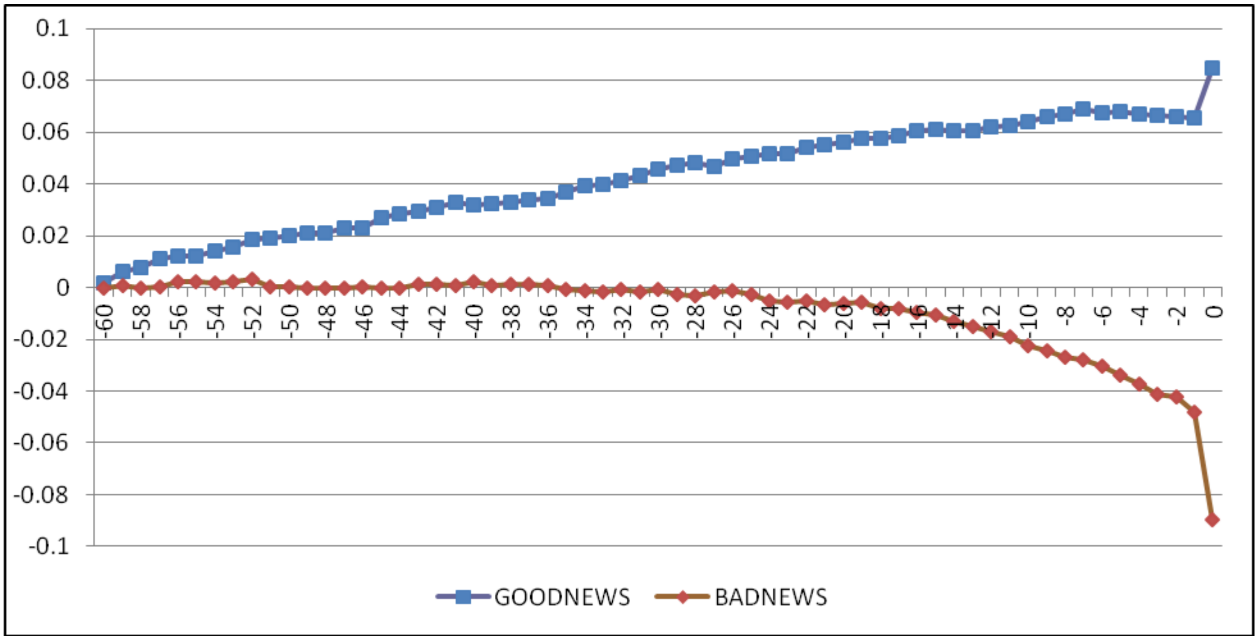


These figures present the coefficients of a Fama-MacBeth regression model – *CAR* = *GoodNews* + *BadNews* (+ *EarnSurp*) – for the period from 1995 to 2007. The sample consists of the 6,817 management forecasts of quarterly earnings (i.e., 2,200 management forecasts between 1995 and 2000; 4,617 management forecasts between 2001 and 2007) that satisfy the sample selection criteria as in Kothari et al. (2009). *News* in Panel A (in Panel B) is defined as the difference between the mid-point (upper-bound) of management earnings forecasts (MEF) and the most recent analyst forecast consensus issued within 30 trading days prior to the MEF, scaled by the analyst forecast consensus. *GoodNews* (*BadNews*) is coded 1 where *News* > 0 (*News* < 0). Panel B additionally controls for the earnings surprise when the management forecast is bundled with an earnings announcement (i.e., *EarnSurp*) (Rogers and Van Buskirk 2013).

