The Interaction between Mandatory Reporting and Voluntary Disclosure and Their Relevance to Equity Market and Credit Market

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Abstract

Mandatory financial reporting is subject to generally accepted accounting principles while voluntary disclosure such as management forecasts could be used by managers to reveal their inside information to market participants. This dissertation examines the interaction between mandatory reporting and voluntary disclosure and their relevance to equity market and credit market.

Chapter 1 provides the outline of the dissertation and discusses the major contributions. Chapter 2 reviews previous literature and develops the hypotheses. Management forecasts about future earnings are sometimes issued along with current quarter earnings announcements, and these bundled management forecasts have recently become more prevalent. Using a composite measure of ex-ante management forecast accuracy that takes into account forecast ability, forecast difficulty and forecast environment, Chapter 3 shows that the bundled management forecasts can mitigate investors' under-reaction to current earnings and reduce the magnitude of post-earnings announcement drift only when these forecasts have high ex-ante accuracy.

Firms have the incentive to cater for capital market's demand in their financial reporting. Prior research has provided evidence that conservatism can lower the debt cost and solve the interest conflict between bondholders and shareholders. If firms anticipate market's demand for conservatism before they issue public bonds, they will report more conservatively before issuing bonds for the first time. Using alternative measures of accounting conservatism, Chapter 4 shows that firms do report more conservatively before bond IPO. This result highlights the incremental importance of the debt market over the equity market in inducing conservative reporting and supports the argument that conservatism is more closely related with the debt market.

Besides the equity market, management forecasts are also value relevant to the credit market. Chapter 5 provides the evidence that credit default swap (CDS) spreads react significantly and negatively to management forecast news, and that these reactions are stronger than those to actual earnings news. The credit market reactions to bad management forecast news and forecasts issued by credit risky companies are larger, reflecting the asymmetric payoff of debt securities. The impact of management forecasts on CDS spreads, relative to earnings announcements, also becomes stronger during the recent credit crisis when the market uncertainty is greater.

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Five years' PhD life in London Business School is challenging and stimulating. The feeling of being fascinated by the classic papers is still very fresh, before I realize that I spend most of my time in following the recent working papers, identifying the research questions and trying to conduct the analysis independently.

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1. Introduction

This dissertation studies the interaction between mandatory reporting and voluntary disclosure and examines their relevance to equity market and credit market. The interaction between mandatory reporting and voluntary disclosure has not been adequately investigated in the prior literature (Beyer et al 2009), and Chapter 3 contributes to the literature by investigating the effect of bundled management forecasts and their ex-ante forecast accuracy on post-earnings announcement drift. Post-earnings announcement drift is one of the most intriguing market anomalies. Following earnings announcements, stock prices move in the same direction as that of earnings surprises for the subsequent 6 to 12 months. The magnitude of the drift also increases with that of earnings news, which is measured by the standardized unexpected earnings. Two main explanations have been advanced for this anomaly: a failure to adjust abnormal returns for risk, and a delayed response to earnings reports (Bernard and Thomas 1989). Existing evidence is more consistent with the market under-reaction explanation. Investors underestimate the implications of current earnings for future earnings, and their under-reaction is corrected at future earnings announcement dates.[1]

This paper examines a prediction of the under-reaction explanation by investigating the effect of bundled management forecasts and their ex-ante accuracy on post-earnings announcement drift. Management forecasts about future earnings are sometimes issued along with current quarter earnings

[1] See Bernard and Thomas (1989, 1990), Ball and Bartov (1996), Livnat and Mendenhall (2006), and Shivakumar (2006), among others.

announcements, and these bundled management forecasts have recently become more prevalent (Rogers and Van Buskirk 2009). [2] If post-earnings announcement drift is caused by investors' inefficiency in forming the expectations of future earnings upon current earnings news, the management forecasts of future earnings should accelerate investors' reaction. However, if investors perceive that the bundled management forecasts lack accuracy, they place less weight on the forecasts and keep extrapolating future earnings based on their own information sets. The ability of bundled management forecasts to mitigate post-earnings announcement drift should be dependent on their ex-ante (perceived) forecast accuracy.

The main empirical results suggest that the bundled management forecasts on average do not mitigate post-earnings announcement drift when other drift-related variables are controlled. Consistent with the prediction, the bundled management forecasts reduce the magnitude of post-earnings announcement drift only when they have high ex-ante forecast accuracy. The inferences hold after a battery of robustness checks.

This paper adds to the evidence supporting the under-reaction explanation for post-earnings announcement drift. Albeit the long discussion related with the existence and causes of post-earning announcement drift, there is limited evidence about the information provided by managers. More

[2] In this paper, bundled management forecasts refer to the forecasts issued within one trading day around the earnings announcement date, to yield post-earnings announcement drift window consistent with that in previous literature (Livnat and Mendenhall 2006). This approach is similar in spirit to that employed by Rogers and Van Buskirk (2009), in which bundled management forecasts refer to the forecasts issued within two days of the earnings announcement date.

importantly, by showing the different effects of accurate bundled forecasts and inaccurate bundled forecasts on post-earnings announcement drift, this paper goes above and beyond simply focusing on the issuance of bundled management forecasts and incorporates their perceived accuracy into the analysis. The evidence that the bundled management forecasts reduce the magnitude of post-earnings announcement drift only when they have high ex-ante forecast accuracy suggests prompt and accurate management forecasts of future earnings effectively resolve investors' uncertainty towards future earnings and mitigate their under-reaction to announced earnings.

Firms have the incentive to cater for capital market's demand in their financial reporting. Chapter 4 studies whether firms report more conservatively before their bond initial public offerings (Bond IPOs). Here conservatism refers to conditional conservatism- asymmetric timely loss recognition. There is evidence showing that conservative financial reporting exists primarily for the efficiency of debt market contracts. Ball et al. (2008) use an international sample to test the relative roles of debt market and equity market in determining the important properties of financial reporting and they find that the debt market rather than the equity market is associated with conditional conservatism. An alternative argument says that because financial statements serve mainly for public shareholders to solve the information asymmetry, conservatism- as an important reporting quality- exists primarily to serve the equity market.

Compared with public bond, bank debt or private debt is relationship financing and lenders rely on inside access to evaluate firms' financial ability.

However in the bond market, bondholders rely exclusively on financial statements to price the risk of each firm. Public bond market is an important financing source for firms raising capital. Firms need to build their reputation through borrowing from banks before entering the public bond market (Diamond 1991). On one hand, as part of reputation building, reporting quality is expected to increase before bond IPO. On the other hand, if debt market drives the financial reporting quality and firms anticipate market's information demand, firms will adjust their earnings quality right before bond IPO.

This paper uses a sample of US firms and examines whether firms, conditional on being publicly listed, report more conservatively before their bond IPOs. The main empirical test uses Basu regression and the result shows that firms report more conservatively at event year -1, -2 relative to the level of reporting conservatism in years before event year -2. As a confirmation, this paper uses the accrual-based model and shows that firms increase their reporting conservatism both at event year -1 and -2 before bond IPO when the reporting conservatism in years before event year -2 is used as the benchmark.

This paper contributes to the existing literature in several aspects. Compared with the extant research about common stock IPO, there are limited papers focusing on bond IPO. To my knowledge, this is the first paper to talk about how earnings quality changes before bond IPO. More importantly this paper provides evidence on the incremental importance of the debt market over the equity market for conservatism in reporting. Conditional on being publicly listed, firms increase their reporting conservatism significantly before bond IPO,

supporting the argument that conservatism is more closely related with the debt market (Ball et al 2008).

Chapter 5 is based on a paper, co-authored with Dr Lakshmanan Shivakumar, Dr Oktay Urcan and Dr Florin Vasvari. The recent credit crisis has emphasized the importance of a good understanding of the information based on which credit instruments are priced. A significant body of literature examines the role of macroeconomic, industry-specific and firm-specific factors on the credit market in a variety of settings.[3] However, surprisingly little empirical research has been done to evaluate the credit market's use of earnings-related information, even though prior research finds that earnings information can predict firm bankruptcies (e.g. Altman 1968; Beaver 1968; Ohlson 1980).[4] In addition, there is no evidence on the use of earnings-related information when credit markets are under heightened uncertainty and information asymmetry. We extend the literature by investigating the credit market's response to management earnings forecasts during the periods before and after the onset of the credit crisis.

In contrast to equity markets, credit markets would ignore management forecasts if the forecasts are viewed primarily as mechanisms to communicate with shareholders and contain little credible information about a firm's downside risks, which is the major concern of debtholders. Prior studies yield mixed evidence on whether earnings forecasts contain credible information for

[3] There are lots of papers that investigate drivers of bond spreads. More recently, papers have started to focus on the spreads of credit default swaps (e.g., Jorion and Zhang 2007; Zhang et al. 2009).

[4] Callen et al. (2009) and Easton et al. (2009) investigate the role of earnings in credit markets. We discuss these papers in greater detail later.

evaluating downside risks. On the one hand, Hayn (1995) shows that accounting earnings do not reflect the liquidation option value, which causes negative earnings to have little impact on firm valuation. Similarly, Barth, Beaver and Landsman (1998) find that earnings are less important in valuation when the financial health of the firm deteriorates. In addition, Koch (2002) and Rogers and Stocken (2005) document that management earnings forecasts issued by firms in poorer financial health tend to be biased, lowering their credibility. On the other hand, Kasznik and Lev (1995) and Hutton and Stocken (2009) show that managers often use management forecasts to release bad news on a timely basis.

Several studies document that market reactions to information are time-varying, depending on, among other factors, the level of uncertainty among market participants and investor sentiment (e.g., Lang 1991; Mian and Sankaraguruswamy 2008). These studies find that the equity market reaction to news is larger during periods characterized by greater uncertainty, and that investors react more negatively to bad news when investor sentiment is low. Extrapolating these findings to the credit markets, we argue that management forecasts are likely to be most informative during periods of greater uncertainties and low investor sentiment, and that their reaction to bad forecast news, in particular, would be stronger during these periods. We test our prediction by exploiting the exogenous shocks to market uncertainty and investor sentiment that occurred during the recent credit crisis. This analysis, besides potentially helping us better understand the credit market's use of

management forecasts, could also provide insights into the tools that managers could employ during such crises to mitigate informational asymmetries.

We evaluate the credit market's use of management forecasts in two ways. First, we evaluate the credit market reactions to management forecasts by themselves. Second, we compare the importance of management forecast to credit markets with that of mandated earnings announcements. We document that CDS spread changes are significantly and negatively associated with management forecast news, calculated as the proportional deviation of the management earnings forecasts from the most recent consensus analyst earnings forecasts. The economic effects of the forecasts on CDS spreads are also highly significant. The results show that credit markets view management forecasts as a credible source of information on a firm's downside risks.

Consistent with arguments that greater uncertainty amplifies the sensitivity of prices to news, we find that the credit market reactions are significantly greater for the forecasts issued during the credit crisis period (July 2007 to December 2008) than for forecasts issued in the pre-crisis period. However, in contrast to the predictions based on the investor sentiment, we find that CDS spread reactions are larger for both good news and bad news during the crisis period.

In addition, during the crisis, the coefficient of accounting earnings news becomes insignificant while the coefficient of management forecasts news doubles in magnitude, suggesting that, during periods of higher uncertainty, information about future earnings, even if unverifiable, is more price relevant to

credit markets than the backward-looking audited earnings numbers.

We also examine whether management forecasts by a firm convey information on credit risks to other firms within the same industry. This analysis is motivated by the theoretical model of Giesecke (2004), which predicts a contagion effect in credit spreads when borrowers' fundamentals are not publicly known, and by prior empirical research that documents industry-level clustering in Chapter 11 bankruptcies (e.g., Lang and Stulz 1992; Jorion and Zhang 2007). We find that mean industry CDS spread changes are significantly negatively related to forecast news issued by a firm within the same industry. Although we observe this contagion effect in both the pre-crisis and the crisis periods, the magnitude of the effect is significantly larger during the crisis period, consistent with the analytical framework of Giesecke (2004), which indicates larger contagion when uncertainty and information asymmetry are higher.

To the best of our knowledge we are the first to study the relevance of management earnings forecasts to credit market, and by doing so we make a significant contribution both to the literature on credit pricing and to that on management earnings forecasts. Our findings provide strong evidence that management earnings forecasts are an important disclosure mechanism for the credit market, especially during the financial crisis.

2. Literature Review and Hypothesis Development

2.1. Bundled Management Forecasts and Post-Earnings

Announcement Drift

In addition to disclosing mandatory financial reports, firms also release financial information voluntarily. Voluntary disclosure, especially forecasts of future performance, is not subject to the restrictions of accounting standards, and could be used to pass managers' private information to investors. Management forecasts are good proxies for voluntary disclosure because they can be precisely measured and their issuance time is known (Healy and Palepu 2001).

Managers may choose to provide voluntary disclosures along with their mandatory financial reports, such as announcing current earnings and their forecasts about future earnings simultaneously. After the enactment of Regulation Fair Disclosure, this phenomenon has become more common, partly because of the greater use of earnings-related conferences calls along with or close to an earnings announcement (Rogers and Van Buskirk 2009).[5] If investors think these bundled management forecasts are accurate, such forecasts will help them understand the implications of current earnings for future earnings, and therefore decrease post-earnings announcement drift. This is a prediction that stems from the under-reaction explanation for post-earnings announcement drift.

The previous literature has provided two main explanations for

[5] Using a sample of 518 firms that initiated conference calls between 1994 and 2000, Kimbrough (2005) shows that the initiation of conference calls leads investors to respond in a more timely fashion to current earnings surprises. Initiation of conference calls is a one-off event in a firm's life, which says very little about how firms can affect post-earnings announcement drift on a regular basis by issuing bundled management forecasts.

post-earnings announcement drift: a failure to adjust abnormal returns for risk, and a delayed response to earnings reports (Bernard and Thomas 1989). Existing evidence is more consistent with the market under-reaction explanation. Bernard and Thomas (1990) find that a disproportionate fraction of the drift is concentrated around future earnings announcement dates, which suggests that the under-reaction to current earnings is corrected when future earnings are announced. Bartov et al. (2000) use the institutional holding as a proxy for investor sophistication, and provide evidence that investor sophistication can decrease the magnitude of post-earnings announcement drift because sophisticated investors can 'characterize correctly the process underlying earnings'. Shivakumar (2006) shows that the unexpected cash flows induce greater magnitude of post-earnings announcement drift than the unexpected accruals, which cannot be explained by the risk-based theory.

The under-reaction explanation suggests that the speed with which investors incorporate the implications of current earnings into their expectations of future earnings is associated with the magnitude of the drift. Soffer and Lys (1999) provide the evidence that investors' expectations of future earnings do not reflect the implications of the current earnings up to 15 trading days after they are announced. If managers issue their forecasts of future earnings along with the current earnings announcements, investors have more information to resolve the uncertainty related with the future period's earnings. To the extent that the bundled management forecasts can mitigate investors' under-reaction and post-earnings announcement drift, I expect investors to evaluate the

accuracy of the bundled management forecasts.

Managers' intentional bias leads to lower level of the forecast accuracy. Managers have incentives to bias their forecasts opportunistically. Rogers and Stocken (2005) examine the incentives caused by litigation environment, insider transactions, financial distress and industry concentration. Ertimur et al. (2007) find management forecasts to be less optimistically biased if forecasts are made and verified by actual earnings before IPO lockup expiration, when insider selling is forbidden. Rational expectations theory implies investors will use all the available information to estimate the bias and adjust for it. Managers anticipate investors' response, but continue to behave myopically (Stein 1989). In practice, investor's adjustment may not be complete. Some incentives, such as insider trading incentive, are not observed by investors. Fischer and Verrecchia (2000) find that the adjustment for the reporting bias decreases when the uncertainty towards the manager's objective increases. Managers are also more inclined to bias their forecasts if it is more difficult for market participants to detect this (Rogers and Stocken 2005).

Absent the incentive to bias the forecasts, managers may not be able to forecast accurately when the forecast difficulty level is high. When firms' operating environment is volatile, earnings generating process is subject to multiple contingencies. Even though managers truthfully reveal the inside information through management forecasts, these forecasts are not adequately accurate. Investors' perceived accuracy towards these forecasts is reduced accordingly.

Managers also have the incentive to build a forecast reputation. Stocken (2000) argues that in a repeated game setting, if accounting reports are useful enough to confirm or refute previously released voluntary information, and voluntary disclosure precision is evaluated over a sufficiently long period, managers have the inclination to release their private information fairly. Current management forecast accuracy is expected to be positively related with prior management forecast accuracy. This accuracy 'momentum' is a reflection of managers' efforts to build a forecast reputation. Prior forecast accuracy could also be viewed as a proxy for forecast ability.

Previous literature documents that investors are likely to estimate the accuracy of the management forecasts and incorporate it into their reaction to the forecast news. Rogers and Stocken (2005) investigate the market's response to the predicted forecast errors and find that, for good news management forecasts, the market's response varies with the predicted forecast errors. Ng et al. (2008) find that forecast credibility mitigates investors' under-reaction to management forecast news. They use various forecast credibility measures, including forecast precision, prior forecast accuracy, firm level litigation risk, industry concentration, and research and development intensity.

This paper argues that only when the bundled management forecasts have high ex-ante accuracy could they mitigate post-earnings announcement drift. [6] If the bundled forecasts are perceived to be accurate, they can

[6] This paper is a joint test of whether investors use all the available forecast properties to estimate the forecast accuracy and whether they incorporate the accurate forecasts of future earnings into their reaction to current earnings news.

significantly resolve the uncertainty towards future earnings and help investors under-react less to current earnings. Whereas the ex-ante inaccurate forecasts stimulate informed investors to form new expectations about future earnings and they may interpret and incorporate the forecasts in diverse ways (Verrecchia 2001). Hence issuing bundled management forecasts is not necessarily associated with the reduction of post-earnings announcement drift.

In a similar vein, Zhang (2008) examines the effect of analyst responsiveness on post-earnings announcement drift. She defines a responsive analyst as one who revises the forecast of next quarter's earnings within two trading days after the current quarter earnings announcement. She finds that analyst responsiveness accelerates market reaction to earnings announcements, and mitigates post-earnings announcement drift. She also finds that when earnings announcements are bundled with conference calls or management forecasts, analysts are more likely to be responsive. Compared with Zhang (2008), this paper focuses directly on the bundled management forecasts of next quarter's earnings, which are the primary information sources, and incorporates their ex-ante accuracy into the analysis.

A concurrent study by Li and Tse (2008) finds that bundled management forecasts in general mitigate post-earnings announcement drift, regardless of the accuracy of the forecasts. This finding raises a puzzling issue of why managers do not always issue optimistic forecasts along with bad earnings news to avoid the under-valuation of their firms' shares. They use the prior forecast accuracy as the proxy for ex-ante forecast accuracy, but fail to find the significant

difference in the drift between earnings announcements bundled with accurate forecasts and inaccurate forecasts.[7] I argue that the ability of bundled management forecasts to mitigate post-earnings announcement drift critically hinges on their ex-ante forecast accuracy.

2.2. The Demand of Debt Market for Conservatism

Conservatism is defined as the differential verifiability required for recognition of profits versus losses (Watts 2003). As an important accounting principle, conservatism requires a higher degree of verification to recognize gains than losses. Timely loss recognition means bad news is more timely incorporated in financial statements than good news (Basu 1997). Besides debt contracting and corporate governance, taxation and accounting regulation are also cited as explanations for the pervasive existence of conservatism in accounting standards and accounting practices. Accounting conservatism is manifested into unconditional and conditional conservatism and the difference between the two is important to understand the role of conservatism in contracting (Ball and Shivakumar 2005). Unconditional conservatism is news independent and always leads to lower earnings and book values, such as accelerated depreciation method for fixed assets and expensing of R&D expenditures. Even though the economic depreciation of fixed assets is low and R&D investments will bring profits for the future, unconditionally conservative accounting accelerates depreciation in the initial years and expenses all R&D

[7] Another contemporaneous paper by Wang (2008) also finds that management forecasts mitigate post-earnings announcement drift regardless of their ex-ante accuracy.

expenditures. In contrast, conditional conservatism is news dependent and requires higher verification for recognizing gains than recognizing losses. For example, when the expected cash flows generated by fixed assts are lower than their book value and inventories' market value is lower than their book value, assets are written down and losses are recognized. While when the market value of the assets is higher than their book value, gains are not recognized unless they are sold.

Conditional conservatism is more preferable for the efficiency of debt and compensation contracts as managers have incentives to report upward-biased accounting numbers (Beaver and Ryan 2005). In the debt contracts, covenants based on accounting ratios tend to be violated more quickly under conservative accounting, therefore lenders' rights are better protected. Unconditional conservatism, in contrast, just provides downward biased accounting numbers, the effect of which is 'at best neutral' for contracting efficiency since the ambiguous bias will extenuate the quality of financial information (Ball and Shivakumar 2005).

Although conditional conservatism increases contracting efficiency and protects lenders' rights, it can be costly to firms which adopt conservative accounting methods. Firstly, it requires more professional judgement to verify when gains can be recognized and to recognize losses on a timely basis. Secondly, conditional conservatism brings down current earnings and covenants are more likely violated under adverse situations. Zhang (2008) finds that conservatism leads to accelerated covenants violations.

The literature provides evidence that firms choosing conservative reporting benefit from lower debt costs. Using both market-based and accrual-based measures of conservatism, Ahmed et al. (2002) find that accounting conservatism decreases debt costs after controlling for other determinants of firms' debt costs but they do not differentiate between conditional and unconditional conservatism. They also find that accounting conservatism could extenuate bondholder-shareholder conflicts over dividend policy. Zhang (2008) analyzes ex-ante and ex-post benefits of conservatism to borrowers and lenders. Ex-ante conservatism benefits borrowers through lower initial interest rates and *ex post* conservatism benefits lenders through timely signal of default risk. While in the competitive capital market, borrowers will finally benefit from the gains generated by contracting efficiency.

Compared with the equity market participants, debt market participants are likely to demand more conservative financial reporting. Ball et al. (2008) argue that debtholders rely exclusively on financial statements while shareholders can also get the information from non financial disclosures. Debt contracts include covenants based on financial statement numbers and the covenants are sensitive only to losses rather than gains. The asymmetry of debt covenants leads to more demand for asymmetrically timely loss recognition instead of timely recognition of gains. By examining the relationship between country-level conservatism and debt/equity market size, they find debt market, not equity market, drives the existence of conditional conservatism. Easton et al. (2009) demonstrate that for debtholders upside benefits are fixed but downside

losses are heavy, hence investment decisions of debtholders are influenced mainly by losses rather than profits.

