BANKRUPTCY CONTAGION AND THE MARKET RESPONSE TO INTRA-INDUSTRY EARNINGS NEWS

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A dissertation submitted to the faculty of the University of North Carolina at Chapel Hill in partial fulfillment of the requirements for the degree of Doctor of Philosophy in the Kenan-Flagler School of Business.

Chapel Hill
2010

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ABSTRACT

EDWARD L. OWENS: Bankruptcy Contagion and the Market Response to Intra-Industry Earnings News
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I seek a deeper understanding of bankruptcy contagion and its impact on the industry information environment by examining how intra-industry bankruptcy affects the equity market response to earnings news subsequently released by industry survivors. I document significant negative asymmetry, i.e., a weaker response to good relative to bad news, in the response to earnings news in the wake of intra-industry bankruptcy. Results are consistent with a combination of model uncertainty-induced asymmetry and an on-average reduction in earnings informativeness stemming from increased cross-firm default risk assessments. Results are further consistent with managerial preemptive disclosure of good news exacerbating, but not completely explaining, this asymmetry. Additional tests provide evidence that these information effects are directly associated with contagion. The results facilitate a deeper understanding of contagion, and provide evidence that bankruptcy imposes broader industry-wide consequences than previously documented.

To Debora, Elena and Evan

ACKNOWLEDGEMENTS

I am grateful for the support and assistance of my dissertation committee: Robert Bushman (co-chair), Wayne Landsman (co-chair), William Beaver, Mark Lang and Edward Maydew. I also thank Jeff Abarbanell, Dan Amiram, Ray Ball, Ryan Ball, Mary Barth, Daniel Beneish, Judson Caskey, Craig Chapman, Hans Christensen, George Foster, Kenneth French, Eva Labro, Richard Lambert, Mark Maffett, Maureen McNichols, Richard Sansing, Katherine Schipper, Cathy Schrand, Abbie Smith, Clifford Smith, Ro Verrecchia, Charles Wasley, Chris Williams, Joanna Wu, Jerry Zimmerman, and workshop participants at Dartmouth, Duke University, Indiana University, Northwestern University, Stanford University, University of California Los Angeles, University of Chicago, University of North Carolina at Chapel Hill, University of Pennsylvania, and University of Rochester for valuable comments. I gratefully acknowledge funding from the Deloitte Doctoral Fellowship.

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1. Introduction

In this study, I seek a deeper understanding of bankruptcy contagion and its impact on the industry information environment by examining whether intra-industry bankruptcy affects the equity market response to surviving firms' subsequent earnings news. I use the term "contagion" to denote the phenomenon where a default by one firm increases the market's assessment of the default risk of other firms. I provide evidence that intra-industry bankruptcy diminishes the response to surviving firms' earnings news, consistent with the economic logic that increased default risk assessments lead to weaker stock price reactions to earnings information (e.g., Subramanyam and Wild, 1996; Barth, Beaver, and Landsman, 1998). I find that the decrease in response primarily reflects a diminished response to good news, i.e., there is negative asymmetry in the response in that good news is weighted less than bad news. Results are consistent with managerial preemptive disclosure of good news exacerbating, but not completely explaining, this asymmetry. Additional tests reveal that the results vary predictably with proxies for contagion intensity, which provides evidence that these information processing effects are directly associated with intra-industry bankruptcy contagion. The results of the study facilitate a richer understanding of contagion, and provide evidence that bankruptcy has broader industry-wide consequences than previously documented.

This study extends the literature that documents a number of contemporaneous cross-firm consequences of intra-industry bankruptcy. Lang and Stulz (1992) documents that when a firm announces Chapter 11 bankruptcy, there is a contemporaneous decrease in the value-weighted portfolio of surviving industry firms, consistent with a dominant on-average contagion effect. Jorion and Zhang (2007) finds that Chapter 11 bankruptcy announcements contemporaneously increase the spreads on credit default swaps of industry competitors, which provides clear evidence that intra-industry Chapter 11 bankruptcy increases the perceived default risk of surviving firms. Consistent with that evidence, Hertzel and Officer (2008) documents that corporate loan spreads are significantly larger when loan originations or renegotiations follow incidence of intra-industry bankruptcy. Importantly, that study documents that intra-industry bankruptcy has cross-firm credit market effects that extend beyond contemporaneous announcement effects. The key innovation of my study is the recognition that not only does bankruptcy contagion have direct cross-firm consequences for prices (e.g., stocks, loans, credit default swaps), but it also can fundamentally affect the information processing environment of an industry and alter the manner in which the equity market responds to subsequent news.

The underlying economics of default contagion are not well understood (e.g., Schönbucher, 2003; Das, Duffie, Kapadia, and Saita, 2007). Credit risk models that focus on the portfolio level typically rely on the assumption that default risk is uncorrelated across firms after controlling for common systematic drivers. However, empirical evidence rejects this assumption (e.g., Das et al., 2007). Several models have been proposed to explain excess default risk correlation across firms, including direct causal models of counterparty risk (e.g., Jorion and Zhang, 2009) and models based on learning-from-default. Learning-from-default models posit that an observed default event (e.g., intra-industry bankruptcy) causes investors to update their beliefs about unobservable variables that affect other firms' assessed default risk, which can explain excess observed correlation in credit spreads (e.g., Giesecke, 2004). I rely on insights from the learning-from-default model, and the updating of beliefs mechanism in particular, to formally develop the prediction that contagion in the wake of bankruptcy can affect subsequent information processing in an industry.

Prior research examines the effect of firm-specific default risk on the equity market response to earnings information. Using a value-relevance framework, Barth et al. (1998) finds that the weight investors place on own-firm earnings (book values) decreases (increases) during the period leading up to bankruptcy, consistent with a decreased focus on continuing earnings relative to liquidation values. Relatedly, several long-window association studies document that the market response to earnings news, i.e., the earnings response coefficient, is negatively related to a firm's default risk and probability of bankruptcy (Dhaliwal, Lee, and Fargher, 1991; Dhaliwal and Reynolds, 1994; Subramanyam and Wild, 1996). Accordingly, I hypothesize that if bankruptcy increases investors' assessment of default risk of surviving industry firms, investors will respond less to earnings news from such firms in the wake of intra-industry bankruptcy. Consistent with this hypothesis, I provide evidence that there is a significantly weaker market response to news in earnings announcements released in the wake of intra-industry bankruptcy relative to earnings announcements not preceded by intra-industry bankruptcy.

Interpretation of the on-average decrease in the response to bankruptcy-wake earnings news requires examination of possible negative asymmetry in the response, i.e., weaker response to good news relative to bad news, which can stem from two independent contagion-related sources. First, a class of economic theories involving model uncertainty (e.g., Epstein and Schneider, 2008) posits that when decision makers are faced with certain types of uncertainty, pessimistic updating of beliefs induces negative asymmetry in the response to information. To the extent such pessimistic behavior is descriptively valid, it is likely that it would manifest during periods of contagion. Second, contagion creates incentives for managers of surviving firms to preemptively disclose good news in response to intra-industry bankruptcy announcements. Any such increase in voluntary good news disclosure could result in an observed negative asymmetry in the response to earnings announcements, driven by contagion-induced changes in information flow rather than changes in information processing.

I provide evidence that intra-industry bankruptcy induces negative asymmetry in the market response to earnings news relative to the response to earnings news in the absence of intra-industry bankruptcy. Moreover, this asymmetry results primarily from a decreased response to good news in the wake of bankruptcy, with bankruptcy having little apparent effect on the response to bad news. Results are consistent with a combination of model uncertainty-induced asymmetry and an on-average reduction in earnings informativeness stemming from increased cross-firm default risk assessments. Results are further consistent with managerial preemptive disclosure of good news exacerbating, but not completely explaining, this asymmetry.

Additional tests reveal that the negative asymmetry in the response to earnings news released in the wake of intra-industry bankruptcy varies predictably based on proxies for contagion intensity. Negative asymmetry is more pronounced during periods where multiple intra-industry bankruptcies exist, and for firms that have a greater degree of economic comparability with their industry peers. Moreover, the effects of bankruptcy on the response to earnings news are insignificant for firms with relatively good financial health, consistent with good financial health mitigating contagion effects. The totality of evidence provides support for the interpretation that the cross-firm effects of intra-industry bankruptcy on the market response to news are directly associated with contagion, and is inconsistent with competing explanations.

In addition to the information transfer and default contagion literatures, this study contributes to the literature that examines how uncertainty in the firm- or market-level information environments affects the market response to firm-level earnings news. Skinner and Sloan (2002) finds evidence of negative asymmetry in the response to earnings news for glamour stocks, where the underlying source of uncertainty is the model of earnings for a particular firm. Conrad, Cornell and Landsman (2002) utilizes insights from regime-switching models (e.g., Veronesi, 1999) to provide evidence that uncertainty about the state of the macroeconomy induces negative asymmetry in the response to firm-level earnings news. Using a different theoretical underpinning (model uncertainty), Williams (2009) finds evidence of negative asymmetry in the response to firm-level earnings news following increases in macroeconomic uncertainty. My study provides evidence that industry-level bankruptcy contagion is an independent driver of asymmetry in the market response to firm-level earnings news.

The remainder of the paper is organized as follows. Section 2 discusses literature related to intra-industry bankruptcy contagion and develops the study's hypotheses. Section 3 discusses the study's empirical setting, including the bankruptcy-related partitioning of earnings announcements, and describes sample construction. Section 4 presents my research design. Section 5 presents and discusses the results of the study. Section 6 concludes.

2. Background and hypothesis development

2.1. Cross-firm consequences of intra-industry bankruptcy

Examination of the intra-industry consequences of bankruptcy is rooted in the information transfer literature (e.g., Foster, 1981). Bankruptcy-related information transfer studies generally distinguish between "contagion" effects and "competitive" effects, which describe whether the information that is transferred is "bad" or "good" for the non-announcing firms.[1] There are two common views of contagion in that literature (Lang and Stulz, 1992). The first view is that bankruptcy makes market participants wary of the financial health of other firms in the same industry, resulting in a contemporaneous decrease in firm values. The second view is that a bankruptcy portends financial distress at other firms with similar business models. In contrast, the competitive effect reflects the change in value of the non-announcing firm that is attributable to wealth redistribution from the bankrupt firm to the survivors. The observed event period stock returns of non-announcers reflect the sum of contagion and competitive effects.

Using a sample of large Chapter 11 bankruptcies, Lang and Stulz (1992) documents a contemporaneous decrease in the equity value of a value-weighted portfolio of industry competitors at the time of bankruptcy announcements, consistent with a dominant on-average contagion effect. In an extension of Lang and Stulz (1992), Ferris, Jayaraman, and Makhija (1997) likewise uses stock returns to provide evidence that the contagion effect is dominant among both large and small firm Chapter 11 bankruptcies, even among the set of firms with the highest likelihood of experiencing a competitive effect.

[1] Warner (1977) and Altman (1984) are credited with coining the terms "contagion" and "competitive" effects, respectively.

Jorion and Zhang (2007) examines contemporaneous intra-industry effects of Chapter 11 bankruptcy announcements on cross-firm credit default swap (CDS) spreads. A CDS is a credit derivative contract between two counterparties, where the buyer makes periodic payments to the seller and in return receives a payoff if a given firm defaults on the underlying debt. The CDS spread is the annual amount the buyer must pay to the seller, where a greater spread implies greater default risk, *ceteris paribus*. Focusing on CDS spread reactions as opposed to stock price responses allows a cleaner test of the effects of intra-industry bankruptcy on surviving firms' assessed default risk.[2] Jorion and Zhang (2007) finds that Chapter 11 bankruptcy announcements contemporaneously increase intra-industry CDS spreads, which provides evidence that Chapter 11 bankruptcies are predominantly contagion events.

Consistent with the results in Jorion and Zhang (2007), Hertzel and Officer (2008) finds that corporate loan spreads are significantly larger when loan originations or renegotiations follow incidence of intra-industry bankruptcy. Further, that study provides evidence that lenders are more likely to require collateral in the wake of intra-industry bankruptcy. Importantly, that study documents that bankruptcy contagion imposes debt market consequences that extend beyond contemporaneous bankruptcy announcement effects. Although extant literature provides convincing evidence that intra-industry Chapter 11 bankruptcy is a contagion-inducing event, that literature has not examined the effects of intra-industry bankruptcy on the subsequent processing of surviving firms' information by the equity market.

[2] As an example of the problems encountered in using stock price as an indicator of default risk, consider that an increase in firm leverage leads to greater credit risk, but can also create a wealth transfer to shareholders, in which case stock price appreciates.

2.2. Correlated default risk

A growing literature seeks to understand cross-firm default risk correlation (e.g., Das et al., 2007; Duffie, Eckner, Horel, and Saita, 2009). Standard credit risk models assume that default risk derives its randomness from a set of observable firm-specific (e.g., leverage, asset volatility) and macroeconomic (e.g., interest rates) state variables, and that firm-specific defaults are independent events after conditioning on these observable covariates (e.g., Duffie, Saita, and Wang, 2007). However, those models have been rejected based on empirical evidence of default risk correlation in excess of that predicted by observable covariates (e.g., Das et al., 2007).