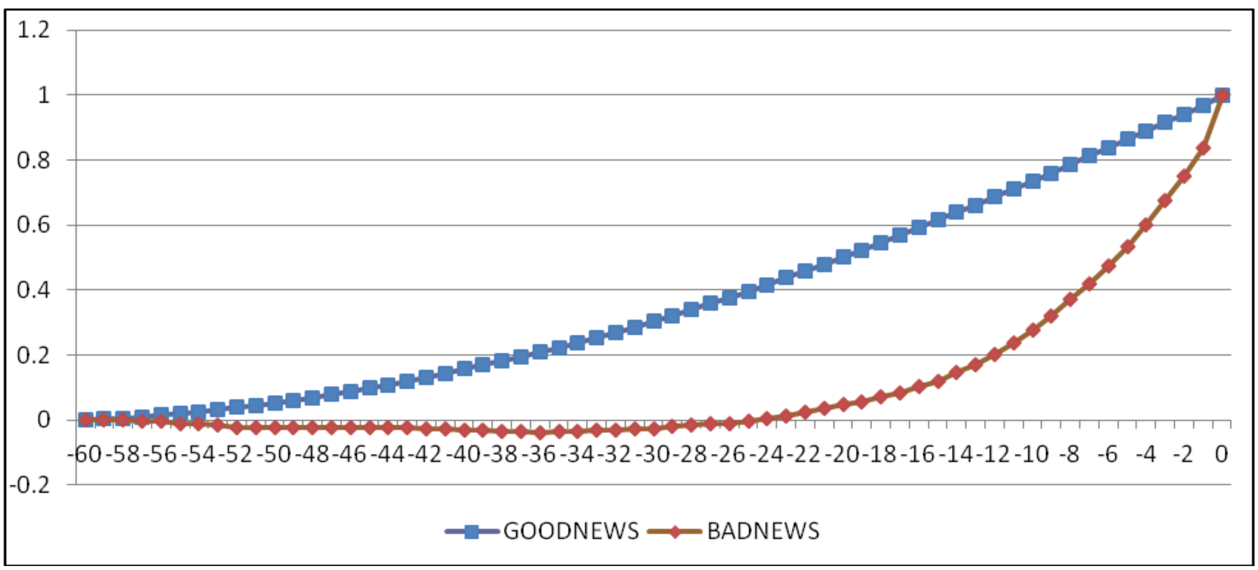
FIGURE 2

Graphs of stock return timing around management forecast releases

Panel A: Mean cumulative market-adjusted return



Panel B: Percentage of news released



These figures present cumulative stock returns prior to the good and bad news management earnings forecasts. The sample consists of the 1,912 post-Reg FD management forecasts of quarterly earnings per share in the sample replication of Kothari et al. (2009) in Table 2. *News* is defined as the difference between the upper bound of management earnings forecasts (Ciconte et al. 2014) and the most recent analyst forecast consensus issued within 30 trading days prior to the MEF, scaled by the analyst forecast consensus. The good news sample consists of all observations where *News* > 0, after adjusting for the news in any bundled forecasts as in Rogers and Van Buskirk (2013). The bad news sample consists of all observations where News < 0, after adjusting for the news in any bundled forecasts as in Rogers and Van Buskirk (2013). Panel A presents the mean cumulative market-adjusted returns for the good vs. bad news samples prior to the management forecast date (day 0). Panel B presents the cumulative news up to day *t* scaled by the total news over the entire period.

TABLE 1

Descriptive statistics for H1

Panel A: Descriptive statistics – Full time period in KSW (2009), 1995 to 2002

Variables	Quarterly earnings forecasts (*N* = 4,112)					Annual earnings forecasts (*N* = 3,403)				
	Mean	Std dev	P25	Median	P75	Mean	Std dev	P25	Median	P75
CAR	-5.87%	15.44%	-13.05%	-2.92%	3.18%	-1.70%	12.37%	-6.33%	-0.16%	5.05%
MFNews	-0.193	0.518	-0.340	-0.083	0.038	-0.011	0.207	-0.059	-0.011	0.039
BadNews	0.666	0.472	0.000	1.000	1.000	0.520	0.499	0.000	1.000	1.000
Range_indicator	0.667	0.471	0.000	1.000	1.000	0.688	0.463	0.000	1.000	1.000
Bundled_indicator	0.035	0.184	0.000	0.000	0.000	0.481	0.499	0.000	0.000	1.000

Panel B: Descriptive statistics – Pre Regulation FD in KSW (2009), 1995 to 2000

Variables	Quarterly earnings forecasts (*N* = 2,200)					Annual earnings forecasts (*N* = 1,301)				
	Mean	Std dev	P25	Median	P75	Mean	Std dev	P25	Median	P75
CAR	-8.05%	15.94%	-16.05%	-5.03%	1.94%	-3.08%	13.17%	-7.58%	-0.90%	4.07%
MFNews	-0.233	0.540	-0.389	-0.113	0.028	-0.019	0.230	-0.069	-0.017	0.032
BadNews	0.701	0.458	0.000	1.000	1.000	0.580	0.494	0.000	1.000	1.000
Range_indicator	0.603	0.489	0.000	1.000	1.000	0.504	0.500	0.000	1.000	1.000
Bundled_indicator	0.043	0.204	0.000	0.000	0.000	0.286	0.452	0.000	0.000	1.000

Panel C: Descriptive statistics – Post Regulation FD in KSW (2009), 2001 and 2002

Variables	Quarterly earnings forecasts (*N* = 1,912)					Annual earnings forecasts (*N* = 2,102)				
	Mean	Std dev	P25	Median	P75	Mean	Std dev	P25	Median	P75
CAR	-3.35%***	14.45%	-9.11%	-1.11%***	4.69%	-0.84%***	11.78%	-5.45%	0.49%***	5.71%
MFNews	-0.147***	0.489	-0.275	-0.050***	0.051	-0.005*	0.191	-0.053	0.011***	0.043
BadNews	0.627***	0.484	0.000	1.000***	1.000	0.482***	0.500	0.000	0.000***	1.000
Range_indicator	0.740***	0.439	0.000	1.000***	1.000	0.802***	0.398	1.000	1.000***	1.000
Bundled_indicator	0.025***	0.158	0.000	0.000***	0.000	0.601***	0.489	0.000	1.000***	0.000