Even though debt contracts include covenants which may adjust financial information, reporting conservatism is not replaced. Nikolaev (2010) investigates whether debt covenants are complements or substitutes of accounting conservatism. He provides evidence that firms with more restrictive covenants in debt contracts recognize losses in a more timely manner and increase reporting conservatism more significantly after bond issuance. Beatty et al. (2008) show that contractual modifications cannot fulfill lenders' demand for conservatism. The residual conservatism, which is unexplained by litigation, tax and equity demand, is positively related with the agency costs of debt.

Since conservatism could bring direct benefits to firms by lowering interest rates, firms would increase their reporting conservatism before their bond IPOs. This hypothesis is inspired by the papers which find earnings quality is determined by the demand of the investors. Ball and Shivakumar (2005) provide evidence that UK public firms report more conservatively than private firms, which is ascribed to different information demands of public and private investors. Public investors rely on arm's length monitoring and require more conservative financial statements to solve information asymmetry. Katz (2009) uses a sample of US firms and finds firms are more conservative in their public stage than in private stage. In a related study, Ball and Shivakumar (2008) show that UK private firms improve their earnings quality before the IPO of common stock to meet the increased reporting demands of equity market. If

conservatism is more closely related with debt market, firms are expected to increase their reporting conservatism before bond IPO, even though their common stock is publicly traded already.

This paper aims to provide evidence about the incremental demand of debt market over equity market for an important aspect of financial reporting quality, namely conditional conservatism. If earnings quality is driven more by the demand of bond market and firms anticipate the demand and the benefits stated above, we should see firms become more conservative preceding the issuance of public bonds, especially before bond IPOs.

This paper focuses on public bond market rather than private debt market because financial statements play a more crucial role in the contracting of public bond market. Lenders in private debt market such as banks and other financial institutions have inside information access whereas public bondholders rely exclusively on financial statements to evaluate the firms. Bond initial public offering is a significant change in firms' financing policy and alters public-private debt mix (Datta et al 2000). Diamond (1991) argues that before firms enter the public debt market, they need to build reputation through bank monitoring. Newberry and Parthasarathy (2008) argue because accounting restatements will decrease firms' reporting credibility and increase information uncertainty, firms suffer from higher yield spreads in public debt market and rely more on private debt in post- restatement periods. Besides meeting the information demand of bond market and decreasing interest costs, firms will promote earnings quality as part of reputation building preceding

bond IPO to decrease adverse selection.

As mentioned above, debt market has asymmetric demands and reactions to reported profits and losses. Jiang (2008) finds that firms could reduce the cost of debt to the largest extent by beating the profit benchmark rather than by reporting increasing earnings or beating consensus analyst forecast. Easton et al. (2009) provide the evidence that reported losses are significantly associated with bond returns while reported profits have no effect on bond returns. In addition, the association between losses and bond returns is significantly more positive for speculative- grade bonds than for investment-grade bonds. If firms cater for the bond market, they are expected to adjust their loss reporting behaviour before bond IPO. Complementary to mandatory reporting, voluntary disclosure behaviour of these firms is analyzed as well.

2.3. Credit Pricing and Management Forecasts

Extensive research in equity markets has found that management forecasts decrease information asymmetry between a firm's managers and its investors. For instance, studies show that management forecasts lower bid–ask spreads on equity prices (Coller and Yohn 1997), increase analyst coverage (e.g., Ajinkya et al. 2005; Graham et al. 2005), generate revisions in analysts' forecasts (e.g., Jennings 1987; Clement et al. 2003; Cotter et al. 2006), attract more transient investors (Bushee and Noe 2000), and, depending on their attributes, significantly influence stock prices (e.g., Penman 1980; Hutton et al. 2003; Ng et al. 2009). However, this evidence from equity markets does not

necessarily imply that management forecasts have information relevance to credit markets, or that credit markets will necessarily respond to news in management forecasts.

Significant differences exist in the information needs of credit markets and equity markets, and it is not obvious that information that is relevant to equity holders is necessarily relevant to debt holders. As holders of a call option on the firm value, equity holders are more interested in a firm's upside potential than in its downside risks. In contrast, credit markets are concerned primarily with a firm's downside risks. Moreover, since managers have incentives to issue forecasts strategically to cater to shareholders, or to avoid adversely affecting their own compensation and reputation, they are likely to voluntarily disclose favorable news, while withholding disclosure of bad news (e.g., Roychowdhury and Sletten 2009). Delayed disclosures of bad news could give managers time to take corrective actions, if possible, as well as allow for subsequently received good news to offset the bad news. Such strategic disclosures are likely to be uninformative to credit market participants, given their greater need for bad news.

Also, credit prices are sensitive to firm-specific information, particularly when firms are risky or are close to financial distress (e.g., Easton et al. 2009). However, for such firms, Hayn (1995) finds that accounting earnings do not reflect the liquidation option value, lowering the relevance of earnings for firm valuation. Similarly, Barth et al. (1998) find that the accounting earnings becomes less important for equity valuation as the financial health of the firm

deteriorates. In addition, for firms in poorer financial health, Koch (2002) and Rogers and Stocken (2005) document that management earnings forecasts are more optimistically biased and less credible. Finally, Waymire (1985) finds that management forecasts are issued more often by less risky firms. But, for these firms, credit prices are not particularly sensitive to earnings news.

Notwithstanding the above, credit markets could still respond significantly to news in management forecasts if credit investors perceive these forecasts to be a credible and timely source of information on a firm's default risks. Consistent with this view, Kasznik and Lev (1995) and Hutton and Stocken (2009), among others, report that firms with bad news are more likely to issue management forecasts than firms with good news. The more frequent disclosure of bad news through management earnings forecasts could potentially make these forecasts highly relevant to credit pricing.

In sum, these arguments indicate that the issue of whether management forecasts are relevant for credit pricing is ultimately an empirical one. However, no empirical evidence exists on this issue. This paper aims to fill this void.

The evidence on the relative informativeness of management forecasts compared to accounting earnings for stock markets is mixed. For instance, Atiase et al. (2005) find that current earnings are more strongly associated with announcement period returns than the concurrently disclosed future earnings guidance presumably because of investors' preference for the reliability of earnings compared to management forecasts. On the other hand, recent studies document that management forecasts are more informative for stock markets

than earnings announcements (e.g., Ball and Shivakumar 2008; Beyer et al. 2009). Ball and Shivakumar (2008) attribute this higher relative importance of management forecasts to the discretion available to managers to issue forecasts only when they are perceived to be informative, and to the fact that mandated earnings have low frequency (quarterly), are not discretionary (announced every quarter regardless of arrival of new information), and are primarily backward-looking. Moreover, mandated earnings numbers face greater regulatory scrutiny and auditing requirements, which make earnings announcements relatively rigid for use in communicating timely information to capital markets. These same arguments suggest that, if any, management forecasts are likely to be more informative for credit markets than earnings announcements. However, relative to stocks, debt securities rely more heavily on reported accounting numbers in their contracts, such as in debt covenants, performance-pricing features, etc. The settlement of these debt contracts occurs at earnings announcements, and not when management forecasts are released. Consequently, earnings announcements could be more informative for credit markets than for stock markets. Hence, even if management forecasts are informative by themselves, it is unclear whether they are more or less informative than earnings announcements for credit market participants.

Several studies examine the time variation in stock market responses to new information, depending on, among other things, the level of uncertainty in the market, and investor sentiment. Veronesi (1999) theoretically analyses this issue in the context of a regime-switching economy, in which investors are

uncertain about the overall state of the economy. In his model, differences in responses to good and bad news arise endogenously as the net effect of a change in uncertainty on discount rates and of a direct effect of the news on firm value. During periods of relatively low (high) uncertainty, where investors place a high probability on the economy being in a good (bad) state, bad (good) news increases uncertainty about the state of the economy, and leads to higher discount rates. This discount rate effect, combined with the direct effect of the news, causes an asymmetrically greater (lesser) market reaction to bad (good) news in good (bad) times. Thus the model predicts asymmetric responses to bad and good news to be attenuated during periods of high uncertainty relative to periods of low uncertainty.

Early empirical studies on the time-varying stock market response to earnings news generally find that the market reaction to earnings news is larger during periods characterized by greater uncertainty (e.g., Lang 1991). However, more recent studies that allow for the market response to vary across good and bad earnings news (e.g., Conrad et al. 2002) document results consistent with the predictions of Veronesi (1999). For instance, Conrad et al. (2002) report that the difference in stock market reactions to good and bad earnings news decreases during periods of declining aggregate-market valuation and greater uncertainties. These findings, extrapolated to the credit markets, suggest that credit prices should be more (less) sensitive to good (bad) news in management forecasts during periods of greater uncertainty and declining aggregate market

value, as was the case during the recent credit crisis.[8]

In contrast to the above, Mian and Sankaraguruswamy (2008) focus on the role of investor sentiment in affecting market responses to news, and document that investors react more negatively (positively) to bad (good) earnings news during periods of high investor sentiment than during periods of low sentiment. To the extent that the recent credit crisis is characterized as a period of low investor sentiment, the findings of Mian and Sankaraguruswamy (2008) imply that credit market responses to good earnings news should be lower and credit market responses to bad news should be greater in the crisis period than in the pre-crisis period.[9]

Prior studies provide good reasons to expect that the credit-related information content of management forecasts is a function of a variety of firm characteristics and forecast attributes. First, consistent with arguments based on asymmetric payoffs facing debtholders, Easton et al. (2009) and Callen et al. (2009) show that credit markets react primarily to bad news rather than good earnings news, and that the reactions are stronger for firms with below investment-grade ratings. These findings have a direct implication for the market response to management forecasts, and imply that market responses would be larger for bad forecast news and for riskier firms. Second, forecasts that are issued sporadically by the management are likely to be viewed as being

[8] Johnson (1999) examines time variation in earnings persistence and earnings response coefficients (ERCs) on account of business cycles, and reports that both earnings persistence and ERCs are weaker during recessions relative to economic expansions. The credit crisis period in our sample includes both an economic expansion and an economic recession.

[9] Consistent with the characterization of the credit crisis as a low investor sentiment period, the Michigan Consumer Index declined from a high of 96.9 in January 2007 to a low of 55.3 by November 2008.

less credible by market participants, as the costs of a firm manipulating such forecasts are lower than for forecasts issued regularly by firms (Stocken 2000). Lastly, Rogers and Stocken (2005), among others, observe that the forecast horizon is a good indicator of the quality of the information underlying the forecasts, because managers are likely to be better informed when making forecasts with shorter horizons, and also because these forecasts are more quickly verified at the subsequent earnings announcements. Hence we expect stronger credit market reactions to short-horizon forecasts, defined as forecasts issued after the fiscal period end but before the earnings announcements (i.e., earnings pre-announcements).

Several studies have investigated the issue of spillovers in earnings news in the context of stock markets (e.g., Foster 1981; Clinch and Sinclair 1987; Tookes 2008), and document a positive correlation between a firm's unexpected earnings and the stock returns of its industry peers on the earnings release date. Similarly, another stream of research has documented positive intra-industry information transfers in the context of management earnings forecasts in the equity market (Baginski 1987; Han et al. 1989). However, little empirical evidence exists on information spillovers generated by management earnings forecasts in the credit market.

A firm's earnings news has the potential to have information content about the default risks faced by an entire industry.[10] This possibility arises for several reasons. First, Giesecke (2004) theoretically documents that credit

[10] Anilowski, Feng and Skinner (2007) study information spillovers from a firm's management forecast to the aggregate stock market.

spreads will display contagion effects when information about the borrower's fundamentals is not publicly known. Second, prior empirical studies document a significant clustering in the filing of Chapter 11 bankruptcies at the industry level (e.g., Lang and Stulz 1992; Jorion and Zhang 2007). This view is also supported by anecdotal evidence from the airline industry, where 22 airlines filed for bankruptcy in the US court in the three-year period between 1979 and 1982, compared with only two airlines in the subsequent 18-year period. Hence, to the extent that management forecasts are informative for pricing the forecasting firm's credit, we could also expect this information to be relevant for pricing credit in the entire industry to which the forecasting firm belongs.

3. The Effect of Ex-Ante Management Forecast Accuracy on Post-Earnings Announcement Drift

3.1. Introduction

This chapter investigates the effect of bundled management forecasts and their ex-ante accuracy on post-earnings announcement drift. Investors are expected to use all available information to estimate the accuracy of the management forecasts. This paper constructs a forecast accuracy prediction model based on the relation between actual forecast accuracy and the forecast properties. The results suggest that the management forecast is more accurate if the prior forecast accuracy is higher, if the forecast horizon is shorter, if the forecast difficulty is lower, if the forecast news is less extreme, if the prior stock return is higher or if the firm's market to book ratio is higher. For each

management forecast, the ex-ante forecast accuracy is measured using the estimated relationship and the current forecast properties.[11] Compared with other ex-ante accuracy measures used in the prior literature (the prior forecast accuracy, the previous four quarters' average forecast accuracy), this estimated ex-ante forecast accuracy measure is more significantly associated with the actual forecast accuracy and is a better proxy for investors' perceived accuracy.

From the second quarter of 1997 to the second quarter of 2007, 10,521 quarterly earnings announcements are bundled with the management forecasts of next quarter's earnings in the sample of 68,569 quarterly earnings announcements made by US firms.[12] 6,237 bundled management forecasts are perceived to be accurate (i.e., the estimated accuracy is higher than the median). Firms which have earnings announcements bundled with accurate management forecasts are significantly larger, have smaller analyst forecast dispersion, more analysts following, higher institutional shareholding, higher trading volume during the past, higher stock price, lower earnings persistence, less negative earnings surprises, and more responsive analysts than firms which have earnings announcements bundled with inaccurate management forecasts.

The main empirical results suggest that the bundled management forecasts reduce the magnitude of post-earnings announcement drift only when they have high ex-ante forecast accuracy, with or without controlling other

[11] To avoid the look-ahead bias, the estimated relationship is based on the data available within four quarters before the management forecast is announced.

[12] The main analysis focuses on the bundled management forecasts that are about the next quarter's earnings. Compared with the bundled management forecasts of two-quarter-ahead (and beyond) earnings, these forecasts are most likely to mitigate investors' under-reaction to current earnings.

drift-related variables. The inferences hold after a battery of robustness checks. The additional analysis shows that when the forecasts of next quarter's earnings are issued between the earnings announcements, a large part of the drift is concentrated around the management forecast announcement dates, especially for the ex-ante accurate forecasts.

The rest of the chapter proceeds as follows. Section 3.2 discusses the sample selection. Section 3.3 constructs the ex-ante forecast accuracy measure. Section 3.4 provides the research design and descriptive statistics. Section 3.5 presents the empirical results. Section 3.6 concludes.

3.2. Sample Selection

This paper obtains management forecasts announced from year 1995 to year 2007 from First Call's Company Issued Guidance database. In the analysis of the association between actual forecast accuracy and forecast properties, only point estimates and range estimates are included, because the explicit forecast numbers are necessary for the calculation of forecast accuracy. Managers might pre-announce earnings after corresponding fiscal period ends, and these pre-announcements are excluded from the management forecast sample.[13] Because management forecast error is deflated by the average stock price one week before the issuance of the management forecast, observations that do not have corresponding stock prices, or which have stock prices smaller than $1, are deleted. Financial firms, and firms whose shares are not traded on

[13] The untabulated analysis shows that the actual accuracy of earnings preannouncements is significantly higher than that of ordinary management forecasts.

NYSE/AMEX/NASDAQ or which have non-ordinary shares, are excluded from the sample as well.

The firms' actual earnings per share (EPS) over the period 1997–2007 are obtained from the I/B/E/S database.[14] The earnings announcement dates in I/B/E/S are cross-checked with those in First Call, and observations that have different earnings announcement dates are deleted.[15] Analyst forecasts are non-split adjusted forecasts from I/B/E/S. The most recent consensus analyst forecast used to calculate the unexpected earnings is the median of the analyst forecasts issued within 90 days before quarterly earnings announcements.[16]

(Insert Figure 1)

Figure 1 shows the timeline for measurement of variables. Consistent with the prior literature (Livnat and Mendenhall 2006), the drift window starts two trading days after the earnings announcement date of Quarter *t* and ends on the first trading day after the earnings announcement date of Quarter *t* + 1. The earnings announcement window is from trading day −1 to trading day 1 around the earnings announcement date (trading day 0) of Quarter *t*. The main analysis focuses on the management forecasts of next quarter's earnings issued within the earnings announcement window, and investigates their ex-ante accuracy on

[14] This time range is selected because ex-ante forecast accuracy can be estimated during this period. Specifically, it is from the second quarter of 1997 to the second quarter of 2007.

[15] This paper uses the earnings numbers in I/B/E/S rather than COMPUSTAT because earnings in COMPUSTAT are restated: therefore they are not the actual earnings that investors observe and react to. See Livnat and Mendenhall (2006) for a comparison between I/B/E/S and COMPUSTAT.

[16] The results are similar if I use the latest individual analyst forecast to calculate the unexpected earnings, following Zhang (2008).

the drift period return.[17]

3.3. Measurement of Ex-ante Management Forecast Accuracy

I propose that investors estimate the ex-ante management forecast accuracy based on a variety of forecast properties. Rogers and Stocken (2005) develops a management forecast error prediction model which identifies the following explanatory variables: forecast horizon, forecast difficulty, forecast news, previous stock return, litigation risk, industry concentration, financial distress, market value to book value of equity ratio and firm size.[18] In addition, I argue that the accuracy of firms' prior forecasts is a proxy for firm-level management forecast ability and affects investors' perceived accuracy of current forecast (Hutton and Stocken 2009). This paper also takes into account the different accuracy levels of bundled management forecasts and non-bundled management forecasts (Rogers and Van Buskirk 2009). The following model is used to estimate the association between the actual management forecast accuracy and the forecast properties.

$$ActualAccuracy_{i,t} = \alpha_0 + \alpha_1 PriorAccuracy_{i,t} + \alpha_2 ForecastHorizon_{i,t} + \alpha_3 ForecastDifficulty_{i,t} + \alpha_4 ForecastNews_{i,t} \times GoodNews_{i,t} + \alpha_5 ForecastNews_{i,t} \times BadNews_{i,t} + \alpha_6 CAR_{i,t} + \alpha_7 Litigation_{i,t} + \alpha_8 Concentration_{i,t-1} + \alpha_9 Distress_{i,t-1} + \alpha_{10} MB_{i,t-1} + \alpha_{11} Size_{i,t-1} + \alpha_{12} Bundle_{i,t} + \varepsilon_{i,t} \quad (1)$$

[17] Annual forecasts which have the next quarter ends as the forecast period end dates are also included.

[18] The prediction model in Rogers and Stocken (2005), specifically, is about forecast bias. This paper focuses on forecast accuracy, which is subject to the effects of intentional forecast bias, forecast ability, and forecast difficulty.

The variables are defined below:

Actual forecast accuracy (*ActualAccuracy*): Following Ng et al. (2008), I calculate management forecast accuracy as *ActualAccuracy* = -1×*ABS* (*ActualEarnings-ManagementForecast*)*/Price. Price* refers to the average stock price one week before the management forecast announcement date. Management forecast is more accurate if *ActualAccuracy* is higher (closer to zero).

Prior forecast accuracy (*PriorAccuracy*) refers to the actual accuracy of the prior management forecast. The actual earnings in relation to the prior forecast need to be announced on or immediately before the current management forecast announcement date.

Forecast horizon (*ForecastHorizon*) is calculated as the management forecast period end date minus the announcement date, deflated by 360.

Forecast difficulty (*ForecastDifficulty*) is developed by performing Principal Axis Factoring over analyst forecast dispersion, standard deviation of previous analyst forecast errors, firm's prior performance, future performance forecasted by the management, stock return volatility, bid-ask spread and forecast width (see Rogers and Stocken 2005).

Forecast news (*ForecastNews*) is calculated as the management forecast minus the consensus analyst forecast, deflated by *Price*. *GoodNews* is an indicator variable which equals 1 if *ForecastNews* is non-negative, and 0 otherwise. *BadNews* is an indicator variable which equals 1 if *ForecastNews* is negative, and 0 otherwise.

Other control variables: *CAR* is the firm's abnormal stock return cumulated from day -120 to day -1 relative to the management forecast announcement day. *Litigation* is an indicator variable which equals 1 if the firm is in industries with high litigation risk (Standard Industrial Classification codes 2833-2836, 3570-3577, 3600-3674, 5200-5961, 7370-7374, 8731-8734), and 0 otherwise. *Concentration* is measured by the Herfindahl index using the revenues of firms sharing the same four-digit SIC code. *Distress* is an indicator variable which equals 1 if the firm is in the most distressed decile predicted by Ohlson's (1980) bankruptcy model, and 0 otherwise. *MB* is the firm's market value of equity deflated by the book value of equity. *Size* is the natural log of the firm's total assets. *Concentration, Distress*, *MB* and *Size* are calculated using the most recent accounting data prior to the announcement of the management forecast. *Bundle* is an indicator variable which equals 1 if the management forecast is issued within one trading day around the earnings announcement date and 0 otherwise.

(Insert Table 1)

Model (1) is estimated separately for 20,701 annual forecasts and 21,790 quarterly forecasts issued from year 1995 to year 2007. The results are presented in Panel A of Table 1. Actual forecast accuracy is positively related with the prior management forecast accuracy. Forecast accuracy is also higher when forecast horizon is shorter. The significantly negative coefficient on *ForecastDifficulty* suggests that this latent variable effectively captures the difficulty for managers to accurately forecast future earnings. The coefficients

on $ForecastNews \times GoodNews$ and $ForecastNews \times BadNews$ indicate that forecast accuracy is negatively associated with the magnitude of the management forecast news. Previous stock return is positively related with forecast accuracy, suggesting managers tend to forecast more accurately if they are under less share price pressure. The positive coefficient on *MB* indicates that firms with good growth opportunities are likely to issue accurate forecasts to build forecast reputation on capital market. The actual accuracy of the annual forecast is higher when the forecast is bundled with the earnings announcement.