A leading class of models that attempts to explain such excess default risk correlation posits a learning-from-default mechanism, where assessed default risk includes the expected effects of random unobservable variables (e.g., Jarrow and Yu, 2001; Schönbucher, 2003; Giesecke, 2004; Collin-Dufresne, Goldstein, and Helwege, 2008). This class of models provides a structured way of thinking about how intra-industry bankruptcy affects default risk assessments of surviving firms. In these models, the observance of a default causes, via Bayes' rule, a jump in the conditional distribution of the unobservable variables, with an associated jump in conditional default probabilities of any other firms whose default probabilities depend on the same unobservable covariates (e.g., firms in the same industry). Oft-cited examples of this are the defaults of Enron and WorldCom, which revealed faulty accounting practices that could have been in use at other firms. These defaults may have caused a sudden reduction in the perceived precision of accounting leverage measures of other firms, consequently affecting their conditional default probabilities.[3] A key implication of the learning-from-default model is that it suggests bankruptcy is an event that can cause investors to update default risk assessments for other firms, consistent with evidence in Jorion and Zhang (2007) and Hertzel and Officer (2008)

2.3. Default risk and the on-average response to earnings news

Prior literature examines the effect of firm-specific default risk on the market response to earnings information. Using a value relevance framework, Barth et al. (1998) documents that the valuation weight investors place on earnings (book values) progressively decreases (increases) over the five-year period directly preceding bankruptcy. That result is consistent with the idea that as default risk increases, investors view liquidation values as more relevant to firm valuation than future economic earnings. Relatedly, Subramanyam and Wild (1996) documents a negative relation between earnings response coefficients and probability of bankruptcy, consistent with the idea that earnings persistence is determined, in part, by the market's view of an entity's going-concern status. Using long window association studies, Dhaliwal et al. (1991) and Dhaliwal and Reynolds (1994) document empirically that earnings response coefficients are negatively related to default risk, consistent with the hypothesis that increases in default risk increase the discount rate applied to unexpected earnings.

These studies suggest that, if intra-industry bankruptcy increases the assessed default risk of surviving firms, it is likely that such contagion will weaken the equity market's response to intra-industry earnings news until implications of the bankruptcy for surviving

[3] Consistent with this, Yu (2005) provides empirical evidence that a reduction in the measured precision of accounting variables is associated with a widening of credit spreads.

firms are better understood. This logic suggests that the on-average market response to earnings news is weaker for earnings announcements that follow intra-industry bankruptcy (hereafter referred to as "bankruptcy-wake" earnings announcements) compared to those that do not (hereafter referred to as "non-bankruptcy-wake" earnings announcements). However, whereas the above-referenced studies provide evidence that earnings responses are affected by default risk as captured by observable firm-specific measures (e.g., leverage, bond ratings), it is unclear whether contagion-induced increases in perceived default risk will induce similar equity market effects. Accordingly, I test the following hypothesis, stated in alternative form:

> ***H_1:*** *The market response to earnings news is weaker for bankruptcy-wake earnings announcements than for non-bankruptcy-wake earnings announcements.*

2.4. Contagion and asymmetry in the response to earnings news

Previous studies that examine the relation between default risk and the market response to earnings information do not examine asymmetry in the responses to good versus bad earnings news. However, contagion raises the potential for negative asymmetry in the response to earnings announcements, i.e., a weaker response to good news relative to bad news, from two independent sources: contagion-induced model uncertainty, and contagion-induced managerial disclosure of good news. Interpretation of the effect of bankruptcy contagion on the response to intra-industry earnings news requires exploration of this potential asymmetry.

2.4.1. Contagion and model uncertainty

The learning-from-default model assumes that investors update their prior beliefs about variables that affect cross-firm conditional default probability assessments in Bayesian fashion. However, a class of economic theories involving model uncertainty (e.g., Hansen and Sargent, 2001; Epstein and Schneider, 2008) assumes that decision makers do not always update priors in accordance with Bayes' rule, and posits that pessimistic updating in the face of uncertainty induces negative asymmetry in the response to information.[4] In brief, investors reflect their pessimism by placing great weight on reported bad news and little weight on reported good news. It is plausible that, to the extent such pessimistic behavior is descriptively valid, it will manifest in the wake of contagion-inducing information shocks.

Intra-industry bankruptcy has the potential to introduce several forms of model uncertainty, each of which implies negative asymmetry in the response to earnings news released in the wake of bankruptcy. First, if investors are surprised by an intra-industry bankruptcy, investors may have concern that the accounting system of the bankrupt firm did not produce financial statements that revealed the impending financial distress. This may lead investors to question how precisely the accounting systems in the overall industry capture underlying firm economics, leading to uncertainty over the precision of earnings announcements (e.g., Yu, 2005). Second, intra-industry bankruptcy could make investors uncertain about the volatility in the fundamentals of surviving firms in the industry. For example, bankruptcy may make investors fear that the prior belief they held about fundamental volatility in the industry before the bankruptcy announcement was formed with

[4] Model uncertainty has been linked to numerous aspects of capital market behavior. Epstein and Wang (1994) proves the existence of equilibrium in an asset pricing model that admits a role for ambiguity. Anderson, Ghysels and Juergens (2009) provides empirical evidence that model uncertainty explains a substantial portion of equity returns.

too little consideration of downside risk. In Appendix A, I outline a model from Epstein and Schneider (2008) to illustrate how these types of model uncertainty can induce larger stock price responses for bad relative to good earnings news. Finally, bankruptcy may make investors uncertain about the persistence of surviving firms' reported earnings because of the bankruptcy's unknown future product or labor market implications. This would lead to a pessimistic slant towards models with high (low) persistence in the face of bad (good) news, likewise generating negative asymmetry in the response to earnings news (e.g., Hansen and Sargent, 2001).

2.4.2. Contagion and managerial disclosure incentives

A large literature examines managerial voluntary disclosure incentives and how market participants react to such disclosures (e.g., Patell, 1976; Penman, 1980; Baginski, Conrad, and Hassell, 1993; Skinner, 1994; Miller, 2002; and Kothari, Shu, and Wysocki, 2009). The seminal studies of Grossman (1981) and Milgrom (1981) suggest that managers will fully disclose all information. However, evidence suggests that voluntary earnings forecasts in advance of large earnings surprises are relatively rare (Kasznik and Lev, 1995), and that the frequency of voluntary disclosure is greater for bad news than for good news (Skinner, 1994). In general, the literature has relied on theories of disclosure costs (e.g., Verrecchia, 1983) and investor uncertainty about the possession or content of managers' private information (e.g., Dye, 1985; Jung and Kwon, 1988) to explain the absence of complete voluntary disclosure. In any event, economic logic suggests that managers will not voluntarily disclose news if the all-in cost of disclosure is greater than the all-in cost of non-disclosure. Because preemptive disclosure of good news is observed relatively infrequently, it follows that in many cases the cost of good news disclosure must outweigh the cost of non-disclosure.

Evidence suggests that bankruptcy contagion leads investors to reassess default risk at surviving firms with a broad brush, forcing all firms to bear contagion-related costs (e.g., Hertzel and Officer, 2008). Stated differently, it is likely that contagion increases the cost of non-disclosure of good news. Therefore, some firms that would not preemptively disclose good news in non-contagion equilibrium will find it advantageous to preemptively disclose good news to mitigate broad-brush contagion effects.[5] This relative increase in preemptive disclosure of good news during periods of bankruptcy contagion will result in an observed negative asymmetry in the response to news at the earnings announcement date, as the market response to good news will take place prior to the earnings announcement.[6] In this case, the observed change in earnings response in the wake of intra-industry bankruptcy would be attributable to a contagion-induced change in information flow, rather than a change in information processing.

Consideration of model uncertainty and managerial disclosure incentives in the wake of intra-industry bankruptcy leads to the following hypothesis, stated in alternative form:

> ***H_2:*** *There exists negative asymmetry, i.e., a weaker response to good news compared to bad news, in the response to bankruptcy-wake earnings announcements relative to non-bankruptcy-wake earnings announcements.*

[5] Using related logic, Sletten (2009) provides evidence that managers in firms that experience negative abnormal returns around an intra-industry restatement are more likely to disclose good news between the restatement and the firms' earnings announcement.

[6] As a recent example, in the wake of numerous Chapter 11 bankruptcy filings in the retail industry (e.g., Goody's on 6/9/2008, Steve & Barry's on 7/9/2008), Macy's, Inc. CEO Terry Lundgren filed a letter to Macy's executives with the SEC under Reg. FD on Friday, 7/11/08. The letter stated "On a quarter-to-date and year-to-date basis, our same-store sales trends are better than J.C. Penney, Kohl's, Dillard's, Nordstrom, Bon-Ton, The Gap and Limited Brands, to name a few….Our corporation is financially healthy. Our cash flow remains strong." Macy's shares jumped 3.8% on Monday, 7/14/08, while the broader retail market declined.

Because other forces may be at work in the information environment that affect the baseline symmetry in the response to earnings news, this hypothesis is stated in relative comparison to the baseline response to earnings news observed in the absence of intra-industry bankruptcy (e.g., if there exists negative asymmetry in the response to non-bankruptcy-wake earnings news, H_2 suggests that the negative asymmetry is stronger for bankruptcy-wake earnings news).

3. Empirical setting, sample selection and data

3.1. Empirical setting

My empirical strategy is to compare the market reaction to bankruptcy-wake earnings announcements vis-à-vis the corresponding reaction for non-bankruptcy-wake earnings announcements. Accordingly, a key ingredient of my setting is the definition of "bankruptcy-wake." Given the lack of theory or empirical evidence concerning post-announcement effects of intra-industry bankruptcy, it is unclear how long any information processing effects will persist. A relatively short window is necessary to minimize the incidence of confounding events between a bankruptcy and a particular earnings announcement under study. Conversely, the definition of bankruptcy-wake must be broad enough to achieve a reasonable sample size of bankruptcy-wake earnings announcements. With these competing considerations in mind, I classify a given earnings announcement as a bankruptcy-wake announcement if there was at least one intra-industry bankruptcy during the thirty calendar days immediately preceding the earnings announcement date, where industries are defined according to the Fama-French 48 classification (Fama and French, 1997).[7] In complementary fashion, I classify an earnings announcement as a non-bankruptcy-wake announcement if

[7] Experimentation with variations in the length of this window ranging from 10 to 45 days does not change the tenor of the primary results reported below.

there were no intra-industry bankruptcies during the thirty calendar days immediately preceding the earnings announcement date.

3.2. *Data and sample selection*

I compile a broad sample of Chapter 11 bankruptcy filings from SDC Platinum, which contains data for Chapter 11 bankruptcy filings beginning in 1980. Specifically, I begin with all initial firm bankruptcy filing observations in SDC where the filing date is between January 1, 1980 and June 30, 2007. I then delete all bankruptcy filings for which I cannot obtain an industry identifier from Compustat (*dnum*), leaving a sample of 1,873 Chapter 11 bankruptcies.[8]

I obtain all quarterly earnings announcements from Compustat that have a valid CRSP-Compustat link and a non-missing earnings announcement date in Compustat (*rdqe*). Using the bankruptcy sample, I delete all earnings announcements for bankruptcy filers that occur after their bankruptcy filing date. I further delete all earnings announcements where the three-day period centered on the earnings announcement date overlaps with any three-day period centered on an intra-industry bankruptcy announcement. Next, I categorize each earnings announcement as either a bankruptcy-wake earnings announcement (hereafter denoted $BW_{i,t}$ = 1), or a non-bankruptcy-wake earnings announcement (hereafter denoted $BW_{i,t}$ = 0).[9] I delete all observations for firms that are not present for at least one firm-quarter observation in both the BW = 0 and BW = 1 partitions, so that the partitions are comprised of identical firm sets.

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[8] I link SDC to CRSP using both the CRSPQ and CUSIP master files. If an exact name match does not exist, I manually search for matches using the CUSIP lookup function on the CRSP-Compustat merged database. I also delete some observations where the accuracy of the SDC filing date is in question, after failed attempts to validate the suspect dates via www.bankruptcydata.com. Additional details are available upon request.

[9] The subscript *i* indexes firm, and the subscript *t* indexes quarter. Hereafter, I frequently suppress the *i,t* subscript when referring to variables measured at the firm-quarter level for expositional simplicity.