Panel D: Descriptive statistics – Expanded Post Regulation FD Period, 2001 to 2007

Variables	Quarterly earnings forecasts (*N* = 4,617)					Annual earnings forecasts (*N* = 9,800)				
	Mean	Std dev	P25	Median	P75	Mean	Std dev	P25	Median	P75
CAR	-1.86%***	10.83%	-6.21%	-0.73%***	3.90%	0.09%***	8.63%	-3.85%	0.39%***	4.72%
MFNews	-0.074***	0.360	-0.167	-0.031***	0.056	0.003***	0.123	-0.038	0.011***	0.042
BadNews	0.601***	0.490	0.000	1.000***	1.000	0.488***	0.500	0.000	0.000***	1.000
Range_indicator	0.797***	0.402	1.000	1.000***	1.000	0.883***	0.322	1.000	1.000***	1.000
Bundled_indicator	0.035*	0.184	0.000	0.000*	0.000	0.694***	0.461	0.000	1.000***	1.000

This table presents descriptive statistics of the variables for diverse sets of sample firms used for testing H1. The sample sets in the left column include management forecasts of quarterly earnings, while those in the right column include management forecasts of annual earnings. In constructing each set of the sample, we follow the sample selection procedures of Kothari et al. (2009): (1) there must be an analyst EPS forecast for the current period within 30 days prior to the MEF, (2) the analyst forecast consensus is larger than 5 cents, (3) the news in the management forecast is larger than 1% of the analyst forecast consensus, and (4) the top 1% of observations are curtailed. The definitions of *CAR, MFNews,* and *BadNews* are presented in Appendix A. *Range_indicator* is coded 1 if the management forecast was issued as a range forecast, while *Bundled_indicator* is coded 1 if the management forecast was issued along with an earnings announcement. The mean and median values in Panel C and D are compared to those in Panel B. ***, **, and * represent significance at the two-tailed one, five, and ten percent levels, respectively.

TABLE 2
Replication of Kothari, Shu, and Wysocki (2009)

Panel A: Univariate analysis of quarterly management forecasts and market reactions

	Good news forecasts			Bad news forecasts		
		MFNews	*CAR*		*MFNews*	*CAR*
		Mean	Mean		Mean	Mean
	# Obs.	(Median)	(Median)	# Obs.	(Median)	(Median)
Full sample	1,372	0.197	0.023	2,740	-0.388	-0.100
N = 4,112		(0.087)	(0.024)		(-0.223)	(-0.068)
Pre Reg FD	658	0.197	0.017	1,542	-0.416	-0.122
N = 2,200		(0.087)	(0.018)		(-0.256)	(-0.093)
Post Reg FD	714	0.196	0.029	1,198	-0.351	-0.071
N = 1,912		(0.086)	(0.030)		(-0.188)	(-0.036)
Post Reg FD – Pre Reg FD:		*MFNews*	*CAR*		*MFNews*	*CAR*
	Mean	-0.001	0.012**	Mean	0.065***	0.051***
	Median	-0.001	0.012**	Median	0.068***	0.057***

Panel B: Regression analysis of market reactions to quarterly management forecasts

	Model (1)		Model (2)	
Variable	Coefficient	(*t-stat*)	Coefficient	(*t-stat*)
Intercept	0.023***	(*7.59*)	0.013***	(*3.55*)
BadNews	-0.123***	(*-28.77*)	-0.076***	(*-15.09*)
MFNews			0.053***	(*4.21*)
MFNews×BadNews			0.041***	(*2.84*)
Adj. R^2	14.10%		20.34%	
N	4,112		4,112	
F-test: *Intercept+BadNews* < -1×*Intercept*				
Difference / p-value:	-0.100***	*< 0.001*	-0.063***	*< 0.001*

This table presents the results from the univariate and regression analyses of market reactions to quarterly management earnings forecasts, replicating the documented results in Kothari et al. (2009). The definitions of each variable are presented in Appendix A. Presented *t*-statistics are adjusted for heteroskedasticity (White 1980). ***, **, and * represent significance at the two-tailed one, five, and ten percent levels, respectively. The *p*-values in the *F*-tests are assessed at the two-tailed level.