Using the above prediction model, I build the ex-ante forecast accuracy measure by applying the estimated coefficients to the current forecast properties. For every consecutive four calendar quarters, the actual management forecast accuracy and forecast properties are used to estimate the coefficients of Model (1). To avoid the look-ahead bias, these estimated coefficients are used in the following calendar quarter to obtain the ex-ante accuracy of the management forecasts issued within this quarter. For example, for quarterly earnings announcements made from the first quarter of 1999 to the fourth quarter of 1999, all the management forecasts in relation to these earnings are obtained to estimate Model (1). The estimated coefficients are then used to compute the ex-ante accuracy of quarterly forecasts announced in the first quarter of 2000. In this way, the ex-ante accuracy of 20,148 annual forecasts and 20,929 quarterly forecasts issued from the second quarter of 1997 to the second quarter

of 2007 is estimated.[19] Previous literature has used other ex-ante accuracy measures. Ng et al. (2008) use the actual accuracy of the prior management forecast, and Li and Tse (2008) use the previous four quarters' average management forecast accuracy. The association between the actual forecast accuracy (*ActualAccuracy*) and the multiple ex-ante accuracy measures, including the estimated accuracy (*EstimatedAccuracy*) developed by Model (1), the prior forecast accuracy (*PriorAccuracy*), the previous four quarters' average forecast accuracy (*AverageAccuracy*), is shown in Panel B of Table 1. *EstimatedAccuracy* has the highest association with the actual forecast accuracy. The coefficient of the univariate regression is 0.752, and the adjusted R^2 is 24.14%.

In the following analyses, the *EstimatedAccuracy* values are transformed into indicator variables based on the previous four quarters' cut-offs. The indicator variable *Dummy_EstimatedAccuracy* equals 1 if *EstimatedAccuracy* is greater than the median and 0 otherwise. Similarly *Dummy_PriorAccuracy* equals 1 if *PriorAccuracy* is greater than the median and 0 otherwise. *Dummy_AverageAccuracy* equals 1 if *AverageAccuracy* is greater than the median and 0 otherwise.

3.4. Research Design and Descriptive Statistics

3.4.1. Test of the Effect of Ex-ante Forecast Accuracy on Post-Earnings

[19] The ex-ante accuracy of forecasts announced before the second quarter of 1997 cannot be estimated because there are no enough observations to generate meaningful coefficients for the prediction model.

Announcement Drift

The effect of ex-ante management forecast accuracy on post-earnings announcement drift is examined by estimating the following model:

$$
\begin{aligned}
ADJ\_RET_{i,t} = {} & \beta_0 + \beta_1 DSUE_{i,t} + \beta_2 D\_BUNDLE_{i,t} + \beta_3 DSUE_{i,t} \times D\_BUNDLE_{i,t} \\
& + \beta_4 DSUE_{i,t} \times D\_BUNDLE_{i,t} \times Accuracy_{i,t} + \beta_5 DSUE_{i,t} \times DFD_{i,t} \\
& + \beta_6 DSUE_{i,t} \times DME_{i,t} + \beta_7 DSUE_{i,t} \times DAC_{i,t} + \beta_8 DSUE_{i,t} \times DVOL_{i,t} \\
& + \beta_9 DSUE_{i,t} \times DPRC_{i,t} + \beta_{10} DSUE_{i,t} \times DINS_{i,t} + \beta_{11} DSUE_{i,t} \times DEP_{i,t} \\
& + \beta_{12} DSUE_{i,t} \times BADNEWS_{i,t} + \beta_{13} DSUE_{i,t} \times 4THQTR_{i,t} + \beta_{14} DSUE_{i,t} \\
& \times RESPONSIVE_{i,t} + \beta_{15} DFD_{i,t} + \beta_{16} DME_{i,t} + \beta_{17} DAC_{i,t} + \beta_{18} DVOL_{i,t} \\
& + \beta_{19} DPRC_{i,t} + \beta_{20} DINS_{i,t} + \beta_{21} DEP_{i,t} + \beta_{22} BADNEWS_{i,t} \\
& + \beta_{23} 4THQTR_{i,t} + \beta_{24} RESPONSIVE_{i,t} + \varepsilon_{i,t} \qquad (2)
\end{aligned}
$$

Where *ADJ_RET* is the size-adjusted return over the drift window and equals the compounded raw return minus the compounded benchmark return of the same CRSP size decile and the same CRSP exchange index (NYSE/AMEX or NASDAQ) that the firm belongs to. Following Shumway and Warther (1999), when a firm is delisted due to poor performance (delisting code is 500 or from 520 to 584), the delisting return is assumed to be −35% if it is traded on NYSE/AMEX, and −55% if it is traded on NASDAQ.[20]

DSUE refers to the decile rank of earnings surprise, which is defined as the actual EPS minus the most recent consensus analyst forecast, scaled by the stock price at the end of the fiscal quarter. The earnings surprises are ranked

[20] The percentage of firms delisted due to poor performance is 0.04% for the whole sample. The results are not affected if I simply delete these observations.

into deciles within each calendar quarter using the cut-off values from the previous quarter, coded from 0 to 1 to yield *DSUE*. The coefficient of *DSUE* can be interpreted as the abnormal return earned on a zero-investment portfolio that takes a long position in the highest *DSUE* decile (*DSUE* = 1) and a short position in the lowest *DSUE* decile (*DSUE* = 0).[21]

When an earnings announcement is bundled with the management forecast of next quarter's earnings, the indicator variable *D_BUNDLE* equals 1, and 0 otherwise. Bundled management forecasts are expected to mitigate post-earnings announcement drift when their ex-ante forecast accuracy (*Accuracy*) is high. Hence $\beta_3 + \beta_4$ should be significantly negative. β_4 should be significantly negative as well under the prediction that accurate bundled forecasts and inaccurate bundled forecasts should have different mitigating effects on post-earnings announcement drift.

The above equation includes several control variables that prior studies have identified as being associated with post-earnings announcement drift. These are analyst forecast dispersion (*DFD*), firm size (*DME*), analyst coverage (*DAC*), trading volume (*DVOL*), price (*DPRC*), institutional shareholding (*DINS*), earnings persistence (*DEP*), negative unexpected earnings (*BADNEWS*), the fourth fiscal quarter earnings announcement (*4THQTR*) and analyst responsiveness (*RESPONSIVE*).[22] The first three variables are proxies for information uncertainty. Analyst forecast dispersion is defined as the

[21] See Doyle et al. (2006) and Zhang (2008).
[22] See Bartov et al. (2000), Mendenhall (2002), Rangan and Sloan (1998), Zhang (2006) and Zhang (2008), among others.

standard deviation of analyst forecasts divided by the stock price at the fiscal quarter end. Firm size is measured by the market value of equity at the fiscal quarter end. Analyst coverage is the number of analysts following the firm based on the analyst estimates from I/B/E/S. Institutional shareholding is the percentage of institutional ownership which is available from CDA/Spectrum. Following Mendenhall (2002), trading volume is estimated by multiplying the closing price and the shares traded from day −272 to day −21 relative to the earnings announcement day, and earnings persistence is the first-order serial correlation of seasonally-differenced earnings estimated over the past 20 quarters. Stock price is the average stock price within one week before the earnings announcement. These drift-related variables are transformed into decile ranks within each calendar quarter using the cut-off values from the previous quarter and coded from 0 to 1(Mendenhall, 2002). In this test, I use the inverse of the standard deviation of analyst forecasts to form *DFD*. Hence, consistent with *DME* and *DAC*, a higher value of *DFD* corresponds to lower analyst forecast dispersion and lower information uncertainty. *BADNEWS* is an indicator variable which equals 1 if the unexpected earnings are negative and 0 otherwise. *4THQTR* is an indicator variable which equals 1 if the earnings announcement is for the fourth fiscal quarter and 0 otherwise. Following Zhang (2008), *RESPONSIVE* equals 1 if there is at least one analyst revising the forecast of next quarter's earnings within two trading days after current quarter earnings announcement and 0 otherwise.

3.4.2. Descriptive Statistics

(Insert Table 2)

Panel A and Panel B of Table 2 present the means and medians of the drift-related variables conditional on the existence of the bundled management forecasts and the ex-ante accuracy of the bundled forecasts. T-tests for means and Wilcoxon-tests for medians are conducted over three sub-samples: earnings announcements bundled with accurate forecasts (*D_BUNDLE* = 1 and *Dummy_EstimatedAccuracy* = 1), earnings announcements bundled with inaccurate forecasts (*D_BUNDLE* = 1 and *Dummy_EstimatedAccuracy* = 0), and standalone earnings announcements (*D_BUNDLE* = 0). From the second quarter of 1997 to the second quarter of 2007, among 68,569 quarterly earnings announcements, 10,521 quarterly earnings announcements are bundled with the management forecasts of next quarter's earnings.[23] 6,237 bundled management forecasts are perceived to be accurate.

Based on the t-tests for mean values, firms which have earnings announcements bundled with accurate management forecasts are significantly larger, have smaller analyst forecast dispersion, more analysts following, higher institutional shareholding, higher trading volume during the past, higher stock price, lower earnings persistence, less negative earnings surprises, and more responsive analysts than firms which have earnings announcements bundled with inaccurate management forecasts. Firms which have earnings

[23] To achieve a clean test, observations which have management forecasts of Quarter t + 1 earnings issued before the earnings announcement date of Quarter t or between the earnings announcement date of Quarter t and the earnings announcement date of Quarter t + 1 are deleted.

announcements bundled with inaccurate management forecasts are not significantly different from firms which have standalone earnings announcements in terms of analyst forecast dispersion, but have better information environment in terms of analyst coverage, institutional shareholding and analyst responsiveness. The former firms are significantly smaller in market value of equity than the latter firms based on the mean values, while the median values suggest the opposite. The different characteristics of these three kinds of firms justify the necessity of controlling the drift-related variables.

(Insert Table 3)

Panel A of Table 3 shows the percentage of earnings announcements with bundled management forecasts and with accurate bundled management forecasts by industry. There are 11 industry categories based on the industry classification in Fama and French (1997) and the finance industry (SIC codes 6000-6999) is excluded. Firms in wholesale and retail industry have the highest percentage (19.77%) of earnings announcements with bundled management forecasts, followed by firms in business equipment industry (19.40%). Firms in chemicals and allied products industry have the highest percentage (12.48%) of earnings announcements bundled with accurate management forecasts. Whereas firms in energy industry and telecommunications industry have quite low percentage (4.04% and 7.10%) of earnings announcements bundled with management forecasts of future earnings, suggesting that firms in high technology firms are less likely to issue bundled management forecasts because

of the possible costs of not being able to issue accurate management forecasts or the proprietary costs as suggested by Verrecchia (1983).

Panel B of Table 3 shows the correlations among *DSUE*, *D_BUNDLE* and other drift-related variables used in the test. The upper-right triangle reports the Pearson product moment and the Spearman rank order is presented in the lower-left triangle. The indicator variable *D_BUNDLE* is significantly correlated with all drift-related variables. It is positively correlated with the inverse of analyst forecast dispersion (*DFD*), market value of equity (*DME*), analyst coverage (*DAC*), trading volume (*DVOL*), stock price (*DPRC*), institutional shareholding (*DINS*), earnings persistence (*DEP*) and analyst responsiveness (*RESPONSIVE*), and negatively correlated with the negative earnings surprise (*BADNEWS*) and the fourth fiscal quarter earnings announcement (*4THQTR*). This univariate analysis indicates that firms in better information environment are more likely to issue bundled management forecasts than firms in relatively uncertain information environment.

3.5. Empirical Results

3.5.1. Main Analysis

(Insert Table 4)

Table 4 shows the regression results of Model (2).[24] Model (2a) tests the effect of the bundled management forecasts on post-earnings announcement

[24] Following Zhang (2008), I delete the observations with absolute value of studentized residuals greater than 2 in all regressions to remove the effects of outliers. This estimation procedure decreases the sample size by about 4% and does not change the inferences.

drift. The average drift is 3.9% if earnings announcements are not bundled with the management forecasts of next quarter's earnings. The coefficient of the interaction term between *DSUE* and *D_BUNDLE* is −0.013, meaning that the bundled management forecasts can decrease the drift by 1.3%; however, the effect is only significant at 10% level (*t*-stat = −1.78). β_3 becomes insignificantly negative when other drift-related variables are controlled, as is shown in the estimation results of Model (2b). Therefore the bundled management forecasts in general do not significantly decrease the magnitude of post-earnings announcement drift when other drift-related variables are controlled.

Models (2c), (2d), (2e), and (2f) test the effect of the ex-ante accuracy of bundled management forecasts on post-earnings announcement drift, using multiple measures of the ex-ante accuracy. Model (2c) and (2d) use *Dummy_ EstimatedAccuracy*, the ex-ante accuracy measure developed by Model (1). When the drift-related variables are not controlled, the coefficient on the interaction term between *DSUE* and *D_BUNDLE* is -0.006 (*t*-stat = −0.80) and the coefficient on $DSUE \times D\_BUNDLE \times Dummy\_EstimatedAccuracy$ is −0.015 (*t*-stat = −2.50). This result suggests that the inaccurate bundled management forecasts have no significant effect on the drift, whereas the accurate bundled management forecasts significantly reduce the magnitude of post-earnings announcement drift by 2.1% ($\beta_3 + \beta_4$ with *p-value* of 0.006). The negative coefficient on $DSUE \times D\_BUNDLE \times Dummy\_EstimatedAccuracy$ suggests the different mitigating effects of accurate bundled management

forecasts and inaccurate bundled management forecasts on post-earnings announcement drift. When the drift-related variables are controlled in Model (2d), the inaccurate bundled management forecasts do not mitigate post-earnings announcement drift at all (β_3 = 0.003 with *t*-stat of 0.44). The accurate bundled forecasts are shown to have a significant mitigating effect on the drift with the coefficient of −0.024 ($\beta_3 + \beta_4$ with *p-value* of 0.003). Consistent with prior research, the magnitude of post-earnings announcement drift is significantly lower if firms have more analysts following, higher trading volume or when the announced earnings correspond to the fourth fiscal quarter.

When the ex-ante management forecast accuracy is measured by *Dummy_PriorAccuracy* or *Dummy_AverageAccuracy* in Model (2e) and Model (2f), β_4 and $\beta_3 + \beta_4$ become insignificant. [25] I interpret the results as the evidence that these ex-ante forecast accuracy measures fail to correctly capture investors' perceived forecast accuracy and show the different effects of accurate bundled forecasts and inaccurate bundled forecasts on post-earnings announcement drift.

In sum, the above results are consistent with the prediction that the bundled management forecasts of next quarter's earnings mitigate post-earnings announcement drift only when the forecasts have high ex-ante accuracy.

3.5.2. Endogeneity

[25] When there are no actual earnings related with the prior management forecasts announced during the previous four quarters, *AverageAccuracy* can not be calculated. 1,871 such observations are deleted from the regression.

Firms which choose to issue management forecasts of next quarter's earnings along with current quarter earnings announcements may have certain characteristics which are associated with the relation between earnings and the stock return. The main analysis does not control for the self-selection of issuing bundled management forecasts. To mitigate the issue of omitted correlated variables, I use the Heckman (1979) two-stage approach. In the first stage, I adapt the Probit model in Rogers and Van Buskirk (2009) to estimate the probability of issuing the bundled management forecast for each quarterly earnings announcement. The explanatory variables include the existence of prior management forecast for current quarter (*MF_EXIST*), the existence of bundled management forecast for the most recent quarterly earnings announcement (*LAG_DBUNDLE*), prior stock return (*PRIOR_RETURN*), the natural log of market value of equity (*LOG_ME*), the natural log of analysts following (*LOG_AC*), the proportion of earnings announcements which meet analyst expectations during the prior four quarters (*MEET*), the absolute value of earnings surprise (*ABS(SUE)*), the indicator variable (*BADNEWS*) which equals 1 if current earnings surprise is negative and 0 otherwise, the indicator variable (*LOSS*) which equals 1 if the reported earnings number is negative and 0 otherwise, and analyst forecast dispersion (*AFD*),

(Insert Table 5)

The results are presented in Table 5. Firms are more likely to issue bundled management forecasts if there is a prior management forecast for current quarter's earnings, if there is a bundled management forecast for the

most recent quarterly earnings announcement, if the prior stock return is higher, if the firm is larger, if there are more analysts following or if the firms are more likely to meet analyst estimates during the prior four quarters. The probability of issuing a bundled management forecast is lower when the absolute value of the earnings surprise is larger, when the current earnings announcement conveys bad news, when the announced earnings number is negative, or when the analyst forecast dispersion is higher. The results are largely consistent with the univariate analysis in Panel B, Table 3. The inverse Mills ratios (IMR) for all firms are calculated accordingly using the estimated coefficients. In the second stage, I include the inverse Mills ratios in Model (2) and present the results in Table 6.

(Insert Table 6)

Under all specifications, the coefficients on the inverse Mills ratios are significant, justifying the endogeneity bias. The coefficient on IMR can be interpreted as the covariance of the error terms of the self-selection model and the post-earnings announcement drift test. It is significantly negative when the drift-related variables are not controlled, and becomes significantly positive when the drift-related variables are added into the regressions. The main results are qualitatively similar with those presented in Table 4. The accurate bundled forecasts continue to significantly mitigate post-earnings announcement drift with the coefficient of -0.022 ($\beta_3+\beta_4$ with *p-value* of 0.006) when the ex-ante accuracy is measured by *Dummy_EstimatedAccuracy*.

3.5.3. Robustness Checks

The main analysis examines the effect of ex-ante management forecast accuracy on post-earnings announcement drift when earnings announcements are bundled with management forecasts of next quarter's earnings. In this section, I test the robustness of the inferences.

First, in many cases, besides the management forecasts of next quarter's earnings, management forecasts of two-quarter-ahead earnings, three-quarter-ahead earnings or four-quarter-ahead (and beyond) earnings are also issued along with current quarter earnings announcements. These management forecasts may provide incrementally valuable information about the persistence of current earnings into future earnings. To erase this confounding effect, I define *D_BUNDLE* = 1 when an earnings announcement is bundled only with the management forecast of next quarter's earnings. The inferences regarding the effect of the ex-ante accuracy of the bundled management forecasts on post-earnings announcement drift remain unchanged.

Second, the main analysis focuses on the management forecasts of next quarter's earnings issued within one trading day around the current quarter earnings announcements. I also widen the event window for the definition of bundled management forecasts, so as to include any management forecast of next quarter's earnings made in the period [-10, +1] or [-30, +1] around the current quarter earnings announcement date (day 0). Investors will incorporate these earlier management forecasts into their reaction to the current quarter earnings announcements as well. I re-estimate Model (2) by treating these

earlier management forecasts as bundled managements, the results are very similar with the ones in the main analysis.

Third, in the post-earnings announcement drift test, the sample period is from the second quarter of 1997 to the second quarter of 2007 which includes the time period before the issuance of Regulation Fair Disclosure. Some earnings announcements might be bundled with private management forecasts which may also reduce post-earnings announcement drift through affecting analyst forecasts, but this effect is not captured due to the limited coverage of First Call CIG database before the issuance of Reg FD. Therefore I use the data after the issuance of Reg FD to estimate Model (2a) and get similar results which suggest that on average the bundled management forecasts do not mitigate post-earnings announcement drift.

Finally, the inferences are unchanged if I define the drift window starting from trading day 2 after the earnings announcement date of Quarter *t* and ending on trading day -2 relative to the earnings announcement date of Quarter *t* + 1, or a 60-day drift window as in Liang (2003).

3.5.4. Clustering of the Drift

If management forecasts of next quarter's earnings accelerate investors' reaction to announced earnings, and if management forecasts of next quarter's earnings are not bundled, but issued between earnings announcements, post-earnings announcement drift should cluster around the management

forecast announcement dates.[26] Following the method in Zhang (2008), this paper estimates the coefficients of the following model:

$$ADJ\_RET_{i,t} = \gamma_0 + \gamma_1 \ DSUE_{i,t} + \varepsilon_{i,t} \quad (3)$$

(Insert Table 7)

The model is estimated, separately, with the dependent variable *ADJ_RET* being the drift period return and the management forecast announcement return, which is accumulated over trading days -1, 0, and +1 around the announcement date. As is shown in Table 7, if *ADJ_RET* refers to the drift period return, γ_1 equals 0.060. If *ADJ_RET* refers to the management forecast announcement return, γ_1 equals 0.027[27], which suggests that a large part of the drift is concentrated around the announcements of the management forecasts, and investors still under-react to last quarter's earnings before that. In addition, γ_2 equals 0.017 if investors regard the management forecasts as inaccurate (*Dummy_EstimatedAccuracy* = 0). The coefficient of *DSUE* equals 0.057 ($\gamma_1 + \gamma_2$) if the ex-ante accuracy of the management forecasts is above the median (*Dummy_EstimatedAccuracy* = 1).

3.6. Conclusion

The under-reaction explanation for post-earnings announcement drift

[26] If the earnings announcements are bundled with management forecasts which are perceived to be inaccurate and there are revised forecasts issued before next earnings announcements, these observations are included as well.

[27] In this regression γ_1 can be interpreted as the abnormal return earned over the management forecast announcement period on a zero-investment portfolio that takes a long position in the highest *DSUE* decile (*DSUE* = 1) and a short position in the lowest *DSUE* decile (*DSUE* = 0).

argues that investors do not fully understand the implications of current earnings for future earnings and this under-reaction is corrected at future earnings announcement dates. Using an estimated ex-ante forecast accuracy measure that is based on forecast ability, forecast difficulty and forecast environment, this paper finds that the bundled management forecasts of next quarter's earnings mitigate post-earnings announcement drift only when they are of high ex-ante accuracy. The results hold after modelling the self-selection of issuing bundled management forecasts and a battery of robustness checks.

The additional analysis shows that a large part of the drift clusters around the forecast announcement period when the management forecasts of next quarter's earnings are issued between the earnings announcements, especially when the forecasts are of high ex-ante accuracy, suggesting that the issuance of management forecasts of future earnings accelerates investors' reaction to announced earnings. This evidence supports the argument that post-earnings announcement drift is at least partly due to investors' failure to fully understand the implications of current earnings for future earnings.

4. Conservatism Adjustment before Bond IPO

4.1. Introduction

Conservatism plays an important role in the debt market because timely loss recognition provides more secure ex-ante financial information for debt pricing and violation of debt covenants is more quickly triggered through ex-post timely loss recognition (Ball and Shivakumar 2005). A lot of papers

provide evidence that ex-ante conservatism leads to lower debt costs (Ahmed 2002; Zhang 2008). Lenders also benefit from timely signalling of default risk via accelerated covenant violations (Zhang 2008). In the competitive market, the benefits of conservative reporting will go to borrowers finally because conservatism decreases the monitoring costs of lenders. Besides conditional conservatism, unconditional conservatism is also important in debt contracts if collateral is required. For example, if tangible assets, such as obsolete machines, are not impaired and they are used as collateral, their high book values are delusive and lenders' rights can not be protected at liquidation.