Unless otherwise noted, accounting and stock price data used in the study are obtained from Compustat and CRSP, respectively (hereafter, *data##* refers to a Compustat variable). I delete observations without data necessary to compute all variables used in my empirical specifications, as well as several control and partitioning variables. Specifically, I require each firm-quarter observation to have non-missing data for seasonally differenced quarterly earnings per share before extraordinary items (*data19*), fiscal quarter-end stock price (*data14*), fiscal quarter-end common shares outstanding (*data61*), book value of common equity (*data59*), total liabilities (*data54*), CRSP stock price two trading days prior to the earnings announcement date, CRSP return for the three trading days centered on the earnings announcement date, and the level of the Chicago Board Options Exchange Volatility Index (VIX) two trading days prior to the earnings announcement date.[10] I further require observations in the $BW = 1$ partition to have non-missing CRSP return for the three trading days centered on the announcement dates of intra-industry bankruptcy filings that occur in the thirty-day window preceding the earnings announcement. Also, I delete observations for which seasonally differenced earnings per share before extraordinary items equals zero, i.e., "no news" observations.

To minimize the effects of market frictions, I delete observations having stock price less than $1 two trading days prior to the earnings announcement date (Ball, Kothari and Shanken, 1993). Following prior literature (e.g. Collins, Maydew and Weiss, 1997; Kothari and Shanken, 2003), I delete observations with negative book value of equity to avoid capturing effects that may be related to extreme financial distress. Finally, I truncate test and

[10] VIX, often referred to as the "fear index," represents one measure of the market's expectation of volatility over the next thirty day period. I obtain daily VIX data from http://www.cboe.com/micro/VIX/historical.aspx. Because VIX data are unavailable prior to 1986, my sample period is so constrained.

control variables defined below at the 1st and 99th percentiles to reduce the effects of database errors and outlier observations. Experimentation with variations in these data restrictions does not change the tenor of the results reported below. The final sample consists of 274,658 firm-quarter earnings announcement observations with report dates that range from January 2, 1986 through June 20, 2007, where 44,987 earnings announcements are bankruptcy-wake earnings announcements (BW = 1), and 229,671 are not (BW = 0). Underlying the BW = 1 earnings announcements are 1,585 distinct Chapter 11 bankruptcy announcements with filing dates that range from December 6, 1985 through June 11, 2007. Of the 1,585 bankrupt firms, 323, 202, and 1,060 were last listed on the NYSE, AMEX, and NASDAQ, respectively.

3.3. Descriptive statistics

Figure 1 presents the frequency of sample bankruptcies by calendar year. Consistent with prior literature (e.g., Beaver, McNichols, and Rhie, 2005), bankruptcies are relatively frequent in the years surrounding 1990 and 2000. Figure 2 presents the distribution of sample bankruptcies by Fama-French 48 industry classification, and reveals that the bankruptcy sample spans a wide range of industries.[11] Consistent with prior research, the sample includes a large number of bankruptcies in the retail and business services industries (e.g., Chava and Jarrow, 2004).

Panels A and B of Table 1 present descriptive statistics for the non-bankruptcy-wake (BW = 0) and bankruptcy wake (BW = 1) earnings announcements, respectively. Interestingly, although the BW = 0 and BW = 1 partitions are comprised of the identical set of firms, there are statistically significant differences in mean values of most firm-level variables. However, most differences appear to be economically insignificant. The overall

[11] Industries not represented are Candy & Soda (FF48 category 3) and Tobacco Products (FF48 category 5).

distributional statistics reveal that the two partitions are quite similar along most firm-level dimensions. However, there is a notable difference across partitions for *VIX*. In particular, the *VIX* distribution is shifted to the right for the $BW = 1$ partition relative to the $BW = 0$ partition (e.g., median of 22.5 versus 18.8), suggesting a positive association between incidence of bankruptcy and market-level uncertainty.

4. Research design

To test the hypothesis that there is a weaker on-average response to news in bankruptcy-wake earnings announcements as compared to that for non-bankruptcy-wake earnings announcements (H_1), I estimate the following regression model via ordinary least squares:

$$CAR_{i,t} = \alpha_0 + \alpha_1 BW_{i,t} + \alpha_2 NEWS_{i,t} + \alpha_3 BW * NEWS_{i,t} + \sum \alpha_k CONTROL_{i,t} + \gamma_{i,t}, \tag{1}$$

where *BW* is an indicator variable that equals one if firm *i*'s earnings announcement in quarter *t* is a bankruptcy-wake earnings announcement and equals zero otherwise, *NEWS* is seasonally differenced quarterly earnings per share before extraordinary items (*data19*) scaled by firm *i*'s stock price two trading days prior to the earnings announcement date (Christie, 1987), and *CAR* is the cumulative abnormal return, i.e. daily return less the CRSP value-weighted return, for firm *i* during the 3-day trading window centered on its quarter *t* earnings announcement date.[12] A significantly negative α_3 coefficient would provide evidence, consistent with H_1, that there is a smaller on-average response to earnings news for bankruptcy-wake earnings announcements.

To test the hypothesis that there is negative asymmetry in the stock price response to earnings news for bankruptcy-wake earnings announcements relative to non-bankruptcy-

[12] As discussed in Section 5.5., results are robust to alternative definitions of *NEWS* and *CAR*.

wake earnings announcements (H_2), I estimate the following regression, which allows for a differential response to bad versus good earnings news, via ordinary least squares:

$$\begin{aligned} CAR_{i,t} = \beta_0 + \beta_1 BAD_{i,t} + \beta_2 NEWS_{i,t} + \beta_3 BAD * NEWS_{i,t} + \beta_4 BW_{i,t} + \beta_5 BW * BAD_{i,t} \\ + \beta_6 BW * NEWS_{i,t} + \beta_7 BW * BAD * NEWS_{i,t} + \sum \beta_k CONTROL_{i,t} + \varepsilon_{i,t}, \end{aligned} \tag{2}$$

where *BAD* is an indicator variable that equals one if *NEWS* is less than zero and equals zero otherwise, and all other variables are as previously defined. This specification permits the slope on *NEWS* to differ conditional on the sign of earnings news, and conditional on whether earnings news is released in the wake of intra-industry bankruptcy. The test of H_2 focuses on the sign and significance of β_7, which estimates the effect of bankruptcy on the incremental response coefficient for bad news versus good news. That is, β_7 estimates directly the degree of negative asymmetry in the response to earnings news introduced by intra-industry bankruptcy. Finding a significantly positive β_7 coefficient would provide evidence, consistent with H_2, that bankruptcy induces negative asymmetry in the stock price response to intra-industry earnings news, relative to the baseline level of asymmetry for non-bankruptcy wake earnings news.[13]

I include several control variables in all specifications. First, following Conrad et al. (2002), I include size (*SIZE*), measured as the natural log of the quarter-end market value of common equity (*data14**data61*), to control for risk differences not reflected in abnormal return (Fama and French, 1993) and for potential scale differences (Barth and Kallapur, 1996). I also include leverage (*LEV*), measured as total liabilities divided by book equity (*data54*/*data59*), to further control for the risk of financial distress and leverage effects (e.g.,

[13] In Appendix B, I outline the empirical modeling through which I determine the necessary interaction variables to include in this specification.

Christie, 1982). Next, I include the cumulative abnormal return, i.e., daily firm return less the CRSP value-weighted market return, for firm *i* over the period beginning two trading days after the report date of earnings for quarter *t*-1 and ending two trading days prior to the report date of earnings for quarter *t* (*PRECAR*) as a control for information associated with recent events prior to earnings announcement that may affect measurement error in *NEWS* (Brown, Griffin, Hagerman, and Zmijewski, 1987). When estimating the asymmetric specification in Eq. (2), I also include an indicator variable, *BADPRE*, that equals one if *PRECAR* is less than zero and equals zero otherwise, and the associated interaction *BADPRE*PRECAR*, to accommodate the possibility that investors react asymmetrically to news received in the pre-earnings announcement period.

I control for the effects of macroeconomic uncertainty throughout my analyses. Otherwise, any effects I attribute to intra-industry bankruptcy may instead be the result of a higher level of uncertainty present in the macro-environment. The necessity of this control is apparent for two reasons. First, Table 1 reveals a correlation between the incidence of bankruptcy and macro-level uncertainty, as captured by *VIX*. Second, Williams (2009) provides evidence that increases in VIX (that study's proxy for macroeconomic uncertainty) are associated with negative asymmetry in the stock price response to earnings news, where the effect is particularly strong during periods of high VIX. I construct an indicator *HVIX* that equals one if *VIX* is above the sample median, and equals zero otherwise, where *VIX* is the level of VIX two trading days prior to firm *i*'s earnings announcement date. I include *HVIX* (and associated interactions with *NEWS*) in any specification where *BW* (and associated interactions with *NEWS*) enters the regression model, to control for both the mean effect of the level of uncertainty in the macro-environment and its effect on the response to earnings news. Finally, I include year fixed effects based on earnings announcement dates to control for time clustering of bankruptcies, and cluster standard errors at the firm level (Petersen, 2009).

When testing alternative explanations and cross-sectional effects, I use either control variables or sample partitioning schemes. When partitioning structures permit, I test the significance of coefficient differences across partitions using untabulated four-way interaction specifications. Otherwise, I employ Monte Carlo randomization tests. I identify the method used when I discuss the associated test results. I present a description of the Monte Carlo randomization test methodology in Appendix C.

5. Empirical analyses and results

5.1. On-average response to bankruptcy-wake earnings news

Table 2 presents the results of estimating Eq. (1). Consistent with the large earnings response literature, there is a strong positive relation between earnings news and the cumulative abnormal return over the earnings announcement window, as shown in model (1). Model (2) reveals that when I allow the response to news to vary based on bankruptcy-wake partitioning, there is a significant negative coefficient (–0.060) on the interaction variable *BW*NEWS* (t-stat of –6.18). That is, the on-average response to earnings news is weaker for bankruptcy-wake earnings announcements than for non-bankruptcy-wake earnings announcements. Model (3) documents that high levels of macroeconomic uncertainty are associated with a significantly diminished response to earnings news, as evidenced by the significant negative coefficient on *HVIX*NEWS*. Critically, the coefficient on *BW*NEWS* remains strongly negative (–0.048 with a t-stat of –4.80), which provides evidence that intra-industry bankruptcy contagion effects are incremental to those related to overarching macroeconomic uncertainty.

These results provide evidence consistent with H_1, and suggest that there is an increase in the assessed default risk for surviving firms in the wake of intra-industry bankruptcy of sufficient magnitude to affect the equity market information processing environment. However, these results do not reveal whether the diminished response to earnings news results from symmetrically decreased responses to both good and bad news, or reflects the net effect of asymmetry in the response to good news relative to bad news.

5.2. Asymmetry in the response to bankruptcy-wake earnings news

Table 3 presents results of estimating Eq. (2). As reported in model (1), there exists a significant negative asymmetry in the full sample of earnings announcements, as indicated by the positive coefficient of 0.028 on *NEWS*BAD* (t-stat of 2.84). However, model (2) provides evidence that there is a symmetric response to bad versus good earnings news for non-bankruptcy-wake earnings announcements, as indicated by the insignificant coefficient (0.0046) on *NEWS*BAD* (t-stat of 0.42). In sharp contrast, consistent with H_2 there exists a marked negative asymmetry in the response to bankruptcy-wake earnings announcements, as indicated by the significantly positive coefficient on *BW*BAD*NEWS* of 0.090 (t-stat of 4.13).

Interestingly, the negative asymmetry appears to result almost entirely from a diminished response to good news for bankruptcy-wake earnings announcements, as reflected by the −0.096 coefficient on *BW*NEWS* (t-stat of −5.66). The total incremental bad news coefficient for bankruptcy-wake earnings announcements of −0.006 is insignificantly different from zero (p-value of 0.66), suggesting that the response to bad news has an equal magnitude for both bankruptcy-wake and non-bankruptcy wake earnings announcements. Model (3) reports results after inclusion of controls for market-level uncertainty. Inferences are unaltered relative to those obtained from model (2). There is a significantly negative coefficient on *HVIX*NEWS* of −0.027 (t-stat of −1.85), which indicates that the response to good news is further suppressed in periods of high macroeconomic uncertainty. However, the insignificant coefficient on *HVIX*BAD*NEWS* of 0.009 provides evidence that, after controlling for asymmetry induced by intra-industry bankruptcy, macroeconomic uncertainty does not induce asymmetry in the market response to firm-level bad versus good earnings news.

These results appear partially inconsistent with theories of model uncertainty. Whereas model uncertainty suggests that negative asymmetry arises from *both* a decreased response to good news *and* an increased response to bad news, I find a decrease in the good news response in the wake of bankruptcy with no apparent effect on the bad news response. However, this interpretation does not consider the potential co-existence of downward pressure on earnings news in general because of the on-average effect of increased default risk on earnings responses (e.g., Barth et al. 1998). It is possible that contagion-induced model uncertainty causes both a decreased response to good news *and* an increased response to bad news, which are then each shifted downward as a result of an on-average decrease in investor sensitivity to earnings news arising from the increase in perceived default risk. Although the observed results are consistent with that combination of forces, my research design cannot disentangle these interpretations.