TABLE 3

Regression analysis using quarterly earnings forecasts (as in KSW 2009)

Panel A: Using the midpoint as investors' interpretation of managements' expectations (based on Kothari, Shu, and Wysocki 2009)

	Prior to Regulation FD (1995 – 2000)		After Regulation FD (2001 – 2007)	
Variable	Coefficient	(*t-stat*)	Coefficient	(*t-stat*)
Intercept	0.007	(*1.20*)	0.023***	(*10.34*)
BadNews	-0.092***	(*-12.35*)	-0.048***	(*-14.64*)
MFNews	0.050**	(*2.51*)	0.049***	(*5.29*)
MFNews×BadNews	0.039*	(*1.74*)	0.065***	(*5.03*)
Adj. R^2	21.75%		21.23%	
N	2,200		4,617	
F-test: *Intercept+BadNews* < -1×*Intercept*				
Difference / p-value:	-0.085***	*< 0.001*	-0.025	(*0.70*)

Panel B: Using the upper bound as investors' interpretation of managements' expectations (based on Ciconte et al. 2014) and controlling for earnings surprise (based on Rogers and Van Buskirk 2013)

	Prior to Regulation FD (1995 – 2000)		After Regulation FD (2001 – 2007)	
Variable	Coefficient	(*t-stat*)	Coefficient	(*t-stat*)
Intercept	0.016***	(*3.88*)	0.014***	(*6.86*)
BadNews	-0.121***	(*-17.86*)	-0.052***	(*-13.91*)
MFNews	0.002	(*1.40*)	0.054***	(*5.27*)
MFNews×BadNews	0.051***	(*4.92*)	0.061***	(*3.84*)
EarnSurp	4.687	(*1.44*)	3.098	(*0.26*)
EarnSurp×BadNews	39.237*	(*1.71*)	44.225**	(*2.20*)
Adj. R^2	21.66%		21.22%	
N	2,200		4,617	
F-test: *Intercept+BadNews* < -1×*Intercept*				
Difference / p-value:	-0.105***	*< 0.001*	-0.038***	*< 0.001*

This table presents the results from the regression analysis of market reactions to management forecasts of quarterly earnings for two time periods: pre- and post-Reg FD. Panel A presents regression results when precisely following the sample selection procedure and the computation of management forecast news used in Kothari et al. (2009). Panel B presents the regression results when re-computing the management forecast news using the upper bound of range forecasts for the same sample (Ciconte et al. 2014), and controlling for the earnings surprise in any concurrent earnings announcement (Rogers and Van Buskirk 2013). Presented t-statistics are adjusted for heteroskedasticity (White 1980). ***, **, and * represent significance at the two-tailed one, five, and ten percent levels, respectively. The *p*-values in the *F*-tests are assessed at the two-tailed level.

TABLE 4
Regression analysis using annual forecasts

Panel A: Using the midpoint as investors' interpretation of managers' expectations (based on Kothari, Shu, and Wysocki 2009)

	Prior to Regulation FD (1995 – 2000)		After Regulation FD (2001 – 2007)	
Variable	Coefficient	(*t-stat*)	Coefficient	(*t-stat*)
Intercept	0.012***	(*2.75*)	0.024***	(*18.67*)
BadNews	-0.053***	(*-7.39*)	-0.034***	(*-17.11*)
MFNews	-0.008	(*-0.52*)	0.054***	(*4.05*)
MFNews×BadNews	0.178***	(*4.13*)	0.191***	(*8.20*)
Adj. R^2	9.89%		14.33%	
N	1,301		9,800	
F-test: *Intercept+BadNews* < -1×*Intercept*				
Difference / p-value:	-0.041***	*< 0.001*	-0.010	(*0.99*)

Panel B: Using the upper bound as investors' interpretation of managers' expectations (based on Ciconte et al. 2014) and controlling for the earnings surprise (based on Rogers and Van Buskirk 2013)

	Prior to Regulation FD (1995 – 2000)		After Regulation FD (2001 – 2007)	
Variable	Coefficient	(*t-stat*)	Coefficient	(*t-stat*)
Intercept	0.008*	(*1.90*)	0.012***	(*10.27*)
BadNews	-0.058***	(*-7.80*)	-0.033***	(*-14.55*)
MFNews	0.000	(*0.01*)	0.054***	(*5.44*)
MFNews×BadNews	0.143***	(*3.10*)	0.189***	(*6.70*)
EarnSurp	9.333***	(*4.10*)	5.000***	(*10.75*)
EarnSurp×BadNews	-9.173**	(*-2.25*)	-1.221	(*-1.12*)
Adj. R^2	10.04%		15.68%	
N	1,301		9,800	
F-test: *Intercept+BadNews* < -1×*Intercept*				
Difference / p-value:	-0.050***	*< 0.001*	-0.021***	*< 0.001*

This table presents the results from the regression analysis of market reactions to management forecasts of annual earnings for the two sets of samples: pre- and post-Reg FD. Panel A presents regression results when precisely following Kothari et al.'s (2009) sample selection procedure and computation of management forecasts news. Panel B presents the regression results when re-computing the management forecast news using the upper bound of range forecasts for the same sample (Ciconte et al. 2014), and controlling for earnings surprise in an accompanying earnings announcement (Rogers and Van Buskirk 2013). Presented t-statistics are adjusted for heteroskedasticity (White 1980). ***, **, and * represent significance at the two-tailed one, five, and ten percent levels, respectively. The *p*-values in the *F*-tests are assessed at the two-tailed level.