Ball et al. (2008) argue that the debt market demands information exclusively from financial reporting while equity market also needs information through non-financial disclosure. Therefore the relation between financial variables and equity returns is more crucial to debt market. Another important property of debt contracts is that the right of lenders is affected only in default and the covenants are sensitive to losses, which demands more timely loss recognition.

Conservatism could also promote corporate governance. Ball and Shivakumar (2005) argue that corporate governance is affected because timely loss recognition forces managers to avoid investments with ex-ante negative NPV and stop operations that generate negative ex-post cash flows. They examine the earnings quality in UK private and public firms and find that public firms report more conservatively. In a related study, they find firms changing from private to public status improve their reporting quality right

before the stock IPO (Ball and Shivakumar 2008).

Descriptive statistics show that the market-to-book ratio increases and the non-operating accruals become more negative before bond IPO, consistent with more conservative reporting before the bond IPO. Firms also report fewer losses before bond IPOs, catering for bond market's asymmetric reaction to losses and profits.

The results of Basu regression suggest that firms increase their reporting conservatism at event year -1, -2 before bond IPO when the reporting conservatism in years before event year -2 is used as the benchmark. The accrual-based model also shows that firms report more conservatively both at event year -1 and -2 relative to the level of reporting conservatism in years before event year -2.

This chapter is organized as follows. Section 4.2 talks about sample selection and research design. Section 4.3 presents the descriptive statistics and relates them with main empirical tests. Section 4.4 reports the main empirical results and sensitivity tests. Section 4.5 concludes the chapter.

4.2. Sample Selection and Research Design

4.2.1. Firms with Bond IPO

Table 8 shows the sample selection process. Unlike multiple databases recording the dates of common stock IPOs, there is no database containing the dates of firms' bond IPOs. So, to identify the bond IPO, I adopt the following procedures. Following the method by Hale and Santos (2008), I restrict the

firms to those that appeared in COMPUSTAT on or after 1970. Under section 13 and 15(d) of the Securities Exchange Act of 1934, a firm needs to file annual report (10K) to SEC if it has public securities. The information source of COMPUSTAT is also 10K. If a firm appears in COMPUSTAT on or after 1970, most likely this firm has not issued any public security- share or bond- before 1970. The domestic new bond issuance database in Securities Data Corporation (SDC) covers all bond issuances starting from 1970 in US market. If a firm does not have financial information in COMPUSTAT before 1970, i.e., it does not issue any bond before 1970, its first public bond issuance in SDC is treated as the firm's bond IPO.

After selecting firms appearing in COMPUSTAT on or after 1970, these firms are merged with their counterparts in CRSP to find out their historical CUSIPs and their industry classification codes. Since COMPUSTAT uses firms' current CUSIPs and SDC uses firms' historical CUSIPs as identifiers, firms' financial information in COMPUSTAT can be connected to their bond issuances in SDC only through historical CUSIPs. Because of different regulation and reporting incentives of firms in financial industry, firms which have standard industry classification codes (SICCD) between 6000 and 6999 are deleted.

Firms may have bond IPOs soon after they made an IPO of equity shares and the sample shows that 107 firms have common stock IPOs within three years before they issue bonds for the first time. As the firms' accruals in this period are likely to be affected by their common stock IPO proceeds, these

firms are excluded. In the accrual based model, to get the precise estimation of total accruals by cash-flow statement items, firms which only have data before 1987 are excluded as well, decreasing the number of firms from 769 to 588.

4.2.2. Research Design

I use the following regression, adapted from the Basu (1997) model, to examine how conditional conservatism changes before bond IPO.

$$E_{i,t}/P_{i,t-1} = \alpha_0 + \alpha_1 DR_{i,t} + \alpha_2 RET_{i,t} + \alpha_3 RET_{i,t} \times DR_{i,t} + \alpha_4 DBE_{i,t} + \alpha_5 DBE_{i,t} \times DR_{i,t} + \alpha_6 DBE_{i,t} \times RET_{i,t} + \alpha_7 DBE_{i,t} \times DR_{i,t} \times RET_{i,t} + \varepsilon_{i,t} \quad (4)$$

This model studies the asymmetric relationship between earnings and contemporaneous stock returns. $E_{i,t}$ is the earnings per share for firm i in fiscal year t, $P_{i,t-1}$ is the price per share at the end of last fiscal year. $RET_{i,t}$ is the stock return of firms i from 3 months after fiscal year end $t-1$ to 3 months after fiscal year end t. $DR_{i,t}$ is the indicator variable which equals 1 if $RET_{i,t}$ is negative and 0 otherwise. Basu (1997) argues that if firms report conservatively, they incorporate bad news (proxied by negative stock returns) into earnings more quickly than good news (proxied by positive stock returns). This reporting property reflected in the coefficients means that $\alpha_2+\alpha_3$ should be bigger than α_2, therefore α_3 which captures the incremental sensitivity of earnings to negative returns should be positive.

This paper modifies the original Basu regression to study whether firms report more conservatively before bond IPO. The time indicator variable $DBE_{i,t}$ is applied and it equals 0 in event year -1, -2 (and -3) and equals 1 before event year -2 (or -3) . Here the event year refers to the year when a firm has its

bond IPO, therefore event year -1 is one year before its bond IPO year and event year -2 is two years before its bond IPO year, etc. Roychowdhury and Watts (2007) point out that Basu measure is biased downward over short horizon and this bias becomes less severe over longer horizon (2 years or more), this paper does not compare firms' reporting conservatism in event year -1, -2 (and -3) separately with that in years before -2 (or -3). If firms do report more conservatively preceding bond IPO, the incremental sensitivity of earnings to negative returns should be less in years before -2 (or -3), therefore α_7 should be negative.

As Basu measure is biased over short horizon, reporting conservatism in event year -1, -2 and -3 can not be compared separately with that in previous years. To solve this problem, this paper follows Ball and Shivakumar (2008) and examines how conditional conservatism changes before bond IPO by studying the relationship between accruals and cash flow.

$$ACC_{i,t} = \beta_0 + \beta_1 CFO_{i,t} + \beta_2 \Delta Sales_{i,t} + \beta_3 FAsset_{i,t} + \beta_4 DCFO_{i,t} + \beta_5 DCFO_{i,t} \times CFO_{i,t} + \beta_6 DBE_{i,t} + \beta_7 DBE_{i,t} \times CFO_{i,t} + \beta_8 DBE_{i,t} \times \Delta Sales_{i,t} + \beta_9 DBE_{i,t} \times FAsset_{i,t} + \beta_{10} DBE_{i,t} \times DCFO_{i,t} + \beta_{11} DBE_{i,t} \times DCFO_{i,t} \times CFO_{i,t} + \varepsilon_{i,t} \quad (5)$$

Where $ACC_{i,t}$ is the total accruals of firm i in year t, $CFO_{i,t}$ is cash flow from operation of firm i in year t. $\Delta Sales_{i,t}$ is the change of sales from year t-1 to year t . $FAsset_{i,t}$ is the book value of fixed assets. $DCFO_{i,t}$ is the indicator variable which equals 1 if $CFO_{i,t}$ is negative and 0 otherwise. If firms recognize losses more timely than gains, $\beta_1 + \beta_5$ should be bigger than β_1,

therefore β_5, the incremental sensitivity of accruals to negative cash flow, should be positive.

To study the reporting conservatism in different time periods, same time indicator variable $DBE_{i,t}$ is applied. It equals 0 in years adjacent to bond IPO and equals 1 otherwise. In the main test, the reporting conservatism in years before event year -2 is treated as the benchmark and the conditional conservatism of firms in event year -1 or -2 is compared separately with the benchmark. If firms report more conservatively before bond IPO, the incremental sensitivity of accruals to negative cash flow will increase significantly preceding bond IPO and β_{11} is expected to be negative both in event year -1 and -2.

4.3. Descriptive Statistics

4.3.1. Bond IPO Summary

Table 9 shows the number of firms which have bond IPOs in each year from 1970 to 2005. The total number of firms is 769. Firms have their bond IPOs mostly in 1980s and 1990s.

Table 10 shows how market-to-book ratio and non-operating accruals change in the pre-bond-IPO years. Market-to-book ratio is a widely used measure of unconditional conservatism since it shows the average understatement of book value of net assets relative to the market value (Beaver and Ryan 2005). This paper does not focus on the change of unconditional conservatism but the accumulation of conditional conservatism will increase the

unconditional conservatism accordingly. The higher the market-to-book ratio, the more conservative the firm is. The comparison of mean market-to-book ratios shows that firms gradually increase their market-to-book ratios approaching bond IPO, i.e., they report more conservatively before bond IPO. The comparison of median market-to-book ratios does not show the increasing conservatism before bond IPO, however if it is combined with the result of mean market-to-book ratios, it means the upper half firms increase their market-to-book ratios significantly in event year -1 and -2.

Non-operating accrual is another measure of conservatism. Management has the discretion to decide the timing and the size of non-operating accrual so more negative non-operating accrual signifies higher reporting conservatism (Givoly and Hayn 2000). Comparison of mean non-operating accrual shows that firms do have more negative non-operating accrual in the pre-bond-IPO years. The comparison of median non-operating accrual shows it is non-increasing before bond IPO.

4.3.2. Financial Reporting Characteristics and Voluntary Disclosure Behaviour before and after Bond IPO

The loss reporting frequency of firms before and after bond IPO is presented in Panel A of Table 11. A firm is reporting loss if its net income is negative. Event year -1, -2, -3, 1, 2, are defined similarly as above. For example, if the earnings announcement is made within one year before the firm's bond IPO year, it is located in year -1. The results show firms do report fewer losses just prior to bond IPO, especially in quarterly earnings announcements. The loss

reporting frequency three years before bond IPO is 18.06% and it decreases to 13%- 14% approaching bond IPO year. It is consistent with the evidence shown in Jiang (2008) that firms want to decrease the cost of debt by avoiding reporting losses, showing that firms cater for bond market. The possible explanation for firms reporting much more losses after issuing bonds is due to the heavy interest expense. Compared with other firms in COMPUSTAT, firms which have bond IPOs are quite profitable, which make them qualified to issue bonds.

Panel B of Table 11 shows the bad news reporting frequency before and after bond IPO. A firm is reporting bad news if its EPS is lower than the consensus analyst forecast. Consensus analyst forecast is the median of analyst forecasts which are issued within one year or one quarter before annual or quarterly earnings announcements. There is no significant adjustment of bad news reporting frequency before bond IPO. While firms report bad news more frequently after bond IPO, especially in annual earnings announcements. In the untabulated analysis, analyst coverage after bond IPO is compared with that of the matched firms which have similar analyst coverage and market equity before bond IPO years. It turns out that the analyst coverage increases significantly after bond IPO.

Mandatory financial reporting is subject to the generally accepted accounting principles while voluntary disclosure is more flexible. Table 12 shows firms report fewer losses in management forecasts than in earnings announcements, even after bond IPOs. There are no clear patterns of loss

reporting frequency and bad news reporting frequency in management earnings forecasts, maybe due to the small number of observations.

4.4. Empirical Results

4.4.1. How Conditional Conservatism Changes before Bond IPO? –Basu Regression

This section presents the results of the modified Basu regression. In the main test, $E_{i,t}$ is the earnings per share (data53) which includes extraordinary items for firm *i* in fiscal year *t* and $RET_{i,t}$ is the cumulative stock return of firms *i* from 3 months after fiscal year end *t-1* to 3 months after fiscal year end *t*. Other variables are defined above. Continuous variables are winsorized at 1% on both sides.

Firstly $DBE_{i,t}$ equals one when firms are in years before event year -2 and equals 0 when firms are in event year -1, -2 . Because reporting conservatism in event year -1, -2 needs to be compared with that in previous years, this test restricts the sample further into firms with at least 4 years data before their bond IPOs, which decreases the sample to 363 firms and 3867 observations. Table 13 presents the regression results.

The results show that firms do report more conservatively before their bond IPOs. When firms are in years before event year -2, the incremental sensitivity of earnings to negative stock returns is 0.071 (0.224-0.153). While when firms are in event year -1, -2, the incremental sensitivity of earnings to negative stock returns is 0.224. The difference is captured by α_7, which is

negative and statistically significant.

Next $DBE_{i,t}$ is defined to equal one when firms are in years before event year -3 and equals 0 when firms are in event year -1, -2 and -3. Similarly to get better comparison, this test restricts the sample further into firms with at least 6 years data before their bond IPOs, which decreases the sample to 270 firms and 3473 observations. Table 14 presents the regression results which are qualitatively similar: the incremental sensitivity of earnings to negative stock returns is 0.049 (0.255-0.206) when firms are in years before event year -3 and the incremental sensitivity of earnings to negative stock returns is 0.255 when firms are in event year -1, -2 and -3. Overall the results support the hypothesis and present the evidence that firms anticipate market's demand for conservatism and report more conservatively approaching bond IPO.

4.4.2. How Conditional Conservatism Changes Before Bond IPO? –Accrual-CFO Model

Next this paper uses the accrual based model to test whether conditional conservatism changes immediately before firms issue bonds for the first time. Table 15 presents the regression results of model (5).

In table 15, accruals are defined as net income including extraordinary items (data172) minus operating cash flow (data308), which avoids the problem discussed in Hribar and Collins (2002). Other variables are defined above. All the continuous variables are winsorized at 1% on both sides.

$DBE_{i,t}$ is the indicator variable which equals 0 when firms are in event year -1 or -2 and equals 1 when firms are in years before event year -2.

Reporting conservatism before event year -2 is treated as benchmark and the reporting conservatism at year -1 or -2 is compared separately with the benchmark to test whether firms report more conservatively adjacent to bond IPO. To get a better test setting for comparison, this test restricts the sample further into firms with at least 4 years data before their bond IPOs, which decreases the sample to 178 firms and 1306 observations.

If firms anticipate market's demand for conservatism and report more conservatively before bond IPO, β_{11}, which measures the change of incremental sensitivity of accruals to negative cash flow, should be negative. The results support the hypothesis: β_{11} is negative and statistically significant both in event year -1 and -2. In event year -2, the incremental sensitivity of accruals to negative cash flow is 0.436 and the benchmark is just 0.007 (0.436-0.429). In event year -1, the incremental sensitivity of accruals to negative cash flow is 0.195 and the benchmark is just 0.028 (0.195-0.167).

In the untabulated analysis, the reporting conservatism before event year -3 is treated as the benchmark and the reporting conservatism in year -1, -2 or -3 is compared with the benchmark. The results are consistent with the results in table 8 for event year -1 and -2, whereas the result for year -3 is not significant.

4.4.3. Sensitivity Tests

For the modified Basu regression, sensitivity tests focus on different definitions of earnings per share and stock returns. All the samples restrict the firms to those with at least 4 years' data before bond IPOs. In the main test

presented in Table 13, earnings per share include extraordinary items and $RET_{i,t}$ is the cumulative stock return of firms *i* from 3 months after fiscal year end *t-1* to 3 months after fiscal year end *t*. The sensitivity tests define $E_{i,t}$ as earnings per share excluding extraordinary items (data58) or $RET_{i,t}$ as the fiscal year stock return. All the results of the sensitivity tests are consistent with the main result in Table 13 but the adjusted R^2 is slightly higher when $RET_{i,t}$ is the fiscal year stock return. The reason might be that the fiscal year return is a better measure of 'contemporaneous' stock return. The results are similar when earnings exclude extraordinary items or not, showing that extraordinary items do not affect the conservative reporting property.

For the accrual based model, sensitivity test changes the net income (data172) to income before extraordinary items (data 123). The untabulated analysis suggests that the results are qualitatively similar with the results in Table 15 as well.

4.5. Conclusion

Overall, this chapter examines whether firms adjust their reporting conservatism before their bond IPOs. Using the sample of US public firms, this paper firstly estimates the change of conditional conservatism via altered Basu regression and the results show that firms do report more conservatively before their bond IPOs. Then the results of the accrual-based model confirm the finding. All the results are significant under different specifications of earnings and stock returns. This paper also finds that firms cater for bond market's

asymmetric reaction to losses and profits and report fewer losses right before bond IPOs. Reporting fewer losses does not conflict with reporting conservatively. These firms are quite profitable and they report conservatively to be qualified to issue bonds.

These results provide the evidence about the incremental importance of debt market over equity market for conservatism in reporting. Conditional on being public listed, firms increase their reporting conservatism before their bond IPOs, which shows that conservatism is more closely related with debt market.

5. The Credit Market Relevance of Management Earnings Forecasts

5.1. Introduction

Management earnings forecasts are voluntary disclosures that have increasingly become an important source of information for capital market participants. An extensive literature in equity markets has found that management earnings forecasts decrease information asymmetry about the firm, and that the issuance of these forecasts is driven mainly by stock price considerations (e.g., Coller and Yohn 1997; Ajinkya et al. 2005; Graham et al. 2005). However, little is known about the usefulness of these forecasts to credit market participants, whose response to management forecasts can differ substantially from those of equity market participants.

Relative to earnings announcements, which face relatively greater

auditing and regulatory requirements, management forecasts offer greater flexibility to communicate information in a more timely manner. Consistent with this argument, management forecasts have been shown to be more important than earnings announcements in conveying information to stock markets (e.g., Ball and Shivakumar 2008; Beyer et al. 2009). However, unlike stocks, contractual features of debt securities, such as covenants and performance pricing agreements, are often written on reported earnings numbers rather than on forecasted earnings. The uncertainty relating to the impact of accounting numbers on debt contracts is fully resolved only at earnings announcements, making earnings releases potentially more informative for credit markets than for stock markets. Thus the issue of whether credit markets react more to management's earnings forecasts or to earnings announcements is an empirical one.

We assess the impact of management earnings forecasts in credit markets by studying the reactions of credit default swap (CDS) spreads to forecasts and earnings announcements. We focus on credit default swaps rather than on secondary bond prices for several reasons.[28] First, CDS spreads provide a relatively pure pricing of the default risk of the underlying entity (e.g., Hull et al. 2004; Longstaff et al. 2005), and CDS contracts are more standardized and homogeneous. In contrast, bond prices are more likely to be affected by

[28] A CDS is in essence an insurance contract protecting against losses arising from a default. In a CDS contract the buyer of credit protection pays to the seller of protection a periodic fee (i.e., the CDS spread) to insure against default of a debt security issued by a third party, called the reference entity. In the event that the reference entity defaults, the buyer typically delivers to the seller the debt owed by the reference entity in return for a lump sum equal to the face value of the debt. CDSs have turned out to clearly dominate other types of credit derivatives such as credit-linked notes or total return swaps in terms of market volume and standardization.

benchmark interest rates and idiosyncratic bond features, such as maturity, seniority, coupon rates, embedded options, and guarantees, which can cause substantial heterogeneity in the market's reactions to management forecasts. Second, institutional features of the CDS market facilitate a continuous flow of trades compared with the bond market, where short positions are difficult to achieve. Recent research by Blanco et al. (2005) and Zhu (2006) has shown that CDS spreads respond more quickly to changes in credit conditions. Third, bond markets are typically characterized by low levels of liquidity, because many bonds are tied up in "buy and hold" portfolios of institutional investors (e.g., Warga, 2004), which causes bond prices to reflect liquidity concerns.[29] These are less of an issue for CDS contracts, which tend to be highly liquid.[30]

Our main findings are as follows. Using a sample of management forecasts issued over the period 2001 to 2008, we document that CDS spread changes are significantly and negatively associated with management forecast news. A change in management forecast news from the 10th to the 90th percentile of the distribution (i.e., from 17.9% below to 7.9% above the analyst consensus estimate) causes the market-adjusted CDS spread reactions to decrease over the five-day announcement window from 0.64% to −0.28%. These reactions are robust to the inclusion of a variety of control variables, such as announcements of credit-rating agency revisions, volatility of daily stock

[29] Longstaff et al. (2005), Chen et al. (2007) and Bushman et al. (2010) find that a large proportion of bond spreads are determined by illiquidity factors, which do not reflect the default risk of the underlying bond. Bushman et al. (2010) further document that the effect of illiquidity factors on bond spreads has increased dramatically during the credit crisis.

[30] The greater liquidity is due partly to the lower capital required to trade in CDS contracts, where only the credit risk premia are paid for, as opposed to the full face value that needs to be paid for in the case of a bond.

returns and CDS spreads, as well as the adjustment of the error terms to control for potential self-selection of the management forecast issuance.[31]

The credit market reactions are significantly greater for the forecasts issued during the credit crisis period than for forecasts issued in the pre-crisis period. However, CDS spread reactions are larger for both good news and bad news during the crisis period, in contrast to the predictions based on the investor sentiment. The lack of sentiment effects in CDS spread changes is possibly due to the almost exclusive trading by institutional investors in credit markets.

Additional exploratory analysis, to uncover whether the market reactions are equally important for all firms and all types of forecasts, documents interesting results. First, consistent with the rational expectations model of Veronesi (1999) that asymmetric reactions between good and bad news are amplified in periods of low uncertainty, we find that the credit market reacts primarily to bad forecast news in the pre-crisis period, but not during the credit crisis.[32] Second, consistent with debtholders' asymmetric payoff, where they bear losses when firms do badly but do not share the profits when firms do well, we also find that credit markets are most responsive to management

[31] The controls for rating agency announcements are important, since rating agencies have privileged access to information about borrowers, and devote considerable resources to analyze that information. While some market participants, particularly banks and large institutional investors, might enjoy similar informational advantages, many other investors in the credit market rely mainly on rating agencies' opinions when assessing the credit quality of borrowers.

[32] The intuition behind Veronesi's (1999) model is that, in a regime-switching economy, if investors have low uncertainty about which regime the economy is in, then good news in a bad state of the economy or bad news in a good state of the economy increases the uncertainty. Consequently, risk-averse investors demand greater compensation for bearing risk. This discount rate effect, combined with the good (bad) news, causes equilibrium prices to react more asymmetrically during periods of low uncertainty.

forecast news when the firms are rated below investment grade than when they are rated investment grade. Third, we find that the credit market reaction to management forecasts is a function of the forecast attributes. Specifically, higher-quality forecasts, measured either as regular forecasts or as short-horizon forecasts, lead to stronger credit market reactions.