5.3. Managerial preemptive good news disclosure

To test whether the observed negative asymmetry is related to voluntary managerial disclosure of good news in the wake of intra-industry bankruptcy, I create a proxy for the likelihood of such disclosure using the magnitude and direction of firm-level abnormal stock return in the window between bankruptcy announcement and earnings announcement. Specifically, for each bankruptcy-wake earnings announcement, I compute the earnings announcer's cumulative market-adjusted abnormal return between the end of the relevant three-day bankruptcy announcement window and the beginning of the three-day earnings announcement window (*DOBEA_CAR*). I then separately estimate Eq. (2) for the non-bankruptcy-wake earnings announcements pooled with each of two sub-partitions of the *BW* = 1 earnings announcements: the first quartile of *DOBEA_CAR*, i.e., *DOBEA_CAR* less than –0.040, and the fourth quartile of *DOBEA_CAR,* i.e., *DOBEA_CAR* greater than 0.037. If managers issue voluntary good news disclosures in the wake of intra-industry bankruptcy that elicit a positive market response of sufficient magnitude to suppress the response to good news at subsequent earnings announcement, it is unlikely (most likely) that any such disclosures are issued for observations within the first (fourth) quartile of *DOBEA_CAR*.

As reported in Table 4, there exists pronounced negative asymmetry in the response to earnings news for earnings announcements released in both the first and fourth quartiles of *DOBEA_CAR* ($\hat{\beta}_7$ of 0.076 and 0.173, with t-stats of 2.04 and 4.69, respectively). Monte Carlo randomization tests indicate that the negative asymmetry is greater in the fourth quartile than in the first quartile (one-tailed p-value of 0.029), consistent with contagion-induced preemptive disclosure of good news exacerbating negative asymmetry in the response to bankruptcy-wake earnings news. However, because there exists significant negative asymmetry in the first *DOBEA_CAR* quartile (where preemptive disclosure of good news is least likely to have occurred), preemptive disclosure of good news does not appear to be a complete explanation for the on-average asymmetry.

5.4. Additional analyses based on contagion intensity

5.4.1. Industry health

Within the $BW = 1$ earnings announcements, there exists variation in the number of intra-industry bankruptcies within the thirty-day window preceding a given earnings announcement. In this section, I exploit this variation to examine whether my results are related to the broader health of a given industry. Lang and Stulz (1992) provides evidence that contagion effects are intensified during periods of multiple intra-industry bankruptcies, consistent with the interpretation that during such times, bankruptcy is more likely caused by industry-wide shocks rather than idiosyncratic factors. If the information processing effects I document are associated with contagion, I therefore expect that the negative asymmetry in the response to bankruptcy-wake earnings news is exacerbated when earnings announcements follow multiple intra-industry bankruptcies. To test this prediction, I estimate Eq. (2) separately for the $BW = 0$ earnings announcements pooled with each of two sub-partitions of the $BW = 1$ earnings announcements. The first (second) sub-partition contains earnings announcements released in the wake of single (multiple) intra-industry bankruptcies ($NBANK = 1$ and $NBANK > 1$, respectively).

Table 5 reports a significant reduction in the response to news for earnings announcements released in the wake of both single and multiple intra-industry bankruptcies ($\hat{\beta}_6$ of −0.058 and −0.164, with t-stats of −2.96 and −5.61, respectively). Table 5 further reveals significant negative asymmetry in the response to earnings news released in the wake of both single and multiple intra-industry bankruptcies ($\hat{\beta}_7$ of 0.054 and 0.167, with t-stats of 2.12 and 4.47, respectively). These results provide evidence that the bankruptcy effects on the market response to earnings news obtain in periods of both relatively good and poor industry health. Importantly, both the decrease in the response to news and the negative asymmetry in the *NBANK* > 1 sub-partition are significantly greater than those for the *NBANK* = 1 sub-partition, as determined by Monte Carlo randomization tests (one-tailed p-values of 0.001 and 0.011, respectively). This finding reinforces the contagion-based explanation, and is consistent with evidence in Lang and Stulz (1992) that contagion effects are more pronounced when there are multiple industry bankruptcies.

5.4.2. Intra-industry economic comparability

Research that investigates intra-industry effects of bankruptcy is based on the premise that any such effects will most likely be felt by related firms, where industry membership is used as a proxy for relatedness. However, there exists variation in economic comparability across firms within a given industry. Lang and Stulz (1992) and Jorion and Zhang (2007) provide evidence that contagion is most pronounced when industry peers have relatively high economic comparability, i.e., similar investments and cash flow patterns, with the bankrupt firm. Similarly, I expect industry-level contagion effects to be more pronounced for firms with tighter economic connections to their industry peers. To test whether my results vary predictably with economic comparability, I follow Morck, Yeung and Yu (2000) and compute a stock return synchronicity measure from the following time-series regression, estimated by firm-quarter using data from the 36 months immediately prior to firm *i*'s fiscal quarter *t* ending date:

$$RET_{i,m} = \gamma_0 + \gamma_1 VWFFRET_m + \nu_{i,m}, \tag{3}$$

where $RET_{i,m}$ is the return for firm i in month m obtained from the CRSP monthly file, and $VWFFRET_m$ is the value-weighted return on firm i's Fama-French 48 industry category in month m.[14] My measure of economic comparability is the adjusted-R^2 obtained from estimating Eq. (3), where a large adjusted-R^2 reflects high economic comparability. I then separately estimate Eq. (2) for two earnings announcement sub-partitions: those with adjusted-R^2 below the sample median, and those with adjusted-R^2 above the sample median.[15]

As reported in Table 6, there is significant bankruptcy-induced negative asymmetry in the response to earnings news for firms with both low and high within-industry economic comparability ($\hat{\beta}_7$ of 0.057 and 0.128, with t-stats of 1.84 and 3.87, respectively). More critically for this analysis, tests based on a four-way interaction specification provide evidence that the negative asymmetry for the high economic comparability observations is greater than that for the low economic comparability observations (one-tailed p-value = 0.055).

5.4.3. Surviving firm financial health

Because contagion effects relate to cross-firm default risk reassessments, it is likely that contagion effects are mitigated for surviving firms in relatively good financial health. To test this conjecture, I estimate Eq. (2) separately for sub-partitions of earnings announcements based on the earnings announcers' own financial health.

[14] I obtain the monthly returns for the Fama-French 48 industry groups from Kenneth French's website at http://mba.tuck.dartmouth.edu/pages/faculty/ken.french/data_library.html.

[15] I remove firm-quarter observations with fewer than twelve monthly observations available to estimate Eq. (3), reducing my overall sample size to 265,245 for this analysis.

To measure financial health, I loosely follow Piotroski (2000) and give each of three firm-level fundamental signals an indicator equal to one (zero) if the signal is "good" ("bad"), and then sum the indicators to create an aggregate financial health measure (*HSCORE*), where higher scores correspond to better financial health. Specifically, I use the following three variables: *CFO*, defined as cash flow from operations (*data108*); *LEVER*, defined as long-term debt (*data51*) divided by total assets (*data44*); *LIQUID*, defined as current assets (*data40*) divided by current liabilities (*data49*). I code *HCFO* equal to one if *CFO* is greater than zero, and code *HLIQUID* (*HLEVER*) equal to one if the associated firm-quarter signal realization is greater than (less than) the industry-fiscal-quarter median signal realization, and equal to zero otherwise. I then compute *HSCORE* as the sum of *HCFO*, *HLIQUID*, and *HLEVER*. Accordingly, *HSCORE* ranges from 0 to 3, where higher *HSCORE* indicates better health.

As reported in Table 7, there is a significant reduction in the response to news for bankruptcy-wake earnings announcements for the sets of firms with low and medium financial health ($\hat{\beta}_6$ of −0.072 and −0.102, with t-stats of −2.54 and −3.21, respectively), as categorized by the partitions *HSCORE* = {0, 1} and *HSCORE* = {2}. Table 7 further reveals significant negative asymmetry in the response to bankruptcy-wake earnings news for these firms ($\hat{\beta}_7$ of 0.077 and 0.094, with t-stats of 2.12 and 2.27, respectively). In sharp contrast, there is neither a reduction in the response to news nor negative asymmetry in the response for bankruptcy-wake earnings announcements of firms with high financial health (*HSCORE* = {3}). Specifically, the coefficients on *BW*NEWS* and *BW*BAD*NEWS* are insignificantly different from zero (t-stats of −1.50 and 0.04, respectively). These results provide evidence that the bankruptcy effects on the market response to intra-industry earnings news are mitigated for firms with relatively good financial health, consistent with a diminished contagion effect. It is also noteworthy that the response to news is monotonically increasing in financial health, consistent with the logic that investors place less weight on earnings the poorer is a firm's financial health (Barth et al., 1998).

5.5. Alternative explanations for bankruptcy-wake negative asymmetry

5.5.1. Contagion-induced earnings management

Managers of surviving firms may have incentives to inflate reported earnings in the wake of intra-industry bankruptcy in an attempt to offset pessimistic market assessments of default risk. Of course, such behavior would reflect an assumption on the part of managers that the market cannot see through earnings management (e.g., Xie, 2001). However, DeFond and Park (2001) concludes that market participants adjust for suspected earnings management at the time of an earnings announcement. Specifically, that study finds that the market response to good (bad) news is stronger for firms reporting income-decreasing (income-increasing) accruals versus income-increasing (income-decreasing) accruals. Those results imply negative asymmetry in the response to earnings news that contains income-increasing accruals, *ceteris paribus*. If investors expect managers to inflate earnings in the wake-of intra-industry bankruptcy, investors may further discount reported good news that is arrived at through income-increasing accruals, leading to an exacerbated negative asymmetry in the response to such earnings news.

To examine whether investor perception of earnings management is a plausible explanation for the asymmetry documented in Table 3, I incorporate a simple measure of upwards earnings management into the analysis. I first compute firm-quarter accruals (*ACC*) by subtracting cash flow from operations (*data108*) from net income (*data69*). I then create an indicator variable, *POSACC*, which equals one if *ACC* is greater than zero and equals zero otherwise.[16] Table 3 indicates that there is a similar proportion of bad news earnings announcements across the *BW* = 0 and *BW* = 1 partitions (0.44 and 0.46, respectively). Moreover, untabulated tests reveal that the proportion of earnings announcements that include upwards earnings management is statistically indistinguishable across the *BW* = 0 and *BW* = 1 partitions (0.304 and 0.307, respectively). Therefore, univariate evidence suggests that managers of surviving firms do not engage in upwards earnings manipulation in the wake of intra-industry bankruptcy. Untabulated results further confirm that inferences are unaltered after adding *POSACC* to Eq. (2) in an interactive fashion, i.e., *POSACC*, *POSACC*BAD*, *POSACC*NEWS*, and *POSACC*BAD*NEWS*.[17]

5.5.2. Torpedo effect

Skinner and Sloan (2002) provides empirical evidence that investors correct overly optimistic expectations for growth stocks when they see negative surprises in earnings announcements, which induces a negative asymmetric response to earnings news (hereafter referred to as the "torpedo effect"). If incidence of intra-industry bankruptcy makes it more likely that investors correct overly optimistic growth expectations for surviving firms, i.e., if the torpedo effect is more pronounced in the wake of intra-industry bankruptcy, asymmetry in the response to bankruptcy-wake earnings news could be related to torpedo effect dynamics.

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[16] Because statement of cash flows data are not available until 1987, for this test my sample is so constrained.

[17] As an alternate proxy for investor perception of earnings management, I replace *POSACC* with *AMACC*, where *AMACC* = 1 if ACC_t is above the sample median ACC for the firm's industry during the fiscal quarter for which earnings are reported, and = 0 otherwise. Untabulated results confirm that inferences are unchanged when I replace *POSACC* with the alternate proxy *AMACC*.

To examine whether my results are related to the torpedo effect, I follow Skinner and Sloan (2002) and control for market-to-book ratio (*MTB*) as a proxy for growth, where I define *MTB* as quarter-end market value of common equity divided by book value of common equity (*data14*data61/data59*). I then estimate Eq. (2) after adding *MTB* in an interactive fashion, i.e., *MTB*, *MTB *BAD*, *MTB *NEWS*, and *MTB *BAD*NEWS*. Untabulated findings confirm that all previously reported inferences are robust to controls for the torpedo effect.

5.5.3. Aggregate sentiment

Extending logic in Baker and Wurgler (2006), Mian and Sankaraguruswamy (2007) finds that the stock price response to good (bad) earnings news that arrives during a high (low) aggregate sentiment period is stronger than the stock price response to good (bad) earnings news that arrives during a low (high) sentiment period. Although the aggregate sentiment literature does not directly predict a stronger stock price response to bad versus good news, it does imply that the response to bad news minus the response to good news, i.e., negative asymmetry, is greater in times of low sentiment than in times of high sentiment. Therefore, if a substantial proportion of bankruptcies occur in times of low sentiment, sentiment could explain my findings.