TABLE 5

Sample selection for H2

Panel A: Sample selection procedure	
	Number of MEF
Annual management earnings forecasts (MEF) for fiscal years 2006 and 2007 (beginning after 12/15/2006) issued by firms with PERMNO, CUSIP, GVKEY, and IBES TICKER	10,138
Less:	
Open-ended or qualitative MEF	259
Those without the CEO's compensation information on Execucomp	1,450
Those without the CEO's severance payment information on the firm's 10-K report	1,770
Those without an analyst EPS forecast issued within a month prior to the MEF	2,803
Those with (1) the analyst forecast consensus less than 5 cents; (2) the news in MEF less than 1% of analyst forecast consensus; or (3) the extreme news in MEF at the top 1% level (Kothari et al. 2009)	950
Management earnings forecasts for empirical tests	2,906

Panel B: Industry Composition		
Two-Digit SIC Industry Sector	Number of MEF	% of Sample Forecasts
Agriculture (01-09)	10	0.34%
Mining (10-14)	37	1.27%
Construction (15-17)	45	1.55%
Manufacturing (20-39)	1,224	42.12%
Telecommunication, Transportation, and Utilities (40-49)	406	13.97%
Wholesale (50-51)	101	3.48%
Retailing (52-59)	386	13.28%
Services (70-88)	402	13.83%
Other	295	10.15%
Total	2,906	100.00%

This table shows our sample selection and the composition of management earnings forecasts in our sample used for testing H2. Panel A provides details of the sample selection procedures. Panel B shows the distribution of management earnings forecasts across two-digit SIC industry sectors.

TABLE 6
Descriptive statistics for H2

Panel A: Descriptive statistics

Variables	Mean	Std dev	P25	Median	P75
CAR	0.004	0.074	-0.032	0.004	0.042
MFNews (using upper bound)	0.035	0.215	-0.012	0.025	0.054
BadNews	0.271	0.444	0.000	0.000	1.000
SevAmt (in thousands)	7,313.8	14,461.7	0.0	2,630.1	8,420.0
CashComp (in thousands)	1,050.3	997.2	657.4	900.0	1,125.0
SevAmt / CashComp	6.817	11.629	0.000	3.304	9.050

Panel B: Correlation matrix

	CAR	*MFNews*	*BadNews*	*SevAmt*	*Cash Comp*
MFNews	**0.168**				
BadNews	**-0.347**	**-0.323**			
SevAmt	0.015	-0.024	**0.052**		
CashComp	-0.002	-0.028	**0.031**	**0.289**	
SevAmt / CashComp	**0.033**	-0.016	**0.032**	**0.860**	0.013

Panel A presents descriptive statistics of the variables used for empirical tests. The definitions of *CAR*, *MFNews*, and *BadNews* are presented in Appendix A. *SevAmt* is the dollar amount (in thousands) of the ex ante severance pay agreement between the CEO and the firm, disclosed in the firm's SEC 10-K filing. *CashComp* is the sum of the CEO's salary and bonus compensation in that year. Panel B exhibits the Pearson correlation among the variables. Bold text in the correlation matrices indicates a statistical significance of the correlation at the 10%, two-tailed level.

TABLE 7

Regression analysis of the effects of severance payment on management forecasts
(Using the upper bound of management forecasts and controlling for earnings surprise)

Panel A: Univariate analysis of management forecasts and market reactions

	Good news forecasts			Bad news forecasts		
	# Obs.	*MFNews* Mean (Median)	*CAR* Mean (Median)	# Obs.	*MFNews* Mean (Median)	*CAR* Mean (Median)
Full sample	2,119	0.078	0.030	787	-0.079	-0.023
N = 2,906		(0.037)	(0.022)		(-0.035)	(-0.013)
SevPay = 0 (Low)	1,589	0.081	0.019	592	-0.079	-0.044
N = 2,181		(0.038)	(0.012)		(-0.037)	(-0.033)
SevPay = 1 (High)	530	0.067	0.020	195	-0.078	-0.021
N = 725		(0.034)	(0.015)		(-0.030)	(-0.013)
High SevPay – Low SevPay:		*MFNews*	*CAR*		*MFNews*	*CAR*
	Mean	-0.014	0.001	Mean	0.001	0.023***
	Median	-0.004**	0.003	Median	0.007**	0.020***