In the pre-crisis period, although changes in credit spreads are negatively correlated with both earnings news and management forecasts news, the coefficient on earnings news is significantly lower. However, during the crisis, the coefficient of accounting earnings news becomes insignificant while the coefficient of management forecasts news doubles in magnitude. These inferences are robust to whether we focus on a sample of "bundled" management forecasts (i.e., forecasts issued during a five-day event window around the earnings-announcement window) or on a sample of "unbundled" forecasts (i.e., forecasts issued outside the earnings-announcement window).[33]

Finally, we examine whether management forecasts by a firm convey information on credit risks to other firms within the same industry. Our intra-industry analysis reveals a significant information spillover effect in the credit market and this contagion effect is significantly larger during the crisis period.

The rest of the chapter is organized as follows. Section 5.2 presents the data and the sample selection process. Section 5.3 presents univariate descriptive statistics and discusses the main results, and Section 5.4 concludes

[33] Rogers and Van Buskirk (2009) observe that bundled forecasts have become more common recently, increasing from approximately 15% of forecasts in the late 1990s to 75% of forecasts in 2007.

the chapter.

5.2. Sample Selection and Data Description

5.2.1 Sample Selection

We obtain CDS data over the period 2001 to 2008 from the database provided by Markit Group. This database contains the daily composite quotes of CDS spreads collected from major market participants, as well as information about debt seniority, restructuring clauses, credit ratings, and contract maturity. Markit achieves high data quality by removing outliers, stale observations, and quotes provided by less than three dealer contributors. To maintain contract homogeneity, we focus only on five-year CDS contracts of senior unsecured debts with modified restructuring clauses written on U.S. non-financial reference entities. The five-year CDS contracts are the most popularly traded ones in the U.S. market (Zhang et al. 2009). In unreported tests, we also use one-year CDS contracts. While this sample is smaller, the results are qualitatively similar to those reported in the paper. We match the CDS data manually with COMPUSTAT based on the name and the location of the reference entity. The matched CDS database consists of 710 firms (reference entities) and 846,261 daily observations of composite CDS spreads.

We collect management earnings forecasts, analysts' forecast and earnings announcement data from the First Call database for our sample. We select only earnings per share forecasts denominated in U.S. dollars. When multiple management earnings forecasts are issued simultaneously, we retain

only the forecasts with the shortest forecast horizon. After requiring the availability of CDS data and information on control variables (discussed below) around the announcements of management earnings forecasts, our sample consists of 3,212 non-bundled management forecasts (i.e., management forecasts that are not issued during an earnings announcement window) for 432 unique firms. Out of these, 2,543 and 669 forecasts are announced in the pre-crisis and crisis periods, respectively. To avoid contamination of management forecasts with earnings announcements, most of our analyses are based on this sample of non-bundled forecasts.

However, for robustness checks, and to evaluate the relative informativeness of earnings announcements and management forecasts, we additionally analyze a sample of 6,137 bundled management forecasts made by 448 unique firms. Of the bundled forecasts, 4,629 are issued in the pre-crisis period and 1,508 in the crisis period.

5.2.2 Variables and Research Design

We investigate short-window reactions of CDS spread changes to the announcements of management earnings forecasts by estimating the following pooled regression:

$$\begin{aligned} CDS\ Return_{it} = {} & \beta_0 + \beta_1 MF\ News_{it} + \beta_2 \sigma(CDS\ Return)_{it} + \beta_3 \sigma(Stock\ Return)_{it} \\ & + \beta_4 S\ \&\ P500\ Return_{it} + \beta_5 3\ Month\ Treasury_{it} + \beta_6 VIX_{it} \\ & + \beta_7\ Good\ Rating\ News_{it} + \beta_8\ Bad\ Rating\ News_{it} + \varepsilon_{it} \end{aligned} \quad (6)$$

The dependent variable (*CDS Return*) is the percentage change in CDS spreads over a five-day window centered on the management forecast

announcement date in excess of the average spread change for a matched basket of CDS contracts calculated over the same five-day window. For each firm announcing a management forecast, we obtain a matched basket of CDS contracts by selecting the CDS contracts with the same credit rating category as the announcing firm. We focus on spread changes in excess of the credit-rating-matched basket's spread changes to isolate the effects of market-wide shifts in spreads, which are likely to be a function of the credit quality of the reference entities. The main independent variable of interest in the regression, the management earnings forecast news (*MF News*), is calculated as the management earnings forecast minus the most recent consensus analyst earnings forecast divided by the absolute value of the most recent consensus analyst earnings forecast.[34] We use only range and point estimates when calculating *MF News*.[35]

The regression controls for a variety of variables that are potentially associated with daily changes in CDS spreads, following prior literature (see, for example, Zhang et al. 2009). Firm-specific control variables include volatility of daily CDS spreads (σ(*CDS Return*)) and prior stock return volatility (σ(*CDS Return*)). We measure these variables as the standard deviation of daily CDS spreads and stock returns over the period [−137, −6] relative to the

[34] We prefer to scale by absolute value of forecast errors rather than by stock price, since Cheong and Thomas (2009) find that forecast errors and seasonally differenced earnings per share do not vary with stock price. However, our results are robust to scaling by stock price, as well as to including the inverse of stock price as an additional explanatory variable in the regression.

[35] In the case of range estimates, we follow Anilowski et al. (2007) and compute management earnings forecasts as the average of high and low estimates when First Call's CIGCODEQ equals 'B', the lower estimate when CIGCODEQ equals 'G' and the higher estimate when CIGCODEQ equals 'H'.

management forecast announcement date (day 0). By including the prior volatility of the daily CDS spreads we control for intrinsic CDS volatility. Stock return volatility is included to capture the unobservable asset volatility, which is an important determinant of default probability. Following Zhang et al. (2009), we also include a set of macro-financial variables as additional controls: the S&P 500 index return (*S&P500 Return*), the risk-free rate (*3 Month Treasury*), and the S&P 500 implied volatility index (*VIX*). These macro-financial variables are included to control for potential effects of the macroeconomy on forecasting the firm's default probability. *S&P500 Return* is the average daily S&P 500 index return during the five-day window over which the dependent variable (change in CDS spread) is measured. Similarly, *3 Month Treasury* is the average three-month treasury rate during the five-day event window. We obtain information on Treasury rates from Federal Reserve Bank Reports, which are publicly available. We compute *VIX* as the average S&P 500 index implied volatility during the five-day event window. We obtain the S&P 500 implied volatility data from the Chicago Board Options Exchange Volatility Index. This measure captures the market's expectation of 30-day volatility of S&P 500 index options.

Prior research documents that the CDS market responds to announcements by credit-rating agencies (Hull et al. 2004; Norden and Weber 2004; Galil and Soffer 2008). To control for the confounding effects of these rating agency announcements, we include two dummy variables, *Good Rating News* and *Bad Rating News.* The *Good Rating News* (*Bad Rating News*) takes the

value of one if the forecasting firm's credit rating is upgraded (downgraded), or if the firm is put on the positive (negative) watch list, or if the firm receives a positive (negative) outlook by Standard & Poor's or Moody's during the five-day event window. Otherwise, it takes the value of zero.[36] We include separate dummy variables for positive and negative rating agency announcements to account for the asymmetric reaction to good rating news versus bad rating news in debt markets that is documented in prior research (e.g., Easton et al. 2009).[37]

To study the relative market reaction to management forecast news and earnings news, we extend Equation (1) by including a measure of earnings news (*EA News*). Earnings news is calculated as the actual earnings minus the most recent consensus analyst earnings forecast divided by the absolute value of the most recent consensus analyst earnings forecast.

Since one of the objectives of this paper is to examine how the informativeness of management forecasts to credit markets changes during the recent credit crisis, we estimate Equation (1) for subsamples separated into whether the forecast was issued before or after 1 July 2007, apart from estimating the regression for the entire sample period. The period before (after) 1 July 2007 is referred to as the "pre-crisis" ("crisis") period.

[36] Outlooks, watchlist additions and ratings are based on both public information about borrowers' operating and financial conditions and private information obtained through confidential discussions with borrowers (e.g., Jorion et al. 2005).

[37] We obtain qualitatively similar results when we drop management forecasts announced simultaneously with rating agency announcements. We lose a total of 111 observations for the unbundled sample and 118 observations for the bundled sample when we remove management forecasts simultaneously disclosed with rating agency announcements.

5.3. Results

5.3.1. Descriptive Statistics and Univariate Evidence

Panels A and B of Table 16 present descriptive statistics for the samples of 3,212 unbundled management forecasts and 6,137 bundled management forecasts. For the unbundled (bundled) management forecasts sample, the average CDS spread change, measured over the five trading days around the management forecast announcement date, is 1.5% (0.2%), which is significant at the 1% (10%) level. This implies that, on average, the credit market reacts to management earnings forecasts by increasing the spread. This increase in spread suggests that management forecasts on average reveal bad news to the credit market – a conclusion also supported by the univariate statistics on management forecast news (*MF News*). The mean *MF News* is −2.2% for unbundled management forecasts and −3.3% for bundled forecasts. In contrast to management forecasts, earnings announcements appear to convey good news, on average, to the market. The average earnings announcement news, *EA News*, is 4.6% compared with the most recent analyst consensus earnings forecast. This observation is consistent with the stylized fact that analysts' earnings estimates start optimistic at the beginning of a fiscal period and tend to become pessimistic as the earnings announcement date approaches (Matsumoto 2002; Richardson et al. 2004). Finally, less than 3.1% (0.3%) of unbundled management earnings forecasts are accompanied by rating agency announcements that convey bad (good) news. The proportions for bundled forecasts are similar, suggesting that the confounding effects of rating agency

announcements are unlikely to be a major concern in our sample.

Panels A and B of Table 17 provide the correlations (Spearman rank order) among the variables for the unbundled and bundled forecasts samples, respectively. As observed in Panel A, which reports the correlations for the unbundled forecasts sample, the CDS spread changes are significantly negatively correlated with the management forecast news, the contemporaneous S&P 500 index return, and the good news rating agency announcements. The CDS spread changes are significantly and positively correlated with the S&P 500 index implied volatility and the bad news rating agency announcements. The results in Panel B reveal that, for the bundled forecasts sample, in addition to the above significant correlations, the CDS spread changes are significantly negatively correlated with the contemporaneously released earnings news, and with the prior volatility of daily CDS spreads.

5.3.2 Main Results

Table 18 presents the results of the regressions of the five-day market-adjusted CDS spread changes on the management forecast news in the unbundled management forecasts sample. In unreported analyses we find similar results when we use CDS spread changes over a three-day window around the management earnings forecast, or only on the day when the management forecast is announced. To mitigate the effect of extreme observations, all regressions are based on winsorizing continuous variables at the top and bottom one-percentile. Our qualitative results are, however, robust to winsorizing at other percentiles (2%, 3% and 5%). Throughout the paper,

t-statistics for the regressions are computed using standard errors clustered at the firm level.

Table 18, Column 1, presents results from a pooled OLS regression using observations from the entire sample period.[38] The coefficient on *MF News* is −0.036 (*t-stat* = −3.09), suggesting that an increase in management news (i.e., news becomes more positive) is associated with a significant decrease in the default premium for the firm, relative to the expected market CDS spread change. In terms of economic magnitude, a change in management forecast news from the 10th to the 90th percentile of the distribution (i.e., from 17.9% below the most recent consensus analyst estimate to 7.9% above it) causes the market-adjusted CDS spread reactions to decrease from 0.64% to −0.28%. Thus management forecasts have both a statistically and an economically significant effect on CDS spreads.

In Columns 2 and 3 we estimate the regressions separately for the pre-crisis and crisis subsamples. While the coefficients on *MF News* continue to be significant and negative in both subsamples, the magnitude of the coefficient on *MF News* during the crisis period is more than twice that during the pre-crisis period. This difference in coefficient across the sub-periods is also statistically significant at less than 5% level. This suggests that the CDS market is more sensitive to firm-specific information provided through management forecasts during the crisis than in the pre-crisis period, which is consistent with

[38] Although the Hausman test and the Breush and Pagan tests indicate that OLS regressions are the most appropriate regressions for our sample relative to either a fixed-effects or a random-effects model, our conclusions remain unaltered when these alternative estimation methods are employed.

the evidence reported in prior studies on the effect of uncertainty on stock market reactions to news (e.g., Lang 1991).

With respect to control variables, the market-adjusted CDS spread changes are generally negatively associated with the average S&P 500 index return and the good news rating agency announcements, and positively associated with the average S&P 500 index implied volatility (*VIX*) and the bad news rating agency announcements during the event window.

Although managers provide earnings forecasts either qualitatively or as numerical estimates, our sample in the earlier analysis is restricted to quantitative forecasts only (i.e., point or range forecasts), because of the need for a forecast number to compute *MF News*. Nonetheless, as a robustness test, we check the sensitivity of our results to the use of qualitative forecasts, which provide directional guidance on future earnings. We implement this test by employing the methodology in Anilowski et al. (2007) to classify qualitative forecasts into good or bad news forecasts. We then replace *MF News* in Equation (6) with a dummy variable for bad news forecasts. In untabulated results we find that, during the pre-crisis period, the coefficient on the indicator for bad forecast news is 0.035 (*t-stat* = 1.62), implying that downward qualitative management forecasts increase the market-adjusted CDS spread changes. During the crisis, the coefficient on the bad forecast news dummy is 0.109 (*t-stat* = 2.13), indicating that the market reaction to downward management forecasts is much stronger during the crisis than before the crisis. The coefficient on the bad news indicator is significantly different across the

pre-crisis and crisis subsamples at the 10% level. Although these results are qualitatively similar to those reported in Table 18, we refrain from drawing strong conclusions from these regressions, as the sample of qualitative forecasts consists of only 387 observations. Of these, only 39 qualitative forecasts are issued during the crisis period.

Firms that choose to issue management forecasts may have certain (unobservable) features that are potentially correlated with the CDS reactions to management forecasts, which would introduce a self-selection bias in the analysis. To control for this, we employ the standard Heckman (1979) two-stage selection approach. We adapt the probit model in Chen et al. (2008) to estimate the probability of issuing a management earnings forecast in the first stage. The dependent variable is an indicator variable that equals one if the firm issues at least one management forecast during the calendar year, and zero otherwise. The independent variables include:

- institutional shareholdings (the percentage of shares held by institutional investors)
- analyst coverage (the number of analysts following the firm)
- the analyst forecast dispersion (the standard deviation of analyst earnings forecasts)
- return volatility (the standard deviation of daily stock returns)
- board independence (an indicator variable that equals one if more than 60% of directors are independent directors, and zero otherwise)
- board size (the number of directors)

- litigation risk (an indicator variable that equals one if the firm is in an industry with a high litigation risk, and zero otherwise)[39]
- firm size (the natural log of total assets)
- the market-to-book ratio (the market value of equity divided by the book value of equity)
- return on assets (net income divided by total assets)
- future equity financing (an indicator variable that equals one if the firm has seasonal equity offerings during the subsequent year, and zero otherwise)
- future debt financing (an indicator variable that equals one if the firm issues bonds during the subsequent year, and zero otherwise).

Except for the equity and debt financing indicators, all variables are computed in the year prior to the management forecast issuance year. Because of the additional data requirements, the number of observations for this analysis decreases to 2,862.

We estimate the probability of issuing a management forecast in each calendar year, and compute the inverse Mills ratio accordingly. Columns 4, 5 and 6 of Table 18 present regressions for the pooled sample as well as for the two subsamples (pre-crisis and crisis) after we include the inverse Mills ratio as an additional control variable. We do not find any evidence of self-selection bias in the OLS regressions, as the coefficient on the inverse Mills ratio is insignificant at conventional levels in all regressions. Moreover, our inferences

[39] Industries with high litigation risk have the following SIC codes: 2833–2836, 3570–3577, 7370–7374, 3600–3674, 5200–5961, and 8731–8734.

from the OLS regressions remain unchanged when self-selection bias is controlled for. The results from the Heckman model continue to reveal a significantly negative coefficient on *MF News* for the entire sample, as well as for each of the subsamples, and the coefficient is significantly greater in the crisis period than in the pre-crisis period. The coefficient on *MF News* is −0.098 (*t-stat* = −3.46) during the crisis, whereas that before the crisis is −0.055 (*t-stat* = −3.82), and these coefficients are significantly different from each other at the 10% level.

5.3.3 Cross-sectional Results

We next investigate whether the CDS market reaction to management forecast news varies with the forecast properties, and with the credit riskiness of the forecasting firms. Apart from helping us understand when credit markets view management forecasts as being more informative, the cross-sectional analysis also provides corroborative evidence that the market reactions reported in Table 18 are driven by management earnings forecasts, rather than reflecting noise or research design choices.

Table 19 presents the cross-sectional results where the dependent variable is the market-adjusted CDS spread change around the five-day announcement window. In this analysis we extend Equation (6) by interacting *MF News* with an indicator variable that captures either the forecast attribute or the characteristics of the forecasting firms. We consider a variety of interactive variables, including the sign of the forecast news, whether the forecasts represent regular forecasts or not, the forecast horizon, and the credit riskiness

of the forecasting firm. We estimate these regressions separately for the pre-crisis and crisis periods.

Column 1 of Table 19 presents the results when the indicator variable takes the value one for bad news (*MF News* is less than or equal to zero) and zero otherwise. Since bad news management forecasts are perceived to be more credible than good news management forecasts (e.g., Hutton et al. 2003; Rogers and Stocken 2005), and since debtholders face an asymmetric payoff function, we expect credit market reactions to be stronger for bad forecast news than for good forecast news. Our sample includes 2,126 bad news forecasts, of which 449 are issued during the financial crisis. Consistent with our expectations, market-adjusted CDS spreads react more negatively to bad news management forecasts in the pre-crisis period. The coefficient on *MF News * Indicator* for bad news is −0.074 with a t-statistic of −2.78 in the pre-crisis period. However, during the financial crisis, there is no significant difference in the credit market reactions to good and bad news. The coefficient on *MF News* is significant and equals −0.046 (*t-stat* = −1.96) during the crisis period, and is also significantly different from the corresponding coefficient from the pre-crisis period at the 5% level. In contrast to the significantly larger market reaction to bad news in the pre-crisis period, the incremental coefficient on bad forecast news is insignificant during the crisis period, suggesting that market participants react similarly to good forecast news and bad forecast news during the crisis. This reduced asymmetry in market reaction to good news and bad news during the crisis is consistent with the predictions of Veronesi (1999).

We next analyze whether credit markets react differently to forecasts issued as part of a regular forecasting strategy, relative to irregular forecasts, by adding an interactive dummy for *Habitual* forecasts, which we define as forecasts made by a firm that has issued at least four management earnings forecasts in the prior year. Since irregular or sporadic forecasts can be used manipulatively by managers, we expect credit market reactions to be stronger for *Habitual* forecasts. Of the entire sample, 2,619 management forecasts are classified as *Habitual* forecasts. In the crisis subsample, 592 forecasts are categorized as *Habitual*.

The results in Column 2 show that credit-spread reactions are not different across sporadic and habitual forecasts in the pre-crisis period. The coefficient on *MF News*Indicator* for *Habitual* is statistically indistinguishable from zero. However, during the financial crisis, we find that the coefficient on *MF News* is statistically insignificant, while the incremental coefficient on *MF News*Indicator* is significant and equals to −0.200 (*t-stat* = −3.78). These findings indicate that during the credit crisis, when managers had potentially greater incentives to issue forecasts strategically and manipulatively, credit markets almost entirely ignored sporadic forecasts, and reacted significantly more strongly to Habitual forecasts than to similar forecasts issued in the pre-crisis period.

Column 3 of Table 19 reports results when the indicator variable captures the forecast horizon. Based on prior studies, such as Pownall et al. (1993) and Rogers and Stocken (2005), we expect forecasts issued with a

shorter horizon to be more informative. We test this prediction by including the interactive indicator variable *Preannouncement* in Equation (6). *Preannouncement* takes the value one if the firm releases management earnings forecast after fiscal period-end but before the earnings announcement date, and zero otherwise. Our sample includes 662 earnings preannouncements, of which 113 are issued during the financial crisis. We find that, although credit markets react both to preannouncements and to longer-horizon forecasts, the credit market response is significantly stronger for preannouncements, both before and during the financial crisis. The coefficient on *MF News* interacted with indicator for *Preannouncement* is −0.044 (*t-stat* = −1.66) for the pre-crisis period and −0.069 for the crisis period (*t-stat* = −2.29).

Finally, we investigate whether CDS market reactions to management forecast news are a function of the forecasting firm's credit-riskiness. Prior research has shown that credit prices are more sensitive to news when a firm has lower credit ratings (e.g., Easton et al. 2009). We examine the impact of credit ratings on the relationship between management earnings forecasts and CDS spread changes by creating an indicator variable *Speculative credit rating*, which equals one if the firm's credit rating is below BBB+, and zero otherwise. The full sample includes 1,705 forecasts made by firms with speculative credit ratings. The corresponding figure for the crisis period is 470. Consistent with the evidence in prior studies of equity markets, we observe from Column 4 of Table 19 that, in both sub-periods, credit markets respond primarily to forecast news from firms with speculative credit ratings, rather than from firms with

investment-level credit ratings.

Overall, these cross-sectional analyses provide evidence that the CDS spreads reactions to management forecast news vary not only over time but also across firms, depending on the forecast attributes chosen by the forecasting firm, and the forecasting firm's characteristics.

5.3.4. Management Forecasts Relative to Earnings Announcements

Management forecasts are often issued along with earnings announcements, and this phenomenon has become more popular recently. For instance, the proportion of bundled forecasts has increased from 46% in 2001 to about 75% in 2007 (Rogers and Van Buskirk 2009). It is possible for credit markets to react more to earnings releases than to management forecast releases, as issues relating to settlement of a debt's contractual features are clarified only at earnings announcements. In contrast, the arguments in Ball and Shivakumar (2008) suggest that earnings announcements are unlikely to be a good source of new information, and that management forecasts are likely to provide more timely information to capital markets.

We test the relative importance of management forecasts and earnings announcements for credit markets by estimating Equation (6) for a sample of bundled forecasts, after including the contemporaneously released earnings news as an additional explanatory variable in the regression. The results from this analysis are reported in Table 20.