To control for aggregate sentiment, I use the Baker and Wurgler (2006) sentiment index for the month during which firm *i* announced earnings for quarter *t* (*SEN*).[18] I then estimate Eq. (2) after adding *SEN* in an interactive fashion, i.e., *SEN*, *SEN*BAD*,

[18] Please refer to Baker and Wurgler (2006) for details on computation of this monthly sentiment index. I obtain monthly sentiment index data from http://pages.stern.nyu.edu/~jwurgler/data/. Because the sentiment index is only available through December 2005, the sample for this analysis is so constrained.

*SEN*NEWS*, and *SEN*BAD*NEWS*. Untabulated findings confirm that previously reported inferences are robust to controls for investor sentiment.

5.6. Robustness tests

5.6.1. Earnings news and abnormal returns

It is unlikely that my results are driven by differential measurement error in *NEWS* across *BW*=0 and *BW*=1 partitions. First, as discussed above, I include controls for information arrival in the quarter prior to earnings announcement (*PRECAR*) in all empirical specifications (Brown et al., 1987). Second, the overall pattern of results makes it difficult to attach a measurement error interpretation. However, to further address this concern I examine whether my primary results are robust to an alternative definition of earnings news based on analyst forecasts.[19] Specifically, I compute earnings news as quarterly I/B/E/S actual earnings less I/B/E/S consensus forecast for firm *i* in quarter *t*, where I compute consensus forecast as the mean of the most recent individual analyst forecasts issued during the 90-day window prior to earnings announcement.[20] Further, I only include bankruptcy-wake observations where the most recent analyst forecast included in the consensus calculation was issued *after* the associated bankruptcy announcement. These restrictions result in a severe reduction of my sample size to 91,375 firm-quarter observations, only 6,866 of which are in the *BW* = 1 partition. Nonetheless, I find results consistent with the main findings reported above. Specifically, there is a significant on-average decrease, as well as significant

[19] Of course, problems exist with using analyst forecasts as a measure of investor earnings expectation, including analyst propensity to issue optimistically biased forecasts (e.g., Francis and Philbrick, 1993) and analyst herding (e.g., Trueman, 1994).

[20] Because there exists a regime shift in I/B/E/S data around 1991-1992 (Abarbanell and Lehavy, 2007), I begin my sample period for this test in 1993.

negative asymmetry, in the response to bankruptcy-wake earnings news when news is measured using analyst consensus forecasts.

One advantage of the short window tests used in this study is that, because daily expected returns are close to zero, the particular model I use for expected returns does not have a large effect on inferences about abnormal returns (Fama, 1998). However, to investigate the robustness of my results to an alternative measure of event window abnormal returns, I estimate a version of the market model used in Collins, Li, and Xie (2009), which would capture any shifts in equity beta that might be induced by periods of contagion. Specifically, I estimate market model abnormal returns (*MMCAR*) as the three-day cumulative market model residual centered on the earnings announcement date (day 0), where the market model parameters are estimated by regressing daily firm return on the value-weighted market return over the 80-day non-announcement period where day (d) varies from -45 to -6 and from day +6 to +45 relative to the quarterly earnings announcement date. Consistent with the logic in Fama (1998) that the choice of the expected return model used in short window studies is relatively unimportant, the correlation (untabulated) between *CAR* and *MMCAR* is 0.97. More importantly, all inferences in the study are unchanged when I use *MMCAR* in place of *CAR*.

5.6.2. Sub-period analysis

Figure 3 plots the coefficients on good news ($\hat{\beta}_2$) and bad news ($\hat{\beta}_2 + \hat{\beta}_3$) obtained from estimating model (1) in Table 3 by calendar year of earnings announcement. Interestingly, there appears to be a downward structural shift in the magnitude of both coefficients in 2001 that is particularly pronounced for good news, followed by a period of prominent negative asymmetry. In this section I consider whether my primary results are affected by this structural shift.

Untabulated results confirm that my primary results separately obtain in both the sub-period 1986-2001 and 2002-2007. That is, there is a significant on-average decrease, as well as significant negative asymmetry, in the response to bankruptcy-wake earnings news relative to non-bankruptcy-wake earnings news in both sample sub-periods. Further, I confirm that my primary inferences are unaltered when I exclude 2001 data from the sample.

5.6.3. Retail and business services industries

Consistent with prior literature, my sample includes a large number of bankruptcies from the business services and retail industries. Specifically, these two industries together account for 14,039 bankruptcy-wake earnings announcements (31% of all *BW*=1 observations), but account for only 25,282 non-bankruptcy wake earnings announcements (11% of all *BW*=0 observations). To ensure that my results are not attributable to the disproportionate influence of these industries on the *BW*=1 partition, I repeat my main tests after exclusion of all observations in the business services and retail industries, and confirm that my inferences are unaltered.

6. Conclusion

I provide novel evidence that intra-industry bankruptcy contagion affects the equity market response to news in surviving firms' subsequent earnings announcements, where I use the term "contagion" to denote the phenomenon where a default by one firm increases the market's assessment of the default risk of other firms. First, I hypothesize and find that intra-industry bankruptcy diminishes the response to surviving firms' earnings news, consistent with the economic logic that increased default risk assessments lead to weaker stock price reactions to earnings information (e.g., Subramanyam and Wild, 1996; Barth, Beaver, and Landsman, 1998).

Next, I investigate whether the effect of bankruptcy contagion on the response to intra-industry earnings news is asymmetric across bad and good news realizations. Interestingly, I find strong evidence that bankruptcy induces negative asymmetry in the response, i.e., a weaker response to good news relative to bad news, which appears to be driven primarily by a reduction in the response to good news. Results are consistent with a combination of model uncertainty-induced asymmetry and an on-average reduction in earnings informativeness stemming from increased perceived default risk. Results are further consistent with managerial preemptive disclosure of good news in the wake of intra-industry bankruptcy exacerbating, but not completely explaining, this asymmetry. I present evidence that the results vary predictably with sample partitioning schemes based on proxies for contagion intensity, and consider several alternative explanations for the results. The totality of evidence supports the interpretation that these effects are directly associated with intra-industry bankruptcy contagion.

This study extends the literature that documents a number of cross-firm consequences of intra-industry bankruptcy. The key innovation of my study is the recognition that not only does intra-industry bankruptcy contagion have direct cross-firm consequences for prices (e.g., stocks, loans, credit default swaps), but it also fundamentally affects the industry information processing environment and alters the manner in which the equity market responds to subsequent news.

I utilize insights from the learning-from-default model of default correlation, and the updating of beliefs mechanism in particular, to formally develop the prediction that contagion in the wake of intra-industry bankruptcy can affect subsequent equity market information processing. In particular, learning-from-default models posit that an observed bankruptcy can cause investors to update their beliefs about unobservable variables that affect other firms' assessed default probabilities, which can affect the response to subsequent earnings news. However, because the learning-from-default models assume Bayesian updating processes, the response asymmetry that I document is difficult to explain within that framework. Therefore, my results suggest that the contagion-related belief updating mechanism is more complicated than that assumed by current models, which can inform the default risk correlation literature.

This study also contributes to the literature that examines how uncertainty in the firm- or market-level information environments affects the stock price response to firm-level earnings news (e.g., Skinner and Sloan, 2002; Conrad et al., 2002; Williams, 2009), by providing evidence that industry-level bankruptcy contagion is an independent driver of response asymmetry. More generally, extant earnings response studies have effectively ignored the possibility that the market response to firm-specific earnings news can be affected by intra-industry economic events. The results of this study suggest that this is a non-trivial omission.

Appendix A – Model uncertainty and information processing

Economic theories of model uncertainty include the theory of ambiguity aversion (e.g., Epstein and Schneider, 2008) and robust control theory (e.g., Hansen and Sargent, 2001), both of which use axioms of max-min utility (Gilboa and Schmeidler, 1989) as their decision theoretic underpinning. Max-min utility is a generalization of the standard or subjective expected utility (SEU) model of decision making under uncertainty (e.g., von Neumann-Morgenstern, 1947; Savage, 1954) that is motivated by evidence (e.g., Ellsberg, 1961) that investors do not always behave in accordance with these standard models. The key difference between the SEU and max-min-based models is that the SEU axioms assert that decision makers maximize expected utility with respect to a unique prior belief over relevant outcomes, whereas the max-min axioms assume that decision makers face a set of possible prior beliefs that cannot be reduced to unity.[21]

Epstein and Schneider (2008) develops a model describing how ambiguity-averse investors process ambiguous signals, where the ambiguity-averse investors behave in accordance with the Gilboa and Schmeidler (1989) axioms. Accordingly, ambiguity-averse investors act as if they solve for the decision rule that maximizes expected utility under each possible probability distribution within their range of signal likelihoods, and then apply the decision rule under which the minimum optimized expected utility is obtained. Because investors evaluate any action using the conditional probability that minimizes the utility of that action, they respond asymmetrically to ambiguous information.

[21] Camerer and Weber (1992) and Kelsey and Quiggin (1992) provide surveys of the literature that uses max-min utility theory, and its distinction from classical subjective expected utility and Bayesian decision theory.

As an example, suppose that an investor wants to learn about a certain fundamental parameter θ (e.g., future economic earnings), where an ambiguous signal s (e.g., earnings announcement released in the wake of intra-industry bankruptcy) is related to θ as follows:

$$s = \theta + \varepsilon, \varepsilon \sim N(0, \sigma_s^2), \sigma_s^2 \in \left[\underline{\sigma}_s^2, \overline{\sigma}_s^2 \right].$$

Also, assume that the investor has a unique prior distribution over θ, where $\theta \sim N(m, \sigma_\theta^2)$. When the ambiguity-averse investor sees the ambiguous signal, she updates her prior as follows:

$$\theta \sim N\left(m + \frac{\sigma_\theta^2}{\sigma_\theta^2 + \sigma_s^2}(s - m), \frac{\sigma_s^2 \sigma_\theta^2}{\sigma_\theta^2 + \sigma_s^2} \right), \sigma_s^2 \in \left[\underline{\sigma}_s^2, \overline{\sigma}_s^2 \right].$$

In particular, if the investor sees bad news ($s < m$), then $\sigma_s^2 = \underline{\sigma}_s^2$, i.e., she views the signal as precise, and if the investor sees good news ($s > m$), then $\sigma_s^2 = \overline{\sigma}_s^2$, i.e., she acts as if the signal is imprecise.[22] In the earnings announcement context, ($s - m$) represents unexpected earnings, and $\frac{\sigma_\theta^2}{\sigma_\theta^2 + \sigma_s^2}$ represents the response coefficient. It follows directly that ambiguity-averse investors will drive larger stock price responses for bad news relative to good news, i.e., negative asymmetry, for earnings announcements that are viewed as ambiguous.

The above example suggests that ambiguity takes the form of "ambiguity in signal quality." However, the Epstein and Schneider (2008) model can be used to generate identical predictions if the ambiguity relates to volatility in fundamentals, i.e., "ambiguity in fundamentals". That is, assume that the investor does not possess a unique prior distribution

[22] In the standard Bayesian case, σ_s^2 takes on a unique value.

over the fundamental parameter θ, but rather $\theta \sim N(m, \sigma_\theta^2), \sigma_\theta^2 \in \left[\underline{\sigma}_\theta^2, \overline{\sigma}_\theta^2\right]$. Holding signal quality constant, if the investor then sees bad news (*s* < *m*), then $\sigma_\theta^2 = \overline{\sigma}_\theta^2$, i.e., she acts as if the signal is very informative about fundamentals, and if the investor sees good news (*s* > *m*), then $\sigma_\theta^2 = \underline{\sigma}_\theta^2$, i.e., she acts as if the signal is not very informative about fundamentals.