Panel B: Regression analysis of market reactions to management forecasts

	Model (1)		Model (2)	
Variable	Coeff.	(*t-stat*)	Coeff.	(*t-stat*)
Intercept	0.013***	(*8.63*)	0.012***	(*8.16*)
BadNews	-0.049***	(*-15.82*)	-0.043***	(*-12.10*)
MFNews			0.006	(*1.22*)
MFNews×BadNew			0.066**	(*2.10*)
EarnSurp	6.925***	(*8.80*)	6.890***	(*8.77*)
EarnSurp×BadNews	2.054	(*1.24*)	1.210	(*0.71*)
Adj. R^2	18.11%		18.50%	
N	2,906		2,906	
F-test: *Intercept+BadNews* < -1×*Intercept*				
Difference / p-value:	-0.036***	*< 0.001*	-0.031***	*< 0.001*

	Model (4)		Model (5)		Model (6)	
Variable	Coeff.	(*t-stat*)	Coeff.	(*t-stat*)	Coeff.	(*t-stat*)
Intercept	0.013***	(*7.59*)	0.012***	(*7.15*)	0.012***	(*7.27*)
BadNews	-0.054***	(*-14.67*)	-0.048***	(*-11.70*)	-0.047***	(*-11.07*)
SevPay	0.001	(*0.49*)	0.002	(*0.51*)	-0.000	(*-0.11*)
SevPay×BadNews	0.019***	(*2.81*)	0.019***	(*2.86*)	0.018**	(*2.51*)
MFNews			0.006	(*1.23*)	0.003	(*0.57*)
MFNews×BadNews			0.066**	(*2.14*)	0.080**	(*2.10*)
MFNews×SevPay					0.028	(*1.50*)
MFNews×BadNews×SevPay					-0.058	(*-0.87*)
EarnSurp	6.928***	(*8.82*)	6.894***	(*8.78*)	6.939***	(*8.89*)

EarnSurp×BadNews	1.820	(*1.13*)	0.964	(*0.58*)	0.876	(*0.53*)
Adj. R^2	18.43%		18.83%		18.85%	
N	2,906		2,906		2,906	

F-test: *Intercept+BadNews* < -1×*Intercept*

	Sum of coeff	*p-value*	Sum of coeff	*p-value*	Sum of coeff	*p-value*
	-0.041	< *0.001*	-0.036	< *0.001*	-0.035	< *0.001*

F-test: *Intercept+BadNews+SevPay* + *SevPay×BadNews* < -1×(*Intercept+SevPay*)

	Sum of coeff	*p-value*	Sum of coeff	*p-value*	Sum of coeff	*p-value*
	-0.021	(*0.20*)	-0.015	(*0.75*)	-0.017	(*0.41*)

This table presents the results from the univariate and regression analyses of the effects of severance payments on market reactions to management earnings forecast (MEF) when using the upper bound of MEF. The definitions of each variable are presented in Appendix A. Presented *t*-statistics are adjusted for heteroskedasticity (White 1980). ***, **, and * represent significance at the two-tailed one, five, and ten percent levels, respectively. The *p*-values in the *F*-tests are assessed at the two-tailed level.

TABLE 8

The effects of severance payment on management forecasts after controlling for litigation risk, insiders' ownership, and determinants of ex-ante severance pay