From Column 1 we find that, when we control for the actual earnings news, management forecast news continues to be significantly negatively

associated with the market-adjusted CDS spread changes. The coefficient on *MF News* is −0.042 (*Preannouncement* = −5.05) and the coefficient on *EA News* is −0.018 (*t-stat* = −3.41). In terms of economic magnitude, a change in management forecast news from the 10th to the 90th percentile of the distribution (i.e., from 14.4% below the analyst consensus estimate to 5.1% above it) decreases the market-adjusted CDS firm spreads over the five-day announcement window by about 0.82%. However, the change in earnings news has a much smaller effect on the CDS spreads. A change in earnings news from the 10th to the 90th percentile of the distribution (i.e., from 7.1% below the analyst consensus estimate to 20.0% above it) decreases the market-adjusted CDS firm spreads over the five-day announcement window by 0.49%. Moreover, as observed earlier in the unbundled sample, the sensitivity of CDS spreads to management forecasts news increases significantly (*p-value* < 5%) in the crisis period. In contrast, the coefficient on actual earnings news turns insignificant in the crisis period, indicating that credit markets ignore backward-looking earnings information during the crisis.

In order to check the robustness of the results obtained using bundled sample, we create a subsample of 953 forecasts from the entire sample of unbundled management forecasts by matching the latest management forecast for a fiscal period with the earnings announcement of that fiscal period. In this matching of unbundled forecasts with earnings announcement, we ensure that the earnings announcements are not bundled with a management forecast.

The results from analyzing the matched unbundled sample are presented

in Table 21. In the entire sample, as well as in each sub-period, we find that the market-adjusted CDS spread changes are significantly negatively associated with the management forecast news, but not with earnings news. For example, in the pre-crisis period, the coefficient on *MF News* is a significant −0.073 with a *t*-statistic of −2.97, whereas the coefficient on *EA News* is insignificant. These results suggest that management forecasts preempt earnings announcements in conveying price-relevant information to the CDS market, consistent with the findings in the equity market (Ball and Shivakumar 2008). Consistent with the results in Table 20, the coefficient on management forecasts is significantly greater (at the 10% level) during the financial crisis, while the coefficient on earnings news continues to remain insignificant.

5.3.5. Intra-industry Contagion Analysis

In our last set of tests we examine the informational value of management forecasts to the broader credit market by investigating information spillovers from a management forecast issued by a firm to its industry peers (defined at three-digit SIC code level). Jorion and Zhang (2007) examine the CDS spread changes of industry rivals around major credit events, such as bankruptcy filings, and report that CDS spreads of industry competitors tend to increase at such major events. Although management forecasts are not as severe an event as bankruptcies, the forecasts may still provide industry-wide information, owing to the greater frequency at which these forecasts are issued.

For each management forecast announcement we calculate the average market-adjusted change in CDS spread of peer firms, which are defined as

firms in the same industry as the forecasting firm, but which do not have an earnings announcement or issue a management forecast in the five-day event window centered on the forecasting firm's forecast announcement date. We refer to the average market-adjusted changes in CDS spread of these peer firms as the changes in industry CDS spread. The sample for this analysis consists of 10,727 management forecasts, out of which 1,638 forecasts are issued during the financial crisis.

We test for industry contagion effects by regressing changes in industry CDS spread on the forecast news released by the forecasting firm. As done in earlier regressions, this analysis also controls for macroeconomic variables that are likely to affect CDS spreads. The results from this analysis are reported in Table 22.

Column 1 of Table 22 shows that, in the pooled sample, the average CDS spreads of industry peers react significantly to the news in forecasts issued by a firm. The coefficient on *MF News* is significant and equals to −0.002 (*t-stat* = −2.86). Thus, even though management forecasts are not major credit events, which potentially explains the smaller magnitude of the coefficients on *MF News* in this analysis relative to that observed in earlier analysis of CDS spreads on an individual firm's debt, we find that management forecasts are pertinent to price industry-level credit risks. Moreover, as observed in earlier tables, the results in Columns 2 and 3 of Table 22 indicate that the coefficient on *MF News* is significantly stronger during the financial crisis than in the pre-crisis period. These results indicate that the information about

industry-level credit risk contained in management forecasts is particularly relevant for credit pricing during periods of greater uncertainty. These results provide corroborative evidence that management forecasts are informative for credit markets, not just for pricing the credit risk of the issuing firm but also for pricing the credit risk of the entire industry.

5.4. Conclusion

We find that credit spreads react significantly and negatively to management forecast news, and that these reactions are stronger than those to actual earnings news. The credit market reactions to bad management forecast news and forecasts issued by credit risky companies are larger, reflecting the asymmetric payoff of debt securities. We also show that the impact of management forecasts on CDS spreads, relative to earnings announcements, becomes stronger during the recent credit crisis, consistent with theoretical arguments that greater market uncertainty increases the sensitivity of credit prices to relevant firm-specific information. Finally, we document that the average industry market-adjusted CDS spreads react significantly and negatively to forecast news issued by a firm in the same industry, and that this effect is magnified during the crisis period. This latter result suggests that management forecast disclosures generate contagion effects at the industry level, and thus provide useful industry-specific information.

Our findings add to the current evidence on the role of earnings-related information in credit markets. Extant research (e.g., Callen et al. 2009; Easton

et al. 2009) finds that accounting earnings affect credit spreads and trading volume in debt markets. To the best of our knowledge, this is the first study to document the relevance of management earnings forecasts to credit markets. Our evidence suggests that prior results based on short-window tests that do not control for management forecasts would exaggerate the importance of earnings announcements for credit markets. The information in management forecasts not only significantly revises the market's expectations about credit risks but also seems to dominate the information released by actual earnings news. Further, by showing that management forecasts change CDS spreads, we document that such voluntary firm disclosures affect the pricing of debt securities issued by the firm.

6. Conclusion

This dissertation is comprised of three inter-related papers. Chapter 3 investigates whether the bundled management forecasts of future earnings, conditional on their ex-ante accuracy, could mitigate post-earnings announcement drift. Using an estimated ex-ante forecast accuracy measure that is based on forecast ability, forecast difficulty and forecast environment, this paper finds that the bundled management forecasts of next quarter's earnings mitigate post-earnings announcement drift only when they have high ex-ante accuracy. By analyzing the effect of bundled management forecasts and their ex-ante accuracy on post-earnings announcement drift, this paper complements the previous papers which focus on the conference calls and the responsive

analysts surrounding the earnings announcement dates. Future study may investigate the effect of other information sources on post-earnings announcement drift, and provide more evidence to explain the formation of the drift.

Chapter 4 examines whether firms cater for bond market's demand for conservatism and increase their reporting conservatism before bond IPOs. Using modified Basu regression and accrual-CFO model, this paper shows that firms, which are already publicly listed, report more conservatively before issuing bonds for the first time, highlighting the incremental importance of debt market over equity market for conservatism in reporting. This paper is the first one to record the conservatism adjustment for bond market. Further research can explore the economic benefits and costs this conservatism adjustment brings to firms.

Chapter 5 examines the credit market reactions, via changes in credit default swap (CDS) spreads, to management forecasts, and compares the credit spread impact of management forecasts with that of mandated earnings for the periods before and during the credit crisis. The results suggest that CDS spread changes are significantly and negatively associated with management forecast news. Management earnings forecasts with more informative attributes and issued by firms with lower credit ratings have a more prominent role in credit markets. Management forecasts news provide more timely and price-relevant information to the CDS market than earnings announcements. Moreover, during the financial crisis, CDS spreads react only to management forecast news,

irrespective of whether management forecasts are issued simultaneously with earnings or not. The intra-industry analysis reveals a significant information spillover effect, which is more prominent during the crisis period. These results provide strong evidence that management earnings forecasts are an important disclosure mechanism for the credit market, especially during the financial crisis.

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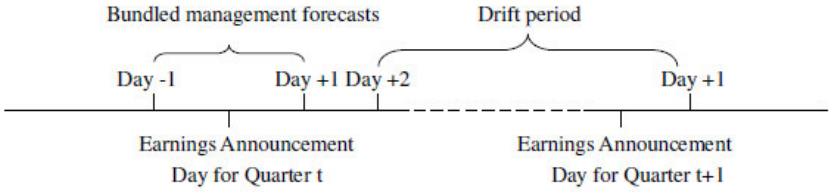

Figure 1. Timeline for measurement of variables. If the management forecast of Quarter t+1's earnings is issued within one trading day around the earnings announcement date of Quarter t, it is treated as the bundled management forecast. Post-earnings announcement drift period starts two trading days after the earnings announcement date of Quarter t and ends on the first trading day after the earnings announcement date of Quarter t + 1.

Table 1. Measurement of ex-ante forecast accuracy

Panel A. Test of the association between actual management forecast accuracy and forecast properties

Dependent variable: *ActualAccuracy*

	Annual Management Forecasts	Quarterly Management Forecasts
PriorAccuracy (α_1)	0.114***	0.181***
	(5.22)	(9.35)
ForecastHorizon (α_2)	-0.013***	-0.007***
	(-23.02)	(-12.26)
ForecastDifficulty (α_3)	-0.005***	-0.002***
	(−10.07)	(−11.33)
ForecastNews×*GoodNews* (α_4)	-0.935***	-0.997***
	(-10.66)	(-16.75)
ForecastNews×*BadNews* (α_5)	0.441***	0.330***
	(8.28)	(15.20)
CAR (α_6)	0.003***	0.0004***
	(7.26)	(3.86)
Litigation (α_7)	-0.0002	0.0001
	(-0.26)	(0.65)
Concentration (α_8)	-0.001	-0.0002
	(-1.26)	(-0.76)
Distress (α_9)	-0.001	-0.002*
	(-0.27)	(-1.85)
MB (α_{10})	0.0004***	0.0001***
	(5.49)	(6.09)
Size (α_{11})	0.0003*	-0.0001**
	(1.68)	(-1.97)
Bundle(α_{12})	0.001***	0.0001
	(4.51)	(1.22)
Adjusted R^2	45.66%	47.57%
Year Fixed-effect	Yes	Yes
Industry Fixed-effect	Yes	Yes
Number of obs.	20,701	21,790

This table presents results from the regression of actual management forecast accuracy (*ActualAccuracy*) on forecast properties. The management forecasts and the related actual earnings are announced from year 1995 to year 2007. The actual forecast accuracy is calculated as: *ActualAccuracy* = −1×*ABS(ActualEarnings* − *ManagementForecast)/Price*. *Price* is the average stock price one week before the management forecast announcement date. *PriorAccuracy* is the actual accuracy of the prior management forecast. The actual earnings in relation to the prior forecast need to be announced on or right before current management forecast announcement day. *ForecastHorizon* is the current management forecast horizon, calculated as the management forecast period end date minus the announcement date, deflated by 360. *ForecastDifficulty* is developed by performing Principal Axis Factoring over analyst forecast dispersion, standard deviation of previous analyst forecast errors, firm performance, predicted loss, stock return volatility, bid-ask spread and forecast width, as in Rogers and Stocken (2005). *ForecastNews* is calculated as the management forecast minus the consensus analyst forecast, deflated by *Price.*

(Notes continue on the following page)

GoodNews is an indicator variable which equals 1 if *ForecastNews* is non-negative, and 0 otherwise. *BadNews* is an indicator variable which equals 1 if *ForecastNews* is negative, and 0 otherwise. *CAR* is the firm's abnormal stock return cumulated from day -120 to day -1 relative to the management forecast announcement day. *Litigation* is an indicator variable which equals 1 if the firm is in industries with high litigation risk (Standard Industrial Classification codes 2833-2836, 3570-3577, 3600-3674, 5200-5961, 7370-7374, 8731-8734), and 0 otherwise. *Concentration* is measured by the Herfindahl index using the revenues of firms sharing the same four-digit SIC code. *Distress* is an indicator variable which equals 1 if the firm is in the most distressed decile predicted by Ohlson bankruptcy model, and 0 otherwise. *MB* is the firm's market value of equity deflated by book value of equity. *Size* is the natural log of the firm's total assets. *Bundle* is an indicator variable which equals 1 if the management forecast is issued along with the earnings announcement and 0 otherwise. All continuous variables are winsorized at the 1% and 99% levels. t-stats are in parentheses and are based on standard errors clustered on the firm level. *, **, and *** indicate significance at 10%, 5% and 1% levels respectively.

Panel B: Test of the association between actual management forecast accuracy and ex-ante forecast accuracy measures

Dependent variable: *ActualAccuracy*

Variable	Coefficients	Coefficients	Coefficients
Intercept (α_0)	-0.002***	-0.005***	-0.004***
	(-14.90)	(-29.76)	(-23.35)
EstimatedAccuracy (α_1)	0.752***		
	(29.08)		
PriorAccuracy (α_1)		0.357***	
		(16.96)	
AverageAccuracy (α_1)			0.373***
			(14.02)
Adjusted R^2	24.14%	9.16%	8.92%
Number of management forecasts	41,077	41,077	30,569[#]

This table presents results from the regression of actual management forecast accuracy (*ActualAccuracy*) on ex-ante forecast accuracy measures. *ActualAccuracy* and *PriorAccuracy* are defined in the notes to Panel A. *EstimatedAccuracy* is the ex-ante accuracy measure developed by Model (1). The *EstimatedAccuracy* of 20,148 annual management forecasts and 20,929 quarterly management forecasts can be estimated using Model (1). These management forecasts are issued from the second quarter of year 1997 to the second quarter of year 2007. *AverageAccuracy* is the prior average management forecast accuracy, and the actual earnings related with the prior management forecasts are announced during the previous four quarters relative to the current announcement quarter. *t*-stats are in parentheses and are based on standard errors clustered on the firm level. *, **, and *** indicate significance at 10%, 5% and 1% levels respectively.

[#] *AverageAccuracy* can not be calculated for 10,513 management forecasts because there are no actual earnings related with the prior management forecasts that are announced during the previous four quarters relative to the current announcement quarter.

Table 2. Descriptive statistics

Panel A: Comparison of firm characteristics based on mean values

Mean	Earnings announcements bundled with accurate forecasts	Earnings announcements bundled with inaccurate forecasts	Standalone earnings announcements	Column1 VS Column2 T-stat (p-Value)	Column1 VS Column3 T-stat (p-Value)	Column2 VS Column3 T-stat (p-Value)
Analyst forecast dispersion	0.002	0.007	0.007	-9.94 (<0.01)	-64.23 (<0.01)	-0.39 (0.69)
Firm size (in millions of dollars)	8,049.4	2,743.7	3,427.98	13.34(<0.01)	12.61(<0.01)	-3.78(<0.01)
Analyst coverage	33.06	27.92	18.60	9.53(<0.01)	39.64(<0.01)	22.12(<0.01)
Institutional shareholding	0.60	0.53	0.41	10.09(<0.01)	40.36(<0.01)	20.02(<0.01)
Trading volume (in millions of dollars)	624.64	439.59	317.37	5.71(<0.01)	12.89(<0.01)	7.51(<0.01)
Stock price	32.88	21.34	23.60	33.13(<0.01)	35.64(<0.01)	-7.96(<0.01)
Earnings persistence	0.35	0.36	0.29	-2.66 (<0.01)	13.74(<0.01)	14.57(<0.01)
BADNEWS	0.18	0.24	0.33	-7.90(<0.01)	-28.54(<0.01)	-12.39(<0.01)
4THQTR	0.19	0.22	0.22	-2.97 (<0.01)	-5.45 (<0.01)	-0.76 (0.45)
RESPONSIVE	0.67	0.58	0.36	9.02(<0.01)	48.39(<0.01)	27.90(<0.01)
Number of Firm-Quarters	6,237	4,284	58,048			

Panel B: Comparison of firm characteristics based on median values

Median	Earnings announcements bundled with accurate forecasts	Earnings announcements bundled with inaccurate forecasts	Standalone earnings announcements	Column1 VS Column2 Z-stat (p-Value)	Column1 VS Column3 Z-stat (p-Value)	Column2 VS Column3 Z-stat (p-Value)
Analyst forecast dispersion	0.001	0.003	0.003	-40.11 (<0.01)	-44.90 (<0.01)	7.54 (<0.01)
Firm size (in millions of dollars)	1,672.58	647.15	517.52	32.35(<0.01)	55.15(<0.01)	10.94(<0.01)
Analyst coverage	25	19	11	13.64(<0.01)	57.88(<0.01)	32.11(<0.01))
Institutional shareholding	0.72	0.62	0.44	10.59(<0.01)	43.23(<0.01)	21.42(<0.01)
Trading volume (in millions of dollars)	195.28	149.61	65.79	10.90(<0.01)	49.55(<0.01)	29.02(<0.01)
Stock price	29.17	17.17	17.67	37.03(<0.01)	49.09(<0.01)	-2.18(0.03)
Earnings persistence	0.39	0.41	0.33	-3.69(<0.01)	11.80(<0.01)	13.90(<0.01)
BADNEWS	0	0	0	-8.05(<0.01)	-24.14(<0.01)	-11.46(<0.01)
4THQTR	0	0	0	-2.99 (<0.01)	-5.23 (<0.01)	-0.76 (0.45)
RESPONSIVE	1	1	1	9.06(<0.01)	46.72(<0.01)	28.32(<0.01)
Number of Firm Quarters	6,237	4,284	58,048			

The sample includes 68,569 quarterly earnings announcements from the second quarter of 1997 to the second quarter of 2007. Analyst forecast dispersion is the standard deviation of analyst forecasts divided by the stock price at the fiscal quarter end. Firm size is measured by the market value of equity at the fiscal quarter end. Analyst coverage is the number of analysts following the firm. Trading volume is estimated by multiplying the closing price and the shares traded from day −272 to day −21 relative to the earnings announcement day. Stock price is the average stock price within one week before the earnings announcement. Earnings persistence is the first-order serial correlation of seasonally-differenced earnings estimated over the past 20 quarters. Institutional shareholding is the percentage of shares held by institutional investors at the quarter end that is closest to the earnings announcement day. *BADNEWS* is an indicator variable which equals 1 if the unexpected earnings are negative and 0 otherwise. *4THQTR* is an indicator variable which equals 1 if the earnings announcement is for the fourth fiscal quarter and 0 otherwise. *RESPONSIVE* equals 1 if there is at least one analyst revising the forecast of next quarter's earnings within two trading days after current quarter earnings announcement and 0 otherwise.

Table 3: Analysis about the bundled management forecasts

Panel A: Percentage of earnings announcements with bundled management forecasts by industry

	No of Earnings Announcements	Percentage of Earnings Announcements Bundled with Management Forecasts	Percentage of Earnings Announcements Bundled with Accurate Management Forecasts
Consumer Non-durables	3,487	15.83%	9.15%
Consumer Durables	1,837	13.88%	7.57%
Manufacturing	8,460	15.19%	8.20%
Energy	3,667	4.04%	2.32%
Chemicals and Allied Products	1,507	17.92%	12.48%
Business Equipment	17,186	19.40%	10.22%
Telecommunications	2,664	7.10%	3.90%
Utilities	2,119	12.36%	9.91%
Wholesale and Retail	8,431	19.77%	11.79%
Healthcare	8,358	11.35%	8.85%
Other	10,853	14.84%	9.28%

This table presents the percentage of earnings announcements with bundled management forecasts and with accurate bundled management forecasts by industry. The sample includes 68,569 quarterly earnings announcements from the second quarter of 1997 to the second quarter of 2007.There are 11 industry categories based on the industry classification in Fama and French (1997) and the finance industry (SIC codes 6000-6999) is excluded.

Panel B: Correlation matrix

	DSUE	*D_BUNDLE*	*DFD*	*DME*	*DAC*	*DVOL*	*DPRC*	*DINS*	*DEP*	*BAD*	*4THQTR*	*RESPONSIVE*
DSUE	1.000	0.031	0.039	0.005 (0.19)	-0.014	0.012	0.026	-0.007 (0.05)	-0.016	0.026	-0.007 (0.05)	-0.016
D_BUNDLE	0.033	1.000	0.142	0.137	0.149	0.114	0.143	0.106	0.057	-0.096	-0.017	0.197
DFD	0.048	0.142	1.000	0.403	0.014	0.082	0.545	0.148	-0.058	-0.201	-0.055	0.102
DME	0.012	0.137	0.406	1.000	0.588	0.653	0.748	0.338	0.051	-0.124	0.052	0.354
DAC	-0.009 (0.01)	0.149	0.018	0.587	1.000	0.581	0.389	0.285	0.152	-0.077	0.151	0.454
DVOL	0.015	0.115	0.083	0.648	0.577	1.000	0.307	0.207	0.059	-0.085	0.056	0.353
DPRC	0.033	0.143	0.548	0.748	0.389	0.304	1.000	0.354	0.067	-0.147	0.052	0.255
DINS	-0.004 (0.28)	0.110	0.153	0.348	0.295	0.215	0.363	1.000	0.162	-0.048	0.012	0.168
DEP	-0.017	0.060	-0.055	0.057	0.156	0.062	0.072	0.169	1.000	0.008 (0.03)	0.004 (0.36)	0.081
BAD	-0.777	-0.096	-0.201	-0.124	-0.078	-0.085	-0.148	-0.050	0.009 (0.01)	1.000	-0.010	-0.070
4THQTR	0.009 (0.02)	-0.017	-0.055	0.052	0.151	0.057	0.052	0.013	0.003 (0.37)	-0.010	1.000	-0.082
RESPONSIVE	0.029	0.197	0.103	0.353	0.454	0.352	0.255	0.174	0.082	-0.070	-0.082	1.000

This table presents the correlations among *DSUE, D_BUNDLE* and other drift-related variables. *DSUE* is the decile rank of the standardized

(Notes continue on the following page)

unexpected earnings, and is coded from 0 to 1. *D_BUNDLE* equals 1 if an earnings announcement is bundled with the management forecast of next quarter's earnings, and 0 otherwise. Analyst forecast dispersion (*DFD*), market value of equity (*DME*), analyst coverage (*DAC*), trading volume (*DVOL*), stock price (*DPRC*), institutional shareholding (*DINS*), and earnings persistence (*DEP*) are transformed into decile ranks within each calendar quarter, and coded from 0 to 1. *BADNEWS*, *4THQTR*, and *RESPONSIVE* are defined above. The upper-right triangle reports the Pearson product moment and the Spearman rank order presented in the lower-left triangle. Unless specified in parentheses with the corresponding two-sided p-values, all other correlations are significant at less than 1% level.