Appendix B – Interaction specification

In this appendix, I outline the model structure that underlies the three-way interaction specification used in Eq. (2). At the most elementary level, the basic process behind my study is that abnormal return is a function of earnings news:

(A1) $$CAR = \left(\alpha_0 + \alpha_1 NEWS\right) + \sum \alpha_k CONTROL,$$

where *CONTROL*s affect *CAR*, but are not modeled as having interactive effects with *NEWS*. I first conjecture that the sign of news affects this process, creating the secondary process

(A2) $$CAR = \left(\alpha_0 + \alpha_1 NEWS\right)\left(\alpha_2 + \alpha_3 BAD\right) + \sum \alpha_k CONTROL.$$

Algebraic expansion of (A2) leads to the following interactive model structure:

$$CAR = \theta_0 + \theta_1 BAD + \theta_2 NEWS + \theta_3 BAD * NEWS + \sum \alpha_k CONTROL.$$

Next, the premise of this paper is that bankruptcy affects (A2), creating the tertiary process

(A3) $$\begin{aligned} CAR &= \left(\alpha_0 + \alpha_1 NEWS\right)\left(\alpha_2 + \alpha_3 BAD\right)\left(\alpha_4 + \alpha_5 BW\right) + \sum \alpha_k CONTROL \\ &= \left[\theta_0 + \theta_1 BAD + \theta_2 NEWS + \theta_3 BAD * NEWS\right]\left(\alpha_4 + \alpha_5 BW\right) + \sum \alpha_k CONTROL. \end{aligned}$$

Further algebraic expansion of (A3) leads to the following triple interactive structure:

$$\begin{aligned} CAR = {} & \delta_0 + \delta_1 BAD + \delta_2 NEWS + \delta_3 BAD * NEWS \\ & + \delta_4 BW + \delta_5 BW * BAD + \delta_6 BW * NEWS + \delta_7 BW * BAD * NEWS \\ & + \sum \alpha_k CONTROL. \end{aligned}$$

Controlling for the level of *VIX* as a competing explanation against *BW* alters (A3) as follows:

(A4) $$\begin{aligned} CAR &= \left(\alpha_0 + \alpha_1 NEWS\right)\left(\alpha_2 + \alpha_3 BAD\right)\left(\alpha_4 + \alpha_5 BW + \alpha_6 HVIX\right) + \sum \alpha_k CONTROL \\ &= \left[\theta_0 + \theta_1 BAD + \theta_2 NEWS + \theta_3 BAD * NEWS\right]\left(\alpha_4 + \alpha_5 BW + \alpha_6 HVIX\right) + \sum \alpha_k CONTROL, \end{aligned}$$

where algebraic expansion leads to the following specification, which I estimate using Eq. (2):

$$
\begin{aligned}
CAR = {} & \beta_0 + \beta_1 BAD + \beta_2 NEWS + \beta_3 BAD * NEWS \\
& + \beta_4 BW + \beta_5 BW * BAD + \beta_6 BW * NEWS + \beta_7 BW * BAD * NEWS \\
& + \beta_8 HVIX + \beta_9 HVIX * BAD + \beta_{10} HVIX * NEWS + \beta_{11} HVIX * BAD * NEWS \\
& + \sum \alpha_k CONTROL.
\end{aligned}
$$

Finally, in certain untabulated tests within the paper I use a four-way interaction specification to test the significance of coefficient differences across sample partitions. In other words, I model the possibility that a given partitioning variable affects the process in (A4), as follows:

$$
\begin{aligned}
CAR &= \left[(\alpha_0 + \alpha_1 NEWS)(\alpha_2 + \alpha_3 BAD)(\alpha_4 + \alpha_5 BW + \alpha_6 HVIX) + \sum \alpha_k CONTROL \right] (\alpha_{k+1} + \alpha_{k+2} PART) \\
&= \begin{bmatrix} \beta_0 + \beta_1 BAD + \beta_2 NEWS + \beta_3 BAD * NEWS \\ + \beta_4 BW + \beta_5 BW * BAD + \beta_6 BW * NEWS + \beta_7 BW * BAD * NEWS \\ + \beta_8 HVIX + \beta_9 HVIX * BAD + \beta_{10} HVIX * NEWS + \beta_{11} HVIX * BAD * NEWS \\ + \sum \alpha_k CONTROL \end{bmatrix} (\alpha_{k+1} + \alpha_{k+2} PART),
\end{aligned} \tag{A5}
$$

where *PART* is an indicator for a given sample partition. Algebraic expansion leads to the following four-way-interactive specification:

$$
\begin{aligned}
CAR = {} & \gamma_0 + \gamma_1 BAD + \gamma_2 NEWS + \gamma_3 BAD * NEWS \\
& + \gamma_4 BW + \gamma_5 BW * BAD + \gamma_6 BW * NEWS + \gamma_7 BW * BAD * NEWS \\
& + \gamma_8 HVIX + \gamma_9 HVIX * BAD + \gamma_{10} HVIX * NEWS + \gamma_{11} HVIX * BAD * NEWS \\
& + \gamma_{12} PART + \gamma_{13} PART * BAD + \gamma_{14} PART * NEWS + \gamma_{15} PART * BAD * NEWS \\
& + \gamma_{16} PART * BW + \gamma_{17} PART * BW * BAD + \gamma_{18} PART * BW * NEWS + \gamma_{19} PART * BW * BAD * NEWS \\
& + \gamma_{20} PART * HVIX + \gamma_{21} PART * HVIX * BAD + \gamma_{22} PART * HVIX * NEWS \\
& + \gamma_{23} PART * HVIX * BAD * NEWS + \sum \alpha_k CONTROL + PART * \sum \gamma_k CONTROL.
\end{aligned}
$$

For example, a test of significance on γ_{19} provides a test of whether bankruptcy-induced negative asymmetry is different across the tested sample partitions.

Appendix C – Monte Carlo randomization test methodology

It is straightforward to use Monte Carlo randomization to test whether a coefficient estimate is statistically greater for sample partition B than sample partition A. To illustrate, I use the example of testing the difference in the coefficient estimate for *BW*BAD*NEWS* across sample partitions as in Table 5, although the methodology is generalizable to any cross-partition test statistic. In step one, I compute the relevant test statistic (*DIFF*) as the estimated coefficient in partition B (e.g., *NBANK* > 1) minus the estimated coefficient in partition A (e.g., *NBANK* = 1), where each coefficient is obtained from estimation of Eq. (2) on the actual sample partitions. In step two, from the original $N_{BW=1}$ (44,987) observations in my data set I randomly assign n_1 (the original number of *NBANK* > 1 observations, i.e., 8,233) observations without replacement to a random *NBANK** > 1 partition, leaving the $N_{BW=1}$ sample complement of n_2 (36,754) randomly assigned observations as the corresponding *NBANK** = 1 partition, and combine each randomly permuted *NBANK** partition with the actual 229,671 *BW*=0 observations (which remain constant across partitions). In Step Three, I estimate equation (2) for each of the randomly permuted sample partitions, and compute the difference in the *BW*BAD*NEWS* coefficient estimates across the permuted partitions (*DIFF**). I then repeat Steps Two and Three 1,000 times. The one-tailed p-value of the test of whether *DIFF* is greater than zero, i.e., whether the coefficient estimate on *BW*BAD*NEWS* in partition B is greater than that in partition A, is calculated as the proportion of the 1,000 sampled permutations where *DIFF** > *DIFF*.[23]

[23] For more details concerning the theory and technical details behind randomization testing, see Edgington and Onghena (2007).

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Table 1: Descriptive statistics

Panel A - Non-Bankruptcy Wake Earnings Announcements ($BW_{i,t}$ = 0)

Variables	N	Mean	Std	Min	P25	Median	P75	Max
$CAR_{i,t}$	229,671	0.002	0.066	-0.307	-0.031	0.000	0.032	0.346
$NEWS_{i,t}$	229,671	-0.002	0.040	-0.566	-0.008	0.001	0.007	0.499
$SIZE_{i,t}$	229,671	5.499	1.878	1.009	4.085	5.396	6.825	11.308
$MTB_{i,t}$	229,671	2.589	2.524	0.245	1.246	1.836	2.951	39.751
$LEV_{i,t}$	229,671	2.625	3.767	0.033	0.523	1.196	2.522	28.297
$PRECAR_{i,t}$	229,671	0.004	0.188	-0.664	-0.102	-0.010	0.090	2.102
VIX_{t}	229,671	20.057	7.893	9.490	14.180	18.760	23.980	113.330

Panel B - Bankruptcy Wake Earnings Announcements ($BW_{i,t}$ = 1)

Variables	N	Mean		Std	Min	P25	Median	P75	Max
$CAR_{i,t}$	44,987	0.002		0.074	-0.306	-0.035	0.000	0.037	0.346
$NEWS_{i,t}$	44,987	-0.003	***	0.052	-0.567	-0.010	0.001	0.007	0.500
$SIZE_{i,t}$	44,987	5.409	***	1.883	1.026	3.972	5.282	6.708	11.281
$MTB_{i,t}$	44,987	2.578		2.747	0.245	1.153	1.766	2.930	39.291
$LEV_{i,t}$	44,987	2.586	**	3.854	0.034	0.482	1.108	2.424	27.775
$PRECAR_{i,t}$	44,987	0.011	***	0.219	-0.666	-0.114	-0.006	0.110	2.059
VIX_{t}	44,987	22.817	***	7.851	9.490	17.000	22.540	26.750	113.330

Panel C - Pearson correlation matrix - *BW* =0 (*BW* =1) above (below) the diagonal

	$CAR_{i,t}$	$NEWS_{i,t}$	$SIZE_{i,t}$	$MTB_{i,t}$	$LEV_{i,t}$	$PRECAR_{i,t}$	VIX_{t}
$CAR_{i,t}$		***0.117***	***0.013***	***-0.013***	***0.006***	***-0.037***	0.002
$NEWS_{i,t}$	***0.094***		***0.042***	***0.043***	***-0.005***	***0.109***	***-0.014***
$SIZE_{i,t}$	***0.014***	***0.046***		***0.221***	***0.060***	***0.049***	***-0.069***
$MTB_{i,t}$	***-0.017***	***0.045***	***0.257***		***-0.054***	***0.098***	***-0.029***
$LEV_{i,t}$	***0.012***	-0.009	***0.042***	***-0.043***		***0.015***	***0.009***
$PRECAR_{i,t}$	***-0.026***	***0.113***	***0.057***	***0.097***	***0.010***		***-0.010***
VIX_{t}	0.007	-0.009	***-0.027***	***-0.039***	***0.112***	***-0.013***	

Panels A and B of Table 1 present descriptive statistics for key variables used in the study. Panel C presents a Pearson correlation matrix, where correlations for the *BW*=0 (*BW*=1) partition are presented above (below) the diagonal, and correlations significant at the 0.05 level or better are presented in bold italics. *BW* is an indicator variable = 1 if there was at least one intra-industry bankruptcy in the thirty calendar days immediately preceding firm *i*'s quarter *t* earnings announcement and = 0 otherwise. Tests of differences in means across *BW* partitions are indicated in Panel B, where *, **, and *** indicate a significance at the 0.10, 0.05, and 0.01 level (two-tailed), respectively. *CAR* is firm *i*'s three-day cumulative abnormal return centered on its quarter *t* earnings announcement. *NEWS* is firm *i*'s seasonally differenced quarterly earnings per share scaled by stock price two trading days prior to quarter *t* earnings announcement. *SIZE* is the natural log of firm *i*'s quarter *t* market value of common equity. *MTB* is firm *i*'s market-to-book ratio at the end of quarter *t*. *LEV* is firm *i*'s total liabilities divided by book equity at the end of quarter *t*. *PRECAR* is firm *i*'s cumulative abnormal return beginning two trading days after its quarter *t*-1 earnings announcement and ending two trading days prior to its quarter *t* earnings announcement. *VIX* is the level of the VIX index two trading days prior to firm *i*'s quarter *t* earnings announcement.

Table 2: Market response to earnings news in the wake of intra-industry bankruptcy

Dependent Variable: $CAR_{i,t}$

	(1)		(2)		(3)	
Parameter	**Coeff**	**T-Stat**	**Coeff**	**T-Stat**	**Coeff**	**T-Stat**
Intercept	-0.0033 ***	(- 3.63)	-0.0033 ***	(- 3.58)	-0.0032 ***	(- 3.51)
NEWS	0.1861 ***	(42.23)	0.2011 ***	(39.96)	0.2421 ***	(35.68)
BW			-0.0002	(- 0.54)	-0.0002	(- 0.49)
BW*NEWS			**-0.0603 *****	**(- 6.18)**	**-0.0476 *****	**(- 4.80)**
HVIX					0.0011 **	(2.23)
*HVIX*NEWS*					-0.0650 ***	(- 7.71)
SIZE	0.0005 ***	(6.38)	0.0005 ***	(6.36)	0.0005 ***	(6.33)
LEV	0.0001 ***	(4.07)	0.0001 ***	(4.06)	0.0001 ***	(4.08)
PRECAR	-0.0175 ***	(- 19.42)	-0.0175 ***	(- 19.44)	-0.0176 ***	(- 19.54)
Fixed Effects	Year		Year		Year	
Clustered SE	Firm		Firm		Firm	
$N_{BW=0}$	229,671		229,671		229,671	
$N_{BW=1}$	44,987		44,987		44,987	
R^2	0.016		0.016		0.017	

Table 2 presents results from OLS estimation of the following model using firm-quarter earnings announcements over the period 1986-2007:

$$CAR_{i,t} = \alpha_0 + \alpha_1 BW_{i,t} + \alpha_2 NEWS_{i,t} + \alpha_3 BW * NEWS_{i,t} + \sum \alpha_k CONTROL_{i,t} + \gamma_{i,t} \qquad (1)$$

Year fixed effects are based on calendar year of earnings announcement. *BW* is an indicator variable that = 1 if there was at least one intra-industry bankruptcy in the thirty calendar days immediately preceding firm *i*'s quarter *t* earnings announcement and = 0 otherwise. *CAR* is firm *i*'s three-day cumulative abnormal return centered on its quarter *t* earnings announcement date. *NEWS* is firm *i*'s seasonally differenced quarterly earnings per share scaled by stock price two trading days prior to quarter *t* earnings announcement. *HVIX* is an indicator variable that =1 if *VIX* is above the sample median value and = 0 otherwise, where *VIX* is the level of the VIX index two trading days prior to firm *i*'s quarter *t* earnings announcement. *SIZE* is the natural log of firm *i*'s quarter *t* market value of common equity. *LEV* is firm *i*'s total liabilities divided by book equity at the end of quarter *t*. *PRECAR* is firm *i*'s cumulative abnormal return beginning two trading days after its quarter *t*-1 earnings announcement and ending two trading days prior to quarter *t* earnings announcement. *, **, and *** denote two-tailed statistical significance at the 0.10, 0.05, and 0.01 levels, respectively.