	Model (7)											
	Column (1)		Column (2)		Column (3)		Column (4)		Column (5)		Column (6)	
Variable	Coeff.	(*t-stat*)	Coeff.	(*t-stat*)	Coeff.	(*t-stat*)	Coeff.	(*t-stat*)	Coeff.	(*t-stat*)	Coeff.	(*t-stat*)
Intercept	0.007***	(*3.78*)	0.007***	(*3.48*)	0.007**	(*2.14*)	0.005**	(*2.00*)	0.002	(*0.68*)	0.000	(*0.08*)
BadNews	-0.040***	(*-9.11*)	-0.035***	(*-7.09*)	-0.034***	(*-6.75*)	-0.030***	(*-5.51*)	-0.023***	(*-3.36*)	-0.020***	(*-2.91*)
SevPay	0.002	(*0.66*)	0.002	(*0.69*)	0.002	(*0.60*)	0.002	(*0.61*)	0.002	(*0.64*)	0.001	(*0.37*)
SevPay×BadNews	0.015**	(*2.34*)	0.015**	(*2.39*)	0.016**	(*2.39*)	0.017**	(*2.48*)	0.016**	(*2.42*)	0.017**	(*2.55*)
MFNews			0.006	(*1.19*)			0.005	(*0.87*)	0.005	(*1.10*)	0.004	(*0.79*)
MFNews×BadNews			0.064**	(*2.17*)			0.061**	(*2.04*)	0.065**	(*2.13*)	0.061**	(*2.00*)
EarnSurp	6.836***	(*8.80*)	6.803***	(*8.76*)	6.925***	(*8.71*)	6.880***	(*8.63*)	6.708***	(*8.64*)	6.782***	(*8.55*)
EarnSurp×BadNews	1.811	(*1.11*)	0.987	(*0.59*)	1.010	(*0.61*)	0.296	(*0.17*)	0.932	(*0.56*)	0.288	(*0.17*)
HiLitRisk	-0.001	(*-0.48*)	-0.001	(*-0.48*)	-0.001	(*-0.22*)	-0.001	(*-0.23*)	-0.001	(*-0.65*)	-0.001	(*-0.17*)
HiLitRisk×BadNews	0.009	(*1.46*)	0.009	(*1.42*)	0.009	(*1.36*)	0.009	(*1.29*)	0.007	(*1.43*)	0.007	(*0.98*)
HiAsymm	0.011***	(*3.85*)	0.011***	(*3.84*)	0.009***	(*2.91*)	0.009***	(*2.91*)	0.009***	(*3.32*)	0.007**	(*2.33*)
HiAsymm×BadNews	-0.033***	(*-5.39*)	-0.033***	(*-5.30*)	-0.029***	(*-4.41*)	-0.029***	(*-4.33*)	-0.030***	(*-4.41*)	-0.026***	(*-4.02*)
HiDistress	0.011	(*0.34*)	0.011	(*0.34*)	0.032	(*1.31*)	0.032	(*1.31*)	0.007	(*1.31*)	0.022	(*0.93*)
HiDistress×BadNews	-0.038	(*-0.96*)	-0.042	(*-1.04*)	-0.059*	(*-1.66*)	-0.062*	(*-1.74*)	-0.035*	(*-1.70*)	-0.050	(*-1.51*)
HiInsideOwn					0.007**	(*2.53*)	0.007**	(*2.49*)			0.006**	(*2.03*)
HiInsideOwn×BadNews					-0.015**	(*-2.34*)	-0.014**	(*-2.17*)			-0.010	(*-1.61*)
OutsideCEO									0.004	(*1.15*)	0.004	(*1.20*)
OutsideCEO×BadNews									-0.016***	(*-2.59*)	-0.015**	(*-2.15*)
CEOown									0.000*	(*1.81*)	0.000	(*1.08*)
CEOown×BadNews									-0.000***	(*-3.59*)	-0.000***	(*-2.65*)
CEOtenure									-0.000	(*-0.70*)	-0.000	(*-0.83*)
CEOtenure×BadNews									-0.000	(*-0.25*)	-0.000	(*-0.31*)

Noncompete									-0.000	(*-0.10*)	0.001	(*0.29*)
Noncompete×BadNews									-0.007	(*-1.20*)	-0.006	(*-0.87*)
Industry Effects	NO		NO		NO		NO		YES		YES	
Adj. R^2	19.27%		19.64%		18.73%		19.04%		20.33%		19.71%	
N	2,906		2,906		2,736		2,736		2,906		2,736	
F-test: *Intercept* + *BadNews* < -1×*Intercept*												
Intercept + *BadNews*	-0.033		-0.028		-0.027		-0.025		-0.021		-0.020	
p-value	*< 0.001*		*< 0.001*		*< 0.001*		*< 0.001*		(*0.01*)		(*0.01*)	
F-test: *Intercept*+*BadNews*+*SevPay*+*SevPay*×*BadNews* < -1×(*Intercept*+*SevPay*)												
Intercept + *BadNews* + *SevPay* + *SevPay*×*BadNews*	-0.016		-0.011		-0.009		-0.006		-0.003		-0.002	
p-value	(*0.38*)		(*0.93*)		(*0.55*)		(*0.99*)		(*0.88*)		(*0.97*)	

This table presents the results from the regression analyses of the effects of severance payments on market reactions to management earnings forecast (MEF) when using the upper bound of MEF and controlling for factors affecting managers' disclosure of bad news (Kothari et al. 2009) and CEO characteristics influencing the firm's ex ante severance payment agreement (Cadman et al. 2014). The definitions of each variable are presented in Appendix A. Presented *t*-statistics are adjusted for heteroskedasticity (White 1980). ***, **, and * represent significance at the two-tailed one, five, and ten percent levels, respectively. The *p*-values in the *F*-tests are assessed at the two-tailed level.

TABLE 9

Regression analysis of the effects of severance payment on management forecasts
(Using the upper bound of management forecasts and controlling for earnings surprise)

Panel A: Effects of non-competition enforcement (Garmaise 2011) on bad news delay