Table 4. Regression analysis about the effect of ex-ante management forecast accuracy on post-earnings announcement drift

Dependent variable: *ADJ_RET*

Independent Variable	Model (2a)	Model (2b)	Model (2c)	Model (2d)	Model (2e)	Model (2f)
Intercept (β_0)	−0.027***	−0.085***	−0.027***	−0.085***	−0.085***	−0.085***
	(−8.83)	(−13.58)	(−8.83)	(−13.59)	(−13.55)	(−13.41)
DSUE (β_1)	0.039***	0.068***	0.039***	0.066***	0.067***	0.065***
	(13.40)	(8.12)	(13.40)	(7.95)	(8.07)	(7.66)
D_BUNDLE (β_2)	0.018***	0.008*	0.020***	0.011**	0.010**	0.001
	(4.59)	(1.95)	(4.87)	(2.56)	(2.25)	(0.13)
DSUE × *D_BUNDLE* (β_3)	−0.013*	−0.007	−0.006	0.003	−0.006	0.004
	(−1.78)	(−0.96)	(−0.80)	(0.44)	(−0.85)	(0.51)
DSUE × *D_BUNDLE* ×*Dummy_EstimatedAccuracy* (β_4)			−0.015**	−0.027***		
			(−2.50)	(−4.48)		
DSUE × *D_BUNDLE* ×*Dummy_PriorAccuracy* (β_4)					−0.009	
					(−1.50)	
DSUE × *D_BUNDLE* ×*Dummy_AverageAccuracy* (β_4)						−0.009
						(−1.37)
DSUE ×*DFD* (β_5)		0.048***		0.051***	0.048***	0.053***
		(4.63)		(4.86)	(4.64)	(5.01)
DSUE ×*DME* (β_6)		−0.027		−0.026	−0.026	−0.028
		(−1.57)		(−1.50)	(−1.55)	(−1.61)
DSUE ×*DAC* (β_7)		−0.022*		−0.022*	−0.022*	−0.017
		(−1.82)		(−1.81)	(−1.83)	(−1.42)
DSUE ×*DVOL* (β_8)		−0.036***		−0.036***	−0.036***	−0.034***
		(−3.00)		(−2.98)	(−2.99)	(−2.79)
DSUE ×*DPRC* (β_9)		−0.004		−0.004	−0.004	−0.005
		(−0.28)		(−0.28)	(−0.28)	(−0.36)

(Table continues on the following page)

$DSUE \times DINS$ (β_{10})		0.006		0.005	0.006	0.008
		(0.74)		(0.62)	(0.72)	(0.93)
$DSUE \times DEP$ (β_{11})		0.011		0.011	0.011	0.009
		(1.37)		(1.36)	(1.34)	(1.09)
$DSUE \times BADNEWS$(β_{12})		0.001		0.001	0.001	−0.001
		(0.05)		(0.04)	(0.08)	(−0.07)
$DSUE \times 4THQTR$ (β_{13})		−0.022***		−0.022***	−0.022***	−0.020***
		(−3.42)		(−3.49)	(−3.43)	(−3.10)
$DSUE \times RESPONSIVE$ (β_{14})		−0.008		−0.009	−0.008	−0.010*
		(−1.36)		(−1.41)	(−1.39)	(−1.67)
DFD(β_{15})		−0.018***		−0.019***	−0.018***	−0.019***
		(−3.04)		(−3.07)	(−3.01)	(−3.17)
DME (β_{16})		0.044***		0.043***	0.044***	0.045***
		(4.48)		(4.46)	(4.47)	(4.51)
DAC (β_{17})		0.031***		0.032***	0.031***	0.030***
		(4.50)		(4.53)	(4.50)	(4.22)
$DVOL$ (β_{18})		−0.012**		−0.012*	−0.012*	−0.012*
		(−1.75)		(−1.78)	(−1.75)	(−1.75)
$DPRC$ (β_{19})		0.008		0.008	0.008	0.008
		(1.00)		(1.04)	(0.99)	(0.99)
$DINS$ (β_{20})		0.019***		0.020***	0.019***	0.019***
		(4.12)		(4.16)	(4.13)	(4.06)
DEP (β_{21})		0.011**		0.011**	0.011**	0.011**
		(2.42)		(2.42)	(2.44)	(2.48)
$BADNEWS$(β_{22})		0.009**		0.008**	0.008**	0.009**
		(2.12)		(2.05)	(2.02)	(2.22)
$4THQTR$ (β_{23})		0.029***		0.029***	0.029***	0.028***
		(8.00)		(8.00)	(8.03)	(7.60)
$RESPONSIVE$ (β_{24})		0.001		0.001	0.001	0.002
		(0.21)		(0.26)	(0.25)	(0.57)

(Table continues on the following page)

$\beta_3 + \beta_4$ (*p-value*)			−0.021 (0.006)	−0.024 (0.003)	−0.015 (0.10)	−0.005 (0.63)
Adjusted R^2	1.02%	2.15%	1.03%	2.17%	2.15%	2.15%
Year fixed-effect	Yes	Yes	Yes	Yes	Yes	Yes

This table examines the effect of the ex-ante management forecast accuracy on post-earnings announcement drift, controlling other drift-related variables. The time period is from the second quarter of year 1997 to the second quarter of year 2007. The dependent variable (*ADJ_RET*) is the drift-period return, which equals the compounded raw return minus the compounded benchmark return of the same CRSP size decile and the same CRSP exchange index (NYSE/AMEX or NASDAQ) that the firm belongs to. *DSUE* is the decile rank of the standardized unexpected earnings, and is coded from 0 to 1. The indicator variable *D_BUNDLE* equals 1 if an earnings announcement is bundled with the management forecast of next quarter's earnings, and 0 otherwise. *Dummy_EstimatedAccuracy* equals 1 if *EstimatedAccuracy* is greater than the median, and 0 otherwise. *Dummy_PriorAccuracy* equals 1 if *PriorAccuracy* is greater than the median, and 0 otherwise. *Dummy_AverageAccuracy* equals 1 if *AverageAccuracy* is greater than the median, and 0 otherwise. *DFD*, *DME*, *DAC*, *DVOL*, *DPRC*, *DINS*, *DEP*, *BADNEWS*, *4THQTR*, and *RESPONSIVE* are defined above. *t*-stats (in parentheses) are based on standard errors clustered on the firm level. *, **, and *** indicate significance at 10%, 5% and 1% levels respectively. Year dummies are included in the regressions.

Table 5. Determinants of issuing bundled management forecasts of next quarter's earnings

Dependent variable: *D_BUNDLE*

	Coefficients (p-value)
MF_EXIST	0.976 (<0.01)
LAG_DBUNDLE	1.116 (<0.01)
PRIOR_RETURN	0.144 (<0.01)
LOG_ME	0.064 (<0.01)
LOG_AC	0.096 (<0.01)
MEET	0.254 (<0.01)
ABS(SUE)	−10.876 (<0.01)
BADNEWS	−0.077 (<0.01)
LOSS	−0.431 (<0.01)
AFD	−0.220 (0.07)
Year Fixed-effect	Yes
Industry Fixed-effect	Yes
Number of Obs.	68,567

This table presents the results of the Probit regression with the dependent variable *D_BUNDLE* equals one if the earnings announcement is bundled with the management forecast of next quarter's earnings, and zero otherwise. The time period is from the second quarter of year 1997 to the second quarter of year 2007. *MF_EXIST* is an indicator variable which equals one if there is a prior management forecast for current quarter's earnings. The indicator variable *LAG_DBUNDLE* equals 1 if there is a bundled management forecast for the most recent quarterly earnings announcement. *PRIOR_RETURN* is the prior stock return over day -120 to day -1 relative to the earnings announcement date (day 0). *LOG_ME* is the natural log of market value of equity. *LOG_AC* is the natural log of analysts following. *MEET* is the proportion of earnings announcements which meet analyst expectations during the prior four quarters. *ABS(SUE)* is the absolute value of earnings surprise. The indicator variable *BADNEWS* equals one if current earnings surprise is negative and zero otherwise. The indicator variable *LOSS* equals 1 if the reported earnings number is negative and zero otherwise. *AFD* is the standard deviation of analyst forecasts for current quarter's earnings. All continuous variables are winsorized at the 1% and 99% levels.

Table 6. Regression analysis about the effect of ex-ante management forecast accuracy on post-earnings announcement drift, controlling the self-selection bias

Dependent variable: *ADJ_RET*

Independent Variable	Model (2a)	Model (2b)	Model (2c)	Model (2d)	Model (2e)	Model (2f)
Intercept (β_0)	−0.027***	−0.085***	−0.027***	−0.085***	−0.085***	−0.085***
	(−8.80)	(−13.63)	(−8.80)	(−13.62)	(−13.60)	(−13.44)
DSUE (β_1)	0.039***	0.067***	0.039***	0.066***	0.067***	0.064***
	(13.33)	(8.02)	(13.32)	(7.88)	(7.98)	(7.58)
D_BUNDLE (β_2)	0.024	0.003	0.025***	0.006	0.005	−0.004
	(5.29)	(0.63)	(5.61)	(1.23)	(0.96)	(−0.79)
DSUE × *D_BUNDLE* (β_3)	−0.013*	−0.006	−0.006	0.004	−0.005	0.005
	(−1.87)	(−0.82)	(−0.84)	(0.57)	(−0.73)	(0.61)
IMR	−0.005**	0.005**	−0.005**	0.004*	0.004**	0.004*
	(−2.44)	(2.13)	(−2.57)	(1.93)	(2.09)	(1.93)
DSUE × *D_BUNDLE* ×*Dummy_EstimatedAccuracy* (β_4)			−0.016***	−0.026***		
			(−2.71)	(−4.37)		
DSUE × *D_BUNDLE* ×*Dummy_PriorAccuracy* (β_4)					−0.009	
					(−1.45)	
DSUE × *D_BUNDLE* ×*Dummy_AverageAccuracy* (β_4)						−0.009
						(−1.38)
DSUE ×*DFD* (β_5)		0.048***		0.050***	0.048***	0.053***
		(4.61)		(4.83)	(4.61)	(5.03)
DSUE ×*DME* (β_6)		−0.027		−0.025	−0.027	−0.028
		(−1.56)		(−1.48)	(−1.55)	(−1.62)
DSUE ×*DAC* (β_7)		−0.022*		−0.022**	−0.022*	−0.018
		(−1.84)		(−1.83)	(−1.83)	(−1.44)
DSUE ×*DVOL* (β_8)		−0.036***		−0.037***	−0.036***	−0.034***
		(−3.00)		(−3.03)	(−3.01)	(−2.79)

(Table continues on the following page)

$DSUE \times DPRC$ (β_9)	−0.003	−0.003	−0.003	−0.003	−0.005
	(−0.25)	(−0.25)	(−0.25)	(−0.24)	(−0.35)
$DSUE \times DINS$ (β_{10})	0.006	0.006	0.006	0.006	0.008
	(0.72)	(0.68)	(0.72)	(0.72)	(0.95)
$DSUE \times DEP$ (β_{11})	0.011	0.010	0.011	0.011	0.009
	(1.40)	(1.31)	(1.39)	(1.39)	(1.11)
$DSUE \times BADNEWS$(β_{12})	0.003	0.002	0.003	0.003	0.001
	(0.19)	(0.15)	(0.21)	(0.21)	(0.05)
$DSUE \times 4THQTR$ (β_{13})	−0.022***	−0.022***	−0.022***	−0.022***	−0.020***
	(−3.45)	(−3.46)	(−3.48)	(−3.48)	(−3.13)
$DSUE \times RESPONSIVE$ (β_{14})	−0.008	−0.009	−0.008	−0.008	−0.010*
	(−1.37)	(−1.41)	(−1.38)	(−1.38)	(−1.66)
DFD(β_{15})	−0.018***	−0.018***	−0.018***	−0.018***	−0.019***
	(−2.94)	(−2.97)	(−2.91)	(−2.91)	(−3.10)
DME (β_{16})	0.044***	0.043***	0.043***	0.043***	0.045***
	(4.47)	(4.45)	(4.46)	(4.46)	(4.50)
DAC (β_{17})	0.032***	0.032***	0.032***	0.032***	0.030***
	(4.55)	(4.57)	(4.54)	(4.54)	(4.27)
$DVOL$ (β_{18})	−0.012**	−0.012*	−0.012*	−0.012*	−0.012*
	(−1.71)	(−1.71)	(−1.70)	(−1.70)	(−1.71)
$DPRC$ (β_{19})	0.008	0.008	0.008	0.008	0.008
	(0.99)	(1.01)	(0.98)	(0.98)	(0.97)
$DINS$ (β_{20})	0.020***	0.020***	0.020***	0.020***	0.019***
	(4.15)	(4.15)	(4.16)	(4.16)	(4.07)
DEP (β_{21})	0.011**	0.011**	0.011**	0.011**	0.011**
	(2.42)	(2.45)	(2.43)	(2.43)	(2.49)
$BADNEWS$(β_{22})	0.008*	0.008*	0.008*	0.008*	0.009**
	(1.93)	(1.88)	(1.84)	(1.84)	(2.06)
$4THQTR$ (β_{23})	0.030***	0.030***	0.030***	0.030***	0.028***
	(8.10)	(8.08)	(8.13)	(8.13)	(7.69)

(Table continues on the following page)

RESPONSIVE (β_{24})		0.001 (0.24)		0.001 (0.27)	0.001 (0.27)	0.002 (0.59)
$\beta_3 + \beta_4$ (*p-value*)			−0.022 (0.003)	−0.022 (0.006)	−0.014 (0.13)	−0.004 (0.69)
Adjusted R^2	1.02%	2.16%	1.03%	2.18%	2.16%	2.15%
Year fixed-effect	Yes	Yes	Yes	Yes	Yes	Yes

This table examines the effect of the ex-ante management forecast accuracy on post-earnings announcement drift, controlling the self-selection of issuing bundled management forecasts. The time period is from the second quarter of year 1997 to the second quarter of year 2007. IMR is the inverse Mills ratio calculated for all firms from the estimation results in Table 5. All other variables are defined as before. *t*-stats (in parentheses) are based on standard errors clustered on the firm level. *, **, and *** indicate significance at 10%, 5% and 1% levels respectively. Year dummies are included in the regressions.

Table 7. Market reaction to management forecasts issued between earnings announcements

	Dependent variable is the drift period return	Dependent variable is the management forecast announcement return	Dependent variable is the management forecast announcement return
Variable	Coefficients	Coefficients	Coefficients
Intercept (γ_0)	−0.117*** (−6.59)	−0.077*** (−8.04)	−0.076*** (−7.96)
DSUE (γ_1)	0.060*** (3.57)	0.027*** (3.61)	0.017** (2.13)
DSUE × *Dummy_EstimatedAccuracy* (γ_2)			0.040*** (5.42)
Adjusted R^2	4.06%	5.56%	6.18%
Number of quarterly earnings announcements	3,972	3,972	3,972
Dummy_EstimatedAccuracy = 1	1,488	1,488	1,488
Year fixed-effect	Yes	Yes	Yes

This table compares the drift period return and management forecast announcement return when management forecasts of next quarter's earnings are issued between earnings announcements. The time period is from the second quarter of year 1997 to the second quarter of year 2007. Variables are defined as before. *t*-stats (in parentheses) are based on standard errors clustered on the firm level. Year dummies are included in the regressions. *, **, and *** indicate significance at 10%, 5% and 1% levels respectively.

Table 8. Bond IPO sample selection

Firms which appear in COMPUSTAT on or after 1970:	19,655
- Firms lost by merging with CRSP	4,515
- Firms in financial industry	2,881
- Firms which do not have record in SDC	11,383
Firms which have bond IPO year	876
- Firms which have common stock IPO 3 years before bond IPO	107
Firms under study	769
- Firms which do not have at least 4 years' data before bond IPO	406
Firms in the modified Basu regression	363
- Firms which do not have data after 1987	185
Firms in the Accrual-CFO model	178

This table presents the bond IPO sample selection procedure. Starting from the firms which appear in COMPUSTAT on or after 1970, the sample decreases from 876 bond IPOs to 363 firms in the modified Basu regression, and 178 firms in the Accrual-CFO model.

Table 9. Number of firms with bond IPOs in each year

Bond IPO Year	No of Firms	Bond IPO Year	No of Firms
1970	21	1988	25
1971	22	1989	15
1972	12	1990	8
1973	11	1991	17
1974	5	1992	35
1975	16	1993	49
1976	10	1994	27
1977	21	1995	23
1978	24	1996	38
1979	17	1997	35
1980	14	1998	32
1981	9	1999	21
1982	16	2000	9
1983	30	2001	14
1984	26	2002	11
1985	48	2003	14
1986	62	2004	4
1987	26	2005	2

This table tabulates the number of bond IPOs in each year from 1970 to 2005. The total number of firms under study is 769.

Table 10. Market to book ratio and non-operating accrual before bond IPO

	Year -1		Year -2		Year -3		Year< -3	
	Mean	Median	Mean	Median	Mean	Median	Mean	Median
Market to book ratio	3.14	1.95	2.87	1.99	2.60	2.01	2.47	2.01
Non-operating accrual	-0.14	-0.11	-0.14	-0.13	-0.17	-0.11	-0.13	-0.11

This table presents the mean and median values of market to book ratio and non-operating accrual of 769 firms before their bond IPOs. The sample period is from 1970 to 2005. Market to book ratio is market value of equity divided by book value of equity. Non-operating accrual equals total accrual minus operating accrual, divided by total assets at the end of last fiscal year. The two variables are trimmed at 0.5% on both sides.

Table 11. Financial reporting before and after Bond IPO

Panel A Loss reporting frequency

	Total annual earnings announcements	Loss reporting frequency	Total quarterly earnings announcements	Loss reporting frequency
Year <-3	4,749	12.4%	15,531	18.06%
Year -3	349	11.74%	1,443	13.24%
Year -2	399	11.53%	1,649	14.07%
Year -1	458	12.45%	1,876	13.7%
Year 1	611	18.5%	2,442	21.66%
Year 2	610	24.9%	2,391	25.47%
Average level of firms in COMPUSTAT	205,241	27.73%	760,526	29.65%

A firm is reporting loss if its net income is negative. The number of bond IPOs is 769 and the period under study is from 1970 to 2005.

Panel B Bad news reporting frequency

	Total annual earnings announcements	Bad News reporting frequency	Total quarterly earnings announcements	Bad News reporting frequency
Year< -3	169	49.7%	713	31.84%
Year -3	86	51.16%	341	34.9%
Year -2	126	45.24%	458	31.66%
Year -1	166	49.4%	635	32.44%
Year 1	266	61.23%	1,008	36.51%
Year 2	307	59.28%	1,048	37.69%
Average level of firms in I/B/E/S	157,060	58.38%	189,699	35.04%

A firm is reporting bad news if its EPS is lower than the consensus analyst forecast. The number of bond IPOs is 230 and the period under study is from 1994 to 2005 due to I/B/E/S coverage.

Table 12. Voluntary disclosure before and after Bond IPO

	Total management forecast announcements	Loss reporting frequency	Bad news reporting frequency
Year <-3	45	0%	21.95%
Year -3	40	5%	16.21%
Year -2	67	8.96%	46.55%
Year -1	115	8.7%	39.05%
Year 1	251	11.16%	41.44%
Year 2	264	5.3%	50.21%
Average level of firms in Firstcall	47,269	7.9%	44.71%

Managers are forecasting losses if management forecasted EPSs are negative. Managers are forecasting bad news if management forecasts are lower than the consensus analyst forecasts. The number of bond IPOs is 230 and the period under study is from 1994 to 2005 due to Firstcall coverage.

Table 13. Result of Basu Regression when reporting conservatism before year -2 is used as the benchmark

Dependent variable: $E_{i,t}/P_{i,t-1}$

Independent Variable	Predicted	$DBE_{i,t}$ equals 0 in event year -1, -2
Intercept (α_0)	+	0.086 (7.98)
$DR_{i,t}$ (α_1)	?	-0.007 (-0.30)
$RET_{i,t}$ (α_2)	+	0.022 (1.70)
$RET_{i,t} \times DR_{i,t}$ (α_3)	+	0.224 (3.00)
$DBE_{i,t}$ (α_4)	?	0.021 (1.88)
$DBE_{i,t} \times DR_{i,t}$ (α_5)	?	-0.010 (-0.43)
$DBE_{i,t} \times RET_{i,t}$ (α_6)	?	0.025 (2.00)
$DBE_{i,t} \times DR_{i,t} \times RET_{i,t}$ (α_7)	-	-0.153 (-1.90)
Adj-R^2 (%)		5.26%
Number of Observations		3,867

This table presents the results of Basu Regression when reporting conservatism before year -2 is used as the benchmark. The sample period is from 1970 to 2005. $E_{i,t}$ is the earnings per share (including extraordinary items) for firm i in fiscal year t, $P_{i,t-1}$ is the price per share at the end of last fiscal year. $RET_{i,t}$ is the cumulative stock return of firms i from 3 months after fiscal end *t-1* to 3 months after fiscal year end *t*. E/P ratio and return are accumulated over two years. $DR_{i,t}$ is the indicator variable which equals 1 if $RET_{i,t}$ is negative and 0 otherwise. $DBE_{i,t}$ is the indicator variable which equals 0 in event year -1, -2 and equals 1 before year -2. Continuous variables are winsorized at 1% on both sides.

Table 14. Result of Basu Regression when reporting conservatism before year -3 is used as the benchmark

Dependent variable: $E_{i,t}/P_{i,t-1}$

Independent Variable	Predicted	$DBE_{i,t}$ equals 0 in event year -1, -2 and -3
Intercept (α_0)	+	0.063 (6.46)
$DR_{i,t}$ (α_1)	?	0.016 (0.81)
$RET_{i,t}$ (α_2)	+	0.053 (4.24)
$RET_{i,t} \times DR_{i,t}$ (α_3)	+	0.255 (3.51)
$DBE_{i,t}$ (α_4)	?	0.044 (4.18)
$DBE_{i,t} \times DR_{i,t}$ (α_5)	?	-0.029 (-1.32)
$DBE_{i,t} \times RET_{i,t}$ (α_6)	?	-0.001 (-0.08)
$DBE_{i,t} \times DR_{i,t} \times RET_{i,t}$ (α_7)	-	-0.206 (-2.61)
Adj-R^2 (%)		6.11%
Number of Observations		3,473

This table presents the results of Basu Regression when reporting conservatism before year -3 is used as the benchmark. The sample period is from 1970 to 2005. $E_{i,t}$ is the earnings per share (including extraordinary items) for firm i in fiscal year t, $P_{i,t-1}$ is the price per share at the end of last fiscal year. $RET_{i,t}$ is the cumulative stock return of firms i from 3 months after fiscal year end *t-1* to 3 months after fiscal year end *t*. E/P ratio and return are accumulated over three years. $DR_{i,t}$ is the indicator variable which equals 1 if $RET_{i,t}$ is negative and 0 otherwise. $DBE_{i,t}$ is the indicator variable which equals 0 in event year -1, -2 and -3 and equals 1 before year -3. Continuous variables are winsorized at 1% on both sides.