Table 3: Asymmetric response to news in the wake of intra-industry bankruptcy

Dependent Variable: $CAR_{i,t}$

	(1)		(2)		(3)	
Parameter	**Coeff**	**T-Stat**	**Coeff**	**T-Stat**	**Coeff**	**T-Stat**
Intercept	0.0056 ***	(5.73)	0.0052 ***	(5.31)	0.0053 ***	(5.37)
BAD	-0.0175 ***	(-54.32)	-0.0172 ***	(-49.43)	-0.0174 ***	(-41.19)
NEWS	0.0730 ***	(9.60)	0.0977 ***	(11.57)	0.1149 ***	(10.02)
*NEWS*BAD*	0.0279 ***	(2.84)	0.0046	(0.42)	0.0000	(0.00)
BW			0.0016 ***	(2.98)	0.0017 ***	(2.99)
*BW*BAD*			-0.0016 *	(-1.92)	-0.0017 **	(-2.02)
*BW*NEWS*			-0.0960 ***	(-5.66)	-0.0902 ***	(-5.23)
BW*BAD*NEWS			**0.0901 *** **	**(4.13)**	**0.0874 *** **	**(3.93)**
HVIX					0.0010 *	(1.78)
*HVIX*BAD*					0.0008	(1.28)
*HVIX*NEWS*					-0.0269 *	(-1.85)
*HVIX*BAD*NEWS*					0.0090	(0.48)
SIZE	0.0003 ***	(4.36)	0.0003 ***	(4.51)	0.0003 ***	(4.46)
LEV	0.0001 ***	(3.09)	0.0001 ***	(2.98)	0.0001 ***	(3.00)
BADPRE	0.0005	(1.24)	0.0005	(1.26)	0.0005	(1.33)
PRECAR	-0.0179 ***	(-11.74)	-0.0180 ***	(-11.78)	-0.0180 ***	(-11.79)
*BADPRE*PRECAR*	-0.0053 *	(-1.92)	-0.0052 *	(-1.89)	-0.0051 *	(-1.84)
Fixed Effects	Year		Year		Year	
Clustered SE	Firm		Firm		Firm	
$N_{BW=0}$	229,671		229,671		229,671	
$N_{BW=1}$	44,987		44,987		44,987	
% BAD $_{BW=0}$	0.44		0.44		0.44	
% BAD $_{BW=1}$	0.46		0.46		0.46	
R^2	0.029		0.029		0.029	
F-Tests	**Coeff**	**P-Value**	**Coeff**	**P-Value**	**Coeff**	**P-Value**
$\beta_6 + \beta_7 = 0$			-0.0059	0.6607	-0.0028	0.8388

Table 3 presents results from OLS estimation of the following model using firm-quarter earnings announcements over the period 1986-2007:

$$CAR_{i,t} = \beta_0 + \beta_1 BAD_{i,t} + \beta_2 NEWS_{i,t} + \beta_3 BAD*NEWS_{i,t} + \beta_4 BW_{i,t} + \beta_5 BW*BAD_{i,t}$$
$$+\beta_6 BW*NEWS_{i,t} + \beta_7 BW*BAD*NEWS_{i,t} + \sum \beta_k CONTROL_{i,t} + \varepsilon_{i,t}.$$

Year fixed effects are based on calendar year of earnings announcement. *CAR* is firm *i*'s three-day cumulative abnormal return centered on its quarter *t* earnings announcement date. *BAD* is an indicator variable that =1 if *NEWS* < 0 and = 0 otherwise, where *NEWS* is firm *i*'s seasonally differenced quarterly earnings per share scaled by stock price two trading days prior to its quarter *t* earnings announcement. *BW* is an indicator variable that = 1 if there was at least one intra-industry bankruptcy in the thirty calendar days immediately preceding firm *i*'s quarter *t* earnings announcement and = 0 otherwise. *HVIX* is an indicator variable that =1 if *VIX* is above the sample median value and = 0 otherwise, where *VIX* is the level of the VIX index two trading days prior to firm *i*'s quarter *t* earnings announcement. *SIZE* is the natural log of firm *i*'s quarter *t* market value of common equity. *LEV* is firm *i*'s total liabilities divided by book equity at the end of quarter *t*. *BADPRE* is an indicator variable that = 1 if *PRECAR* < 0 and = 0 otherwise, where *PRECAR* is firm *i*'s cumulative abnormal return beginning two trading days after its quarter *t*-1 earnings announcement and ending two trading days prior to its quarter *t* earnings announcement. *, **, and *** denote two-tailed statistical significance at the 0.10, 0.05, and 0.01 levels, respectively.

Table 4: Preemptive disclosure of good news

Dependent Variable: $CAR_{i,t}$

	(1)		(2)		(2)-(1)	
	***DOBEA_CAR* Q1**		***DOBEA_CAR* Q4**		**Diff. in**	**One-Tail**
Parameter	**Coeff**	**T-Stat**	**Coeff**	**T-Stat**	**Coeff.**	**P-Value**
Intercept	0.0060 ***	(5.93)	0.0054 ***	(5.34)	-0.0006 ##	0.015
BAD	-0.0175 ***	(- 40.43)	-0.0172 ***	(- 40.18)	0.0002 ###	0.004
NEWS	0.1056 ***	(9.03)	0.1096 ***	(9.22)	0.0039 #	0.095
*NEWS*BAD*	0.0101	(0.67)	0.0107	(0.70)	0.0006	0.465
BW	0.0084 ***	(6.81)	-0.0017	(- 1.62)	-0.0102 ###	0.000
*BW*BAD*	-0.0027	(- 1.53)	-0.0053 ***	(- 3.05)	-0.0026	0.123
*BW*NEWS*	-0.0790 ***	(- 2.57)	-0.1493 ***	(- 5.58)	-0.0703 ##	0.033
BW*BAD*NEWS	**0.0759 ****	**(2.04)**	**0.1729 *****	**(4.69)**	**0.0970 ##**	**0.029**
*CONTROL*s included	Yes		Yes			
Fixed Effects	Year		Year			
Clustered SE	Firm		Firm			
$N_{BW=0}$	229,671		229,671			
$N_{BW=1}$	11,247		11,246			
$\% BAD_{BW=0}$	0.44		0.44			
$\% BAD_{BW=1}$	0.51		0.43			
R^2	0.031		0.030			
F-Tests	**Coeff**	**P-Value**	**Coeff**	**P-Value**		
$\beta_6 + \beta_7 = 0$	-0.0031	0.8852	0.0236	0.3585		

Table 4 presents results from OLS estimation of the following model using firm-quarter earnings announcements over the period 1986-2007, for the upper and lower quartiles of *BW*=1 earnings announcements based on *DOBEA_CAR*:

$$CAR_{i,t} = \beta_0 + \beta_1 BAD_{i,t} + \beta_2 NEWS_{i,t} + \beta_3 BAD*NEWS_{i,t} + \beta_4 BW_{i,t} + \beta_5 BW*BAD_{i,t}$$
$$+\beta_6 BW*NEWS_{i,t} + \beta_7 BW*BAD*NEWS_{i,t} + \sum \beta_k CONTROL_{i,t} + \varepsilon_{i,t}.$$

DOBEA_CAR is firm *i*'s cumulative abnormal return between the three-day bankruptcy announcement window and the three-day quarter *t* earnings announcement window. Year fixed effects are based on earnings announcement dates. *CAR* is firm *i*'s three-day cumulative abnormal return centered on its quarter *t* earnings announcement date. *BAD* is an indicator variable that =1 if *NEWS* < 0 and = 0 otherwise, where *NEWS* is firm *i*'s seasonally differenced quarterly earnings per share scaled by stock price two trading days prior to quarter *t* earnings announcement. *BW* is an indicator variable that = 1 if there was at least one intra-industry bankruptcy in the thirty calendar days immediately preceding firm *i*'s quarter *t* earnings announcement and = 0 otherwise. *CONTROL*s are *HVIX*, *HVIX*BAD*, *HVIX*NEWS*, *HVIX*BAD*NEWS*, *SIZE*, *LEV*, *BADPRE*, *PRECAR*, and *BADPRE*PRECAR*. *HVIX* is an indicator variable that =1 if *VIX* is above the sample median value and = 0 otherwise, where *VIX* is the level of the VIX index two trading days prior to firm *i*'s quarter *t* earnings announcement. *SIZE* is the natural log of firm *i*'s quarter *t* market value of common equity. *LEV* is firm *i*'s total liabilities divided by book equity at the end of quarter *t*. *BADPRE* is an indicator variable that = 1 if *PRECAR* < 0 and = 0 otherwise, where *PRECAR* is firm *i*'s cumulative abnormal return beginning two trading days after its quarter *t*-1 earnings announcement and ending two trading days prior to quarter *t* earnings announcement. *, **, and *** (#, ##, ###) denote two-tailed (one-tailed) statistical significance at the 0.10, 0.05, and 0.01 levels, respectively.

Table 5: Relative industry health

Dependent Variable: $CAR_{i,t}$

	(1)		(2)		(2)-(1)	
	$NBANK_{i,t} = 1$		$NBANK_{i,t} > 1$		Diff. in	One-Tail
Parameter	**Coeff**	**T-Stat**	**Coeff**	**T-Stat**	**Coeff.**	**P-Value**
Intercept	0.0053 ***	(5.35)	0.0059 ***	(5.84)	0.0007	0.139
BAD	-0.0176 ***	(- 41.41)	-0.0172 ***	(- 40.09)	0.0003 ###	0.001
NEWS	0.1094 ***	(9.45)	0.1097 ***	(9.31)	0.0004 ##	0.029
*NEWS*BAD*	0.0039	(0.26)	0.0119	(0.78)	0.0080	0.584
BW	0.0010	(1.69)	0.0041 ***	(2.87)	0.0031 ##	0.011
*BW*BAD*	-0.0009	(- 1.04)	-0.0046 **	(- 2.23)	-0.0037 ##	0.034
*BW*NEWS*	-0.0583 ***	(- 2.96)	-0.1635 ***	(- 5.61)	-0.1052 ###	0.001
BW*BAD*NEWS	**0.0538 ****	**(2.12)**	**0.1665 *****	**(4.47)**	**0.1128 ##**	**0.011**
*CONTROL*s included	Yes		Yes			
Fixed Effects	Year		Year			
Clustered SE	Firm		Firm			
$N_{BW=0}$	229,671		229,671			
$N_{BW=1}$	36,754		8,233			
$\% BAD_{BW=0}$	0.44		0.44			
$\% BAD_{BW=1}$	0.46		0.48			
R^2	0.030		0.030			
F-Tests	**Coeff**	**P-Value**	**Coeff**	**P-Value**		
$\beta_6 + \beta_7 = 0$	-0.0046	0.7719	0.0030	0.8964		

Table 5 presents results from OLS estimation of the following model using firm-quarter earnings announcements over the period 1986-2007 for two separate partitions of *BW*=1 earnings announcements based on *NBANK*:

$$CAR_{i,t} = \beta_0 + \beta_1 BAD_{i,t} + \beta_2 NEWS_{i,t} + \beta_3 BAD*NEWS_{i,t} + \beta_4 BW_{i,t} + \beta_5 BW*BAD_{i,t}$$
$$+\beta_6 BW*NEWS_{i,t} + \beta_7 BW*BAD*NEWS_{i,t} + \sum \beta_k CONTROL_{i,t} + \varepsilon_{i,t}.$$

NBANK is the number of intra-industry bankruptcies in the thirty-day calendar window preceding the quarter t earnings announcement. Year fixed effects are based on calendar year of earnings announcement. *CAR* is firm i's three-day cumulative abnormal return centered on its quarter t earnings announcement date. *BAD* is an indicator variable that =1 if *NEWS* < 0 and = 0 otherwise, where *NEWS* is firm i's seasonally differenced quarterly earnings per share scaled by stock price two trading days prior to quarter t earnings announcement. *BW* is an indicator variable that = 1 if there was at least one intra-industry bankruptcy in the thirty calendar days immediately preceding firm i's quarter t earnings announcement and = 0 otherwise. *CONTROL*s are *HVIX*, *HVIX*BAD*, *HVIX*NEWS*, *HVIX*BAD*NEWS*, *SIZE*, *LEV*, *BADPRE*, *PRECAR*, and *BADPRE*PRECAR*. *HVIX* is an indicator variable that =1 if *VIX* is above the sample median value and = 0 otherwise, where *VIX* is the level of the VIX index two trading days prior to firm i's quarter t earnings announcement. *SIZE* is the natural log of firm i's quarter t market value of common equity. *LEV* is firm i's total liabilities divided by book equity at the end of quarter t. *BADPRE* is an indicator variable that = 1 if *PRECAR* < 0 and = 0 otherwise, where *PRECAR* is firm i's cumulative abnormal return beginning two trading days after its quarter t-1 earnings announcement and ending two trading days prior to its quarter t earnings announcement. *, **, and *** ($^{\#}$, $^{\#\#}$, $^{\#\#\#}$) denote two-tailed (one-tailed) statistical significance at the 0.10, 0.05, and 0.01 levels, respectively.