	Column (1)		Column (2)	
Variable	Coeff.	(*t-stat*)	Coeff.	(*t-stat*)
Intercept	0.013***	(*8.26*)	0.012***	(*6.04*)
BadNews	-0.041***	(*-10.89*)	-0.032***	(*-7.16*)
Garmaise	0.002	(*0.45*)	0.000	(*-0.09*)
Garmaise×BadNews	-0.019*	(*-1.79*)	-0.007	(*-0.54*)
MFNews	0.018	(*1.62*)	0.034	(*1.64*)
MFNews×BadNews	0.103***	(*2.91*)	0.156***	(*2.68*)
EarnSurp	4.906***	(*8.14*)	5.386***	(*6.12*)
EarnSurp×BadNews	-2.105*	(*-1.82*)	0.347	(*0.19*)
Adj. R^2	15.37%		16.56%	
N	3,934		2,272	
F-test: *Intercept* + *BadNews* < -1×*Intercept*				
Intercept + BadNews	-0.028		-0.020	
p-value	*< 0.001*		(*0.06*)	
F-test: *Intercept* + *BadNews* + *Garmaise* + *Garmaise×BadNews* < -1×(*Intercept* + *Garmaise*)				
Intercept + BadNews + Garmaise + Garmaise×BadNews	-0.044		-0.027	
p-value	(*0.01*)		(*0.20*)	

Panel B: Effects of non-competition agreements vs. severance pay contracts on bad news delay

	Column (1)		Column (2)		Column (3)		Column (4)	
Variable	Coeff.	(*t-stat*)	Coeff.	(*t-stat*)	Coeff.	(*t-stat*)	Coeff.	(*t-stat*)
Intercept	0.013***	(*8.14*)	0.013***	(*5.35*)	0.013***	(*6.73*)	0.013***	(*6.23*)
BadNews	-0.043***	(*-11.66*)	-0.043***	(*-6.23*)	-0.043***	(*-9.47*)	-0.047***	(*-9.68*)
Garmaise	-0.001	(*-0.17*)			-0.001	(*-0.13*)	-0.001	(*-0.12*)
Garmaise×BadNews	-0.005	(*-0.41*)			-0.011	(*-0.69*)	-0.011	(*-0.65*)
Noncomp			-0.001	(*-0.35*)	-0.001	(*-0.32*)	-0.001	(*-0.36*)
Noncomp×BadNews			0.000	(*-0.03*)	-0.001	(*-0.12*)	-0.002	(*-0.31*)
Garmaise×Noncomp×BadNews					0.011	(*0.55*)	0.011	(*0.48*)
SevPay							0.002	(*0.54*)
SevPay×BadNews							0.019***	(*2.87*)
MFNews	0.006	(*1.21*)	0.006	(*1.64*)	0.006	(*1.19*)	0.006	(*1.21*)
MFNews×BadNews	0.065**	(*2.09*)	0.066**	(*2.10*)	0.065**	(*2.07*)	0.066**	(*2.10*)
EarnSurp	6.886***	(*8.80*)	6.886***	(*8.78*)	6.883***	(*8.81*)	6.887***	(*8.82*)
EarnSurp×BadNews	1.217	(*0.72*)	1.212	(*0.72*)	1.195	(*0.71*)	0.947	(*0.57*)
Adj. R^2	18.46%		18.45%		18.39%		18.72%	
N	2,906		2,906		2,906		2,906	

F-test: *Intercept* + *BadNews* < -1×*Intercept*				
Intercept + *BadNews*	-0.030	-0.030	-0.030	-0.034
p-value	*< 0.001*	*< 0.001*	*< 0.001*	*< 0.001*
F-test: *Intercept* + *BadNews* + *Garmaise* + *Garmaise*×*BadNews* < -1×(*Intercept* + *Garmaise*)				
Intercept + *BadNews* + *Garmaise* + *Garmaise*×*BadNews*	-0.036		-0.042	-0.046
p-value	(*0.03*)		(*0.07*)	(*0.05*)
F-test: *Intercept* + *BadNews* + *Noncomp* + *Noncomp*×*BadNews* < -1×(*Intercept* + *Noncomp*)				
Intercept + *BadNews* + *Noncomp* + *Noncomp*×*BadNews*		-0.031	-0.031	-0.037
p-value		*< 0.001*	*< 0.001*	*< 0.001*
F-test: *Intercept*+*BadNews*+*SevPay*+*SevPay*×*BadNews* < -1×(*Intercept*+*SevPay*)				
Intercept + *BadNews* + *SevPay* + *SevPay*×*BadNews*				-0.014
p-value				(*0.94*)

This table presents the results from the regression analyses of the effects of non-competition enforcement of a firm's headquarter state, firm-specific non-competition agreements, and severance payments on market reactions to management earnings forecast (MEF) when using the upper bound of MEF and controlling for concurrent earnings announcement news. Garmaise is coded 1, if the state in which firm's headquarter is located is regarded as more strongly enforcing non-competition agreements relative to other states, as measured by Garmaise (2011) index value equal to or greater than 7. The definitions of other variables are presented in Appendix A. Presented *t*-statistics are adjusted for heteroskedasticity (White 1980). ***, **, and * represent significance at the two-tailed one, five, and ten percent levels, respectively. The *p*-values in the *F*-tests are assessed at the two-tailed level.