Table 15. Regression result of the Accrual-CFO model

Dependent variable: $ACC_{i,t}$

Independent Variable	Predicted	Year -1	Year -2
$Intercept\ (\beta_0)$	?	0.014 (1.78)	0.023 (2.73)
$CFO_{i,t}\ (\beta_1)$	-	-0.534 (-10.68)	-0.592 (-11.66)
$\Delta Sales_{i,t}\ (\beta_2)$	+	0.07 (5.92)	0.070 (5.73)
$FAsset_{i,t}\ (\beta_3)$	-	-0.031 (-4.90)	-0.030 (-4.66)
$DCFO_{i,t}\ (\beta_4)$	?	0.022 (1.10)	0.038 (1.69)
$DCFO_{i,t} \times CFO_{i,t}\ (\beta_5)$	+	0.195 (4.33)	0.436 (6.27)
$DBE_{i,t}\ (\beta_6)$	?	0.013 (1.70)	0.006 (0.83)
$DBE_{i,t} \times CFO_{i,t}\ (\beta_7)$	?	-0.047 (-1.03)	-0.003 (-0.06)
$DBE_{i,t} \times \Delta Sales_{i,t}\ (\beta_8)$	?	-0.038 (-3.73)	-0.038 (-3.64)
$DBE_{i,t} \times FAsset_{i,t}\ (\beta_9)$	?	0.002 (0.41)	0.0003 (0.06)
$DBE_{i,t} \times DCFO_{i,t}\ (\beta_{10})$	?	-0.037 (-1.72)	-0.055 (-2.36)
$DBE_{i,t} \times DCFO_{i,t} \times CFO_{i,t}\ (\beta_{11})$	-	-0.167 (2.02)	-0.429 (-4.36)
Adj-R^2 (%)		46.73%	47.14%
Number of Observations		1306	1306

This table presents the results of Accrual-CFO model when reporting conservatism before year -2 is used as the benchmark. The sample period is from 1987 to 2005. $ACC_{i,t}$ is the total accruals of firm *i* in year *t*, $CFO_{i,t}$ is cash flow from operation of firm *i* in year *t*. $\Delta Sales_{i,t}$ is the change of sales from year *t-1* to year *t*. $FAsset_{i,t}$ is the book value of fixed assets. $ACC_{i,t}$, $CFO_{i,t}$, $\Delta Sales_{i,t}$ and $FAsset_{i,t}$ are all divided by total assets at the end of year *t-1*. $DCFO_{i,t}$ is the indicator variable which equals 1 if $CFO_{i,t}$ is negative and 0 otherwise. $DBE_{i,t}$ is the indicator variable which equals 0 in event year -1, -2 and equals 1 before year -2. Continuous variables are winsorized at 1% on both sides.

Table 16. Descriptive statistics

Panel A: Unbundled Forecasts

	N	Mean	Median	STD	P10	P90
CDS Return	3,212	0.015	0.000	0.102	−0.068	0.100
MF News	3,212	−0.022	−0.005	0.225	−0.179	0.079
σ(*CDS Return*)	3,212	0.187	0.079	0.309	0.015	0.465
σ(*Stock Return*)	3,212	0.019	0.017	0.009	0.010	0.030
S&P500 Return	3,212	−0.013	0.015	0.446	−0.601	0.473
3 Month Treasury (%)	3,212	2.564	2.166	1.463	0.944	4.902
VIX	3,212	19.923	17.493	9.027	11.684	30.934
Good Rating News	3,212	0.003	0.000	0.058	0.000	0.000
Bad Rating News	3,212	0.031	0.000	0.174	0.000	0.000

Panel B: Bundled Forecasts

	N	Mean	Median	STD	P10	P90
CDS Return	6,137	0.002	0.000	0.083	−0.074	0.078
MF News	6,137	−0.033	−0.005	0.154	−0.144	0.051
EA News	6,137	0.046	0.022	0.216	−0.071	0.200
σ(*CDS Return*)	6,137	0.171	0.073	0.266	0.013	0.425
σ(*Stock Return*)	6,137	0.018	0.017	0.008	0.010	0.029
S&P500 Return	6,137	0.020	0.050	0.475	−0.558	0.464
3 Month Treasury (%)	6,137	2.843	2.486	1.588	0.948	5.086
VIX	6,137	19.616	16.124	10.560	11.270	30.566
Good Rating News	6,137	0.005	0.000	0.067	0.000	0.000
Bad Rating News	6,137	0.015	0.000	0.120	0.000	0.000

CDS Return is change in CDS premia around management forecast announcement date over a five-day window ([−2, 2]) minus average CDS premia of the market within the same credit rating group during the same five-day window. *MF News* is management forecast news calculated as management earnings forecast minus the most recent consensus analyst earnings forecast divided by absolute value of the most recent consensus analyst earnings forecast. *Crisis* is an indicator variable taking the value of 1 for observations starting with 1 July 2007 and 0 otherwise. *σ*(*CDS Return*) is the standard deviation of firm's CDS spread during the period [−137, −6] with respect to management forecast announcement date (day 0). *σ*(*Stock Return*) is the standard deviation of firm's stock return during the period [−137, −6] with respect to management forecast announcement date (day 0). *S&P500 Return* is the average daily S&P 500 index return during the window for which the dependent variable (*CDS Return*) is measured. *3 Month Treasury* is the average three-month treasury rate during the window for which the dependent variable (*CDS Return*) is measured. *VIX* is the average S&P 500 index implied volatility during the window for which the dependent variable (*CDS Return*) is measured. *Good Rating News* is an indicator variable taking the value of 1 if the firm's credit rating is upgraded, or if the firm is put in positive watchlist, or if the firm is put in positive outlook during the window for which the dependent variable (*CDS Return*) is measured, and 0 otherwise. *Bad Rating News* is an indicator variable taking the value of 1 if the firm's credit rating is downgraded, or if the firm is put in negative watchlist, or if the firm is put in negative outlook during the window for which the dependent variable (*CDS Return*) is measured, and 0 otherwise. *EA News* is earnings announcement news, calculated as earnings minus the most recent consensus analyst earnings forecast divided by the absolute value of the most recent consensus analyst earnings forecast. When simultaneous management forecasts are announced, the forecasts with the shortest forecast periods are included. Bundled forecasts refer to the forecasts issued within a five-day window of earnings announcements. The sample period is between 2001 and 2008. All non-indicator variables are winsorized at the top and bottom one-percentiles.

Table 17. Correlations

Panel A: Unbundled Forecast

		(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
(1)	*CDS Return*	1.000									
(2)	*MF News*	**–0.103**	1.000								
(3)	*Crisis*	**0.065**	0.010	1.000							
(4)	*σ(CDS Return)*	0.022	–0.024	**0.168**	1.000						
(5)	*σ(Stock Return)*	0.014	–0.023	**0.140**	**0.509**	1.000					
(6)	*S&P500 Return*	**–0.066**	0.019	**–0.117**	**–0.046**	**–0.047**	1.000				
(7)	*3 Month Treasury (%)*	–0.022	**–0.064**	**–0.095**	**–0.295**	**–0.288**	**0.047**	1.000			
(8)	*VIX*	**0.081**	–0.001	**0.427**	**0.308**	**0.493**	**–0.160**	**–0.548**	1.000		
(9)	*Good Rating News*	**–0.060**	–0.001	–0.004	0.002	0.018	–0.015	0.013	–0.004	1.000	
(10)	*Bad Rating News*	**0.093**	**–0.056**	**–0.043**	**0.068**	**0.069**	–0.012	0.022	0.022	–0.011	1.000

Panel B: Bundled Forecast

		(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
(1)	*CDS Return*	1.000										
(2)	*MF News*	**–0.098**	1.000									
(3)	*EA News*	**–0.087**	**0.308**	1.000								
(4)	*Crisis*	**0.053**	–0.007	0.021	1.000							
(5)	*σ(CDS Return)*	**–0.029**	**–0.058**	**0.033**	**0.135**	1.000						
(6)	*σ(Stock Return)*	0.019	**–0.044**	**0.059**	**0.190**	**0.507**	1.000					
(7)	*S&P500 Return*	**–0.023**	0.002	–0.001	**–0.102**	–0.006	0.001	1.000				
(8)	*3 Month Treasury (%)*	–0.018	**–0.032**	**0.024**	**–0.131**	**–0.311**	**–0.354**	**0.044**	1.000			
(9)	*VIX*	**0.055**	**–0.021**	0.003	**0.544**	**0.284**	**0.458**	**–0.126**	**–0.579**	1.000		
(10)	*Good Rating News*	**–0.034**	0.008	0.002	–0.016	–0.002	–0.006	0.008	**0.026**	**–0.022**	1.000	
(11)	*Bad Rating News*	**0.090**	**–0.070**	**–0.033**	–0.004	**–0.070**	**0.046**	0.014	–0.011	0.032	–0.008	1.000

This table provides Spearman correlations among variables of interest in our sample. Correlations in bold are significant at the 10% level or better. All variables are defined in Table 1. The sample period is between 2001 and 2008. All non-indicator variables are winsorized at the top and bottom one-percentiles.

Table 18. Association between management forecast news and CDS return (unbundled sample)

	Dependent variable: *CDS return*					
	(1)	(2)	(3)	(4)	(5)	(6)
	Pooled	Pre-crisis	Crisis	Pooled	Pre-crisis	Crisis
MF News	**−0.036**	**−0.032**	**−0.081**$^{++}$	**−0.058**	**−0.055**	**−0.098**$^{+}$
	(−3.09)	**(−2.55)**	**(−3.50)**	**(−4.43)**	**(−3.82)**	**(−3.46)**
σ(*CDS Return*)	−0.005	−0.002	−0.013	−0.009	−0.001	−0.043^{+}
	(−0.56)	(−0.22)	(−0.86)	(−0.88)	(−0.09)	(−1.88)
σ(*Stock Return*)	−0.372	−0.416	−0.017	−0.460	−0.772	0.711
	(−1.09)	(−1.06)	(−0.23)	(−1.18)	(−1.76)	(0.85)
S&P500 Return	−0.013	−0.008	−0.025	−0.009	−0.004	−0.020
	(−2.59)	(−1.33)	(−2.26)	(−1.80)	(−0.72)	(−1.86)
3 Month Treasury	−0.002	−0.002	−0.002	−0.003	−0.004	−0.002
	(−0.60)	(−0.51)	(−0.24)	(−1.08)	(−1.23)	(−0.25)
VIX	0.001	0.002	0.001	0.002	0.002	0.001
	(3.35)	(2.63)	(1.65)	(3.43)	(2.52)	(1.83)
Good Rating News	−0.086	−0.087	−0.090	−0.064	−0.045	−0.106^{+++}
	(−3.65)	(−3.03)	(−5.09)	(−4.54)	(−2.54)	(−6.76)
Bad Rating News	0.095	0.088	0.142	0.095	0.095	0.096
	(5.33)	(4.64)	(2.54)	(5.10)	(4.66)	(2.30)
Inverse Mills ratio				−0.005	0.003	−0.052
				(−0.28)	(0.18)	(−1.32)
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
R^2	0.058	0.051	0.079	0.068	0.066	0.089
N	3,212	2,543	669	2,862	2,277	585

The dependent variable is change in CDS premia in the five-day window around management forecast announcement date minus average CDS premia of the market within the same credit rating group during the same five-day window. Inverse mills ratio is defined as $\phi(Z_i\beta)/\Phi(Z_i\beta)$, where ϕ and Φ are standard normal p.d.f. and c.d.f., respectively, Z is the row vector of explanatory variables in the management earnings forecast choice model, and β is the column vector of coefficients estimated from management earnings forecast choice model. All other variables are defined in Table 1. When simultaneous management forecasts are announced, the forecasts with the shortest forecast periods are included. The sample period is between 2001 and 2008. All non-indicator variables are winsorized at the top and bottom one-percentiles. The *t*-values, reported in parentheses, are based on standard errors adjusted for clustering at firm level. $^{+++}$, $^{++}$ and $^{+}$ indicate that the coefficients between the pre-crisis and crisis periods are significantly different at the 1%, 5% and 10% levels, respectively, in two-tailed tests.

Table 19. Association between management forecast news and CDS return (unbundled sample): Cross-sectional tests

	Dependent variable: *CDS Return*							
	(1)		(2)		(3)		(4)	
	Indicator = Bad		Indicator = Habitual		Indicator = Preannouncement		Indicator = Speculative credit rating	
	Pre-crisis	*Crisis*	*Pre-crisis*	*Crisis*	*Pre-crisis*	*Crisis*	*Pre-crisis*	*Crisis*
Indicator	0.006	0.010	−0.001	0.029^{+}	0.014	0.026	0.000	−0.011
	(1.28)	(1.19)	(−0.09)	(2.05)	(2.74)	(2.23)	(0.03)	(−0.97)
MF News	0.013	−0.046^{++}	−0.040	0.063^{++}	−0.024	−0.057	0.003	0.002
	(0.69)	(−1.96)	(−1.78)	(1.48)	(−2.20)	(−2.41)	(0.34)	(0.31)
*MF News * Indicator*	−0.074	−0.038	−0.013	−0.200^{+++}	−0.044	−0.069	−0.066	−0.090
	(−2.78)	(−0.86)	(−0.37)	(−3.78)	(−1.66)	(−2.29)	(−2.79)	(−2.13)
σ(CDS Return)	−0.005	−0.014	0.017	−0.027	−0.001	−0.014	0.016	−0.027
	(−0.49)	(−0.93)	(0.88)	(−1.17)	(−0.09)	(−0.90)	(0.76)	(−1.20)
σ(Stock Return)	−0.609	−0.169	−0.841	0.348	−0.576	−0.247	−0.860	0.404
	(−1.56)	(−0.22)	(−1.35)	(0.30)	(−1.46)	(−0.33)	(−1.40)	(0.34)
S&P500 Return	−0.008	−0.025	−0.027	−0.062^{++}	−0.008	−0.024	−0.027	−0.061^{++}
	(−1.35)	(−2.27)	(−2.78)	(−4.43)	(−1.34)	(−2.18)	(−2.75)	(−4.38)
3 Month Treasury	−0.001	−0.002	−0.005	0.009	−0.001	−0.007	−0.005	0.010
	(−0.38)	(−0.22)	(−1.29)	(0.82)	(−0.31)	(−0.82)	(−1.18)	(0.87)
VIX	0.002	0.001	0.003	0.002	0.001	0.001	0.003	0.002
	(2.62)	(1.64)	(2.73)	(2.40)	(2.63)	(1.32)	(2.73)	(2.57)
Good Rating News	−0.085	−0.093	−0.096	−0.111	−0.086	−0.083	−0.096	−0.104
	(−2.97)	(−4.89)	(−2.64)	(−5.60)	(−2.96)	(−4.80)	(−2.63)	(−5.29)
Bad Rating News	0.083	0.142	0.124	0.141	0.086	0.127	0.123	0.151
	(4.34)	(2.51)	(3.75)	(2.51)	(4.44)	(2.14)	(3.73)	(2.70)
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
R^2	0.057	0.081	0.059	0.140	0.057	0.087	0.061	0.124
N	2,543	669	2,543	669	2,527	655	2,543	669

The dependent variable is change in CDS premia in the five-day window around management forecast announcement date minus average CDS premia of the

(Notes continue on the following page)

market within the same credit rating group during the same five-day window. *Bad* is an indicator variable taking the value of 1 if management forecast news is 0 or negative, and 0 otherwise. *Habitual* is an indicator variable taking the value of 1 if the firm releases at least four management earnings forecast over the last year, and 0 otherwise. *Preannouncement* is an indicator variable taking the value of 1 if the firm releases management earnings forecast after fiscal period end but before earnings announcement, and 0 otherwise. *Speculative credit rating* is an indicator variable taking the value of 1 if credit rating is below BBB+, and 0 otherwise. All other variables are defined in Table 1. When simultaneous management forecasts are announced, the forecasts with the shortest forecast periods are included. The sample period is between 2001 and 2008. All non-indicator variables are winsorized at the top and bottom one-percentiles. The *t*-values, reported in parentheses, are based on standard errors adjusted for clustering at firm level. $^{+++}$, $^{++}$ and $^{+}$ indicate that the coefficients between the pre-crisis and crisis periods are significantly different at the 1%, 5% and 10% levels, respectively, in two-tailed tests.

Table 20. Association between management forecast news and CDS return (bundled sample)

	Dependent variable: *CDS Return*		
	(1)	(2)	(3)
		Pre-crisis	Crisis
MF News	−0.042	−0.035	-0.076^{++}
	(−5.05)	(−4.02)	(−3.23)
EA News	−0.018	−0.016	−0.022
	(−3.41)	(−2.61)	(−1.31)
σ(*CDS Return*)	−0.013	−0.016	−0.007
	(−2.12)	(−2.32)	(−0.62)
σ(*Stock Return*)	0.083	0.203	−0.116
	(0.40)	(0.85)	(−0.31)
S&P500 Return	−0.003	−0.005	0.001
	(−1.07)	(−1.24)	(0.13)
3 Month Treasury	0.000	0.000	0.004
	(0.12)	(0.13)	(0.69)
VIX	0.001	0.002	0.001
	(5.70)	(3.70)	(3.84)
Good Rating News	−0.046	−0.047	−0.041
	(−2.61)	(−2.49)	(−0.72)
Bad Rating News	0.084	0.096	0.044^{+}
	(6.05)	(6.02)	(1.65)
Year fixed effects	Yes	Yes	Yes
p-value (*MF News* = *EA News*)	0.010	0.061	0.025
R^2	0.046	0.047	0.041
N	6,137	4,629	1,508

The dependent variable is change in CDS premia in the five-day window around management forecast announcement date minus average CDS premia of the market within the same credit rating group during the same five-day window. *EA News* is earnings announcement news calculated as earnings minus the most recent consensus analyst earnings forecast divided by absolute value of the most recent consensus analyst earnings forecast. All other variables are defined in Table 1. When simultaneous management forecasts are announced, the forecasts with the shortest forecast periods are included. The sample period is between 2001 and 2008. All non-indicator variables are winsorized at the top and bottom one-percentiles. The *t*-values, reported in parentheses, are based on standard errors adjusted for clustering at the firm level. $^{+++}$, $^{++}$ and $^{+}$ indicate that the coefficients between the pre-crisis and crisis periods are significantly different at the 1%, 5% and 10% levels, respectively, in two-tailed tests.

Table 21. Comparison of CDS market reaction to management forecast news versus earnings announcement news (unbundled sample)

	Dependent variable: *CDS Return*		
	(1)	(2)	(3)
		Pre-crisis	Crisis
MF News	−0.088	−0.073	−0.125^{+}
	(−3.09)	(−2.97)	(−1.71)
EA News	0.008	0.008	0.009
	(0.70)	(0.69)	(0.09)
σ(*CDS Return*)	0.001	0.002	0.046
	(0.13)	(0.13)	(0.75)
σ(*Stock Return*)	−0.510	−0.582	−0.148
	(−0.71)	(−0.75)	(−0.10)
S&P500 Return	−0.029	−0.027	−0.038
	(−2.83)	(−2.26)	(−1.78)
3 Month Treasury	0.002	0.002	0.014
	(0.42)	(0.28)	(0.69)
VIX	0.001	0.001	0.001
	(1.08)	(0.50)	(1.03)
Good Rating News	0.127	0.128	0.000
	(0.95)	(0.96)	(0.02)
Bad Rating News	0.120	0.118	0.245^{++}
	(2.87)	(2.75)	(6.83)
Year fixed effects	Yes	Yes	Yes
p-value (*MF News* = *EA News*)	0.001	0.001	0.183
R^2	0.081	0.079	0.165
N	953	874	79

The dependent variable is change in CDS premia in the five-day window around management forecast announcement date minus average CDS premia of the market within the same credit rating group during the same five-day window. *EA News* is earnings announcement news calculated as earnings minus the most recent consensus analyst earnings forecast divided by absolute value of the most recent consensus analyst earnings forecast. All other variables are defined in Table 1. When simultaneous management forecasts are announced, the forecasts with the shortest forecast periods are included. The sample period is between 2001 and 2008. All non-indicator variables are winsorized at the top and bottom one-percentiles. The *t*-values, reported in parentheses, are based on standard errors adjusted for clustering at the firm level. $^{+++}$, $^{++}$ and $^{+}$ indicate that the coefficients between the pre-crisis and crisis periods are significantly different at the 1%, 5% and 10% levels, respectively, in two-tailed tests.

Table 22. Association between management forecast news and CDS return of other firms in the same three-digit SIC code (unbundled sample)

	Dependent variable: *CDS Return*		
	(1)	(2)	(3)
		Pre-crisis	Crisis
MF News	−0.002	−0.002	−0.005^{+}
	(−2.86)	(−2.30)	(−1.86)
S&P500 Return	−0.010	−0.010	−0.010
	(−6.90)	(−6.39)	(−2.72)
3 Month Treasury	0.002	0.001	0.002
	(1.48)	(0.66)	(0.84)
VIX	0.001	0.001	0.001^{+++}
	(7.93)	(8.74)	(2.70)
Year fixed effects	Yes	Yes	Yes
R^2	0.027	0.026	0.022
N	10,727	9,089	1,638

The dependent variable is the average market-adjusted change in CDS premia of firms in the same three-digit SIC code on the five-day window around management forecast announcement date of a sample firm. All other variables are defined in Table 1. When simultaneous management forecasts are announced, the forecasts with the shortest forecast periods are included. The sample period is between 2001 and 2008. All non-indicator variables are winsorized at the top and bottom one-percentiles. The *t*-values, reported in parentheses, are based on standard errors adjusted for clustering at firm level. $^{+++}$, $^{++}$ and $^{+}$ indicate that the coefficients between the pre-crisis and crisis periods are significantly different at the 1%, 5% and 10% levels, respectively, in two-tailed tests.