Table 6: Firm economic comparability with industry peers

Dependent Variable: $CAR_{i,t}$

	(1)		(2)		(2)-(1)	
	Adj. R^2 < Median		Adj. R^2 > Median		Diff. in	One-Tail
Parameter	**Coeff**	**T-Stat**	**Coeff**	**T-Stat**	**Coeff.**	**P-Value**
Intercept	0.0068 ***	(4.61)	0.0053 ***	(3.90)	-0.0014	0.204
BAD	-0.0211 ***	(- 33.73)	-0.0141 ***	(- 25.67)	0.0070 ###	0.001
NEWS	0.1266 ***	(8.04)	0.0924 ***	(5.44)	-0.0342 #	0.061
*NEWS*BAD*	-0.0155	(- 0.76)	0.0215	(1.00)	0.0370 #	0.081
BW	0.0000	(0.03)	0.0031 ***	(4.06)	0.0031 ###	0.006
*BW*BAD*	-0.0017	(- 1.33)	-0.0018	(- 1.53)	-0.0001	0.465
*BW*NEWS*	-0.0627 ***	(- 2.69)	-0.1206 ***	(- 4.56)	-0.0578 ##	0.048
BW*BAD*NEWS	**0.0571 ***	**(1.84)**	**0.1279 *****	**(3.87)**	**0.0708 #**	**0.055**
*CONTROL*s included	Yes		Yes			
Fixed Effects	Year		Year			
Clustered SE	Firm		Firm			
$N_{BW=0}$	111,511		110,535			
$N_{BW=1}$	21,111		22,088			
$\% BAD_{BW=0}$	0.44		0.44			
$\% BAD_{BW=1}$	0.46		0.46			
R^2	0.037		0.022			
F-Tests	**Coeff**	**P-Value**	**Coeff**	**P-Value**		
$\beta_6 + \beta_7 = 0$	-0.0057	0.781	0.0073	0.711		

Table 6 presents results from OLS estimation of the following model using firm-quarter earnings announcements over the period 1986-2007 for two partitions of BW=1 earnings announcements:

$$CAR_{i,t} = \beta_0 + \beta_1 BAD_{i,t} + \beta_2 NEWS_{i,t} + \beta_3 BAD*NEWS_{i,t} + \beta_4 BW_{i,t} + \beta_5 BW*BAD_{i,t}$$
$$+\beta_6 BW*NEWS_{i,t} + \beta_7 BW*BAD*NEWS_{i,t} + \sum \beta_k CONTROL_{i,t} + \varepsilon_{i,t},$$

based on the adj-R^2 from the following firm-specific regression using up to 36 monthly observations ending with the month of firms i's quarter t end date: $RET_{i,m} = \alpha_0 + \alpha_1 VWFFRET_m + v_{i,m}$. *RET* is firm i's raw return in month m. *VWFFRET* is the value-weighted return on firm i's Fama-French 48 industry category in month m. *CAR* is firm i's three-day cumulative abnormal return centered on its quarter t earnings announcement. *BAD* is an indicator variable =1 if $NEWS < 0$ and = 0 otherwise, where *NEWS* is firm i's seasonally differenced quarterly earnings per share scaled by stock price two trading days prior to its quarter t earnings announcement. *BW* is an indicator variable = 1 if there was an intra-industry bankruptcy in the thirty calendar days immediately preceding firm i's quarter t earnings announcement and = 0 otherwise. *CONTROL*s are *HVIX*, *HVIX*BAD*, *HVIX*NEWS*, *HVIX*BAD*NEWS*, *SIZE*, *LEV*, *BADPRE*, *PRECAR*, and *BADPRE*PRECAR*. *HVIX* is an indicator variable that =1 if *VIX* is above the sample median value and = 0 otherwise. *VIX* is the level of the VIX index two trading days prior to firm i's quarter t earnings announcement. *SIZE* is the natural log of firm i's quarter t market value of common equity. *LEV* is firm i's total liabilities divided by book equity at the end of quarter t. *BADPRE* is an indicator variable that = 1 if $PRECAR < 0$ and = 0 otherwise. *PRECAR* is firm i's cumulative abnormal return beginning two trading days after its quarter t-1 earnings announcement and ending two trading days prior to its quarter t earnings announcement. *, **, and *** (#, ##, ###) denote two-tailed (one-tailed) statistical significance at the 0.10, 0.05, and 0.01 levels, respectively.

Table 7: Earnings announcer financial health

Dependent Variable: $CAR_{i,t}$

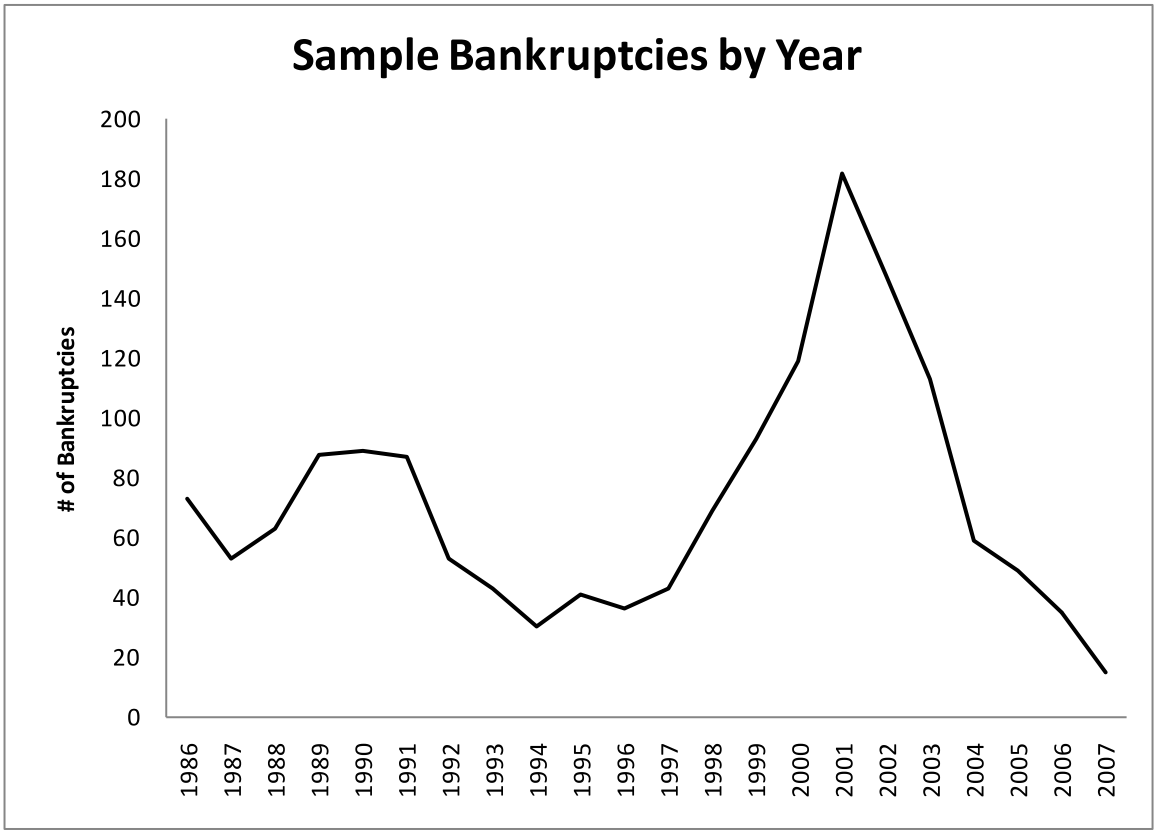
	(1)		(2)		(3)	
	HSCORE = {0, 1}		*HSCORE* = {2}		*HSCORE* = {3}	
Parameter	**Coeff**	**T-Stat**	**Coeff**	**T-Stat**	**Coeff**	**T-Stat**
Intercept	0.0063 ***	(3.09)	0.0045 **	(2.11)	0.0082 ***	(2.94)
BAD	-0.0212 ***	(- 24.94)	-0.0192 ***	(- 22.41)	-0.0213 ***	(- 18.30)
NEWS	0.0695 ***	(3.57)	0.1100 ***	(5.25)	0.1973 ***	(4.58)
*NEWS*BAD*	0.0155	(0.62)	0.0266	(0.95)	-0.0778	(- 1.44)
BW	0.0010	(0.89)	0.0009	(0.75)	0.0025 *	(1.65)
*BW*BAD*	-0.0017	(- 1.01)	-0.0018	(- 1.03)	-0.0010	(- 0.43)
*BW*NEWS*	-0.0720 **	(- 2.54)	-0.1015 ***	(- 3.21)	-0.0792	(- 1.50)
BW*BAD*NEWS	**0.0767 **	**(2.12)**	**0.0937 **	**(2.27)**	**0.0026**	**(0.04)**
*CONTROL*s included	Yes		Yes		Yes	
Fixed Effects	Year		Year		Year	
Clustered SE	Firm		Firm		Firm	
$N_{BW=0}$	62,508		58,060		33,640	
$N_{BW=1}$	13,316		12,500		6,993	
$\% BAD_{BW=0}$	0.46		0.44		0.41	
$\% BAD_{BW=1}$	0.48		0.46		0.44	
R^2	0.030		0.029		0.033	
F-Tests	**Coeff**	**P-Value**	**Coeff**	**P-Value**	**Coeff**	**P-Value**
$\beta_6 + \beta_7 = 0$	0.0047	0.8355	-0.0078	0.774	-0.0767	0.101

Table 7 presents results from OLS estimation of the following model using firm-quarter earnings announcements over the period 1987-2007:

$$CAR_{i,t} = \beta_0 + \beta_1 BAD_{i,t} + \beta_2 NEWS_{i,t} + \beta_3 BAD * NEWS_{i,t} + \beta_4 BW_{i,t} + \beta_5 BW * BAD_{i,t}$$
$$+\beta_6 BW * NEWS_{i,t} + \beta_7 BW * BAD * NEWS_{i,t} + \sum \beta_k CONTROL_{i,t} + \varepsilon_{i,t},$$

where the sample is partitioned based on financial health (*HSCORE*). *HSCORE* = *HCFO* + *HLIQUID* + *HLEVER*, where *HCFO* is an indicator = 1 if cash flow from operations > 0 and = 0 otherwise, and *HLIQUID* (*HLEVER*) is an indicator = 1 if liquidity (leverage) is greater (less) than the industry-fiscal-quarter median. Year fixed effects are based on calendar year of earnings announcement. *CAR* is firm *i*'s three-day cumulative abnormal return centered on its quarter *t* earnings announcement date. *BAD* is an indicator variable that =1 if *NEWS* < 0 and = 0 otherwise, where *NEWS* is firm *i*'s seasonally differenced quarterly earnings per share scaled by stock price two trading days prior to its quarter *t* earnings announcement. *BW* is an indicator variable that = 1 if there was at least one intra-industry bankruptcy in the thirty calendar days immediately preceding firm *i*'s quarter *t* earnings announcement and = 0 otherwise. *CONTROL*s are *HVIX*, *HVIX*BAD*, *HVIX*NEWS*, *HVIX*BAD*NEWS*, *SIZE*, *LEV*, *BADPRE*, *PRECAR*, and *BADPRE*PRECAR*. *HVIX* is an indicator variable that =1 if *VIX* is above the sample median value and = 0 otherwise, where *VIX* is the level of the VIX index two trading days prior to firm *i*'s quarter *t* earnings announcement. *SIZE* is the natural log of firm *i*'s quarter *t* market value of common equity. *LEV* is firm *i*'s total liabilities divided by book equity at the end of quarter *t*. *BADPRE* is an indicator variable that = 1 if *PRECAR* < 0 and = 0 otherwise, where *PRECAR* is firm *i*'s cumulative abnormal return beginning two trading days after its quarter *t*-1 earnings announcement and ending two trading days prior to its quarter *t* earnings announcement. *, **, and *** denote two-tailed statistical significance at the 0.10, 0.05, and 0.01 levels, respectively.

Figure 1: Distribution of bankruptcy filings by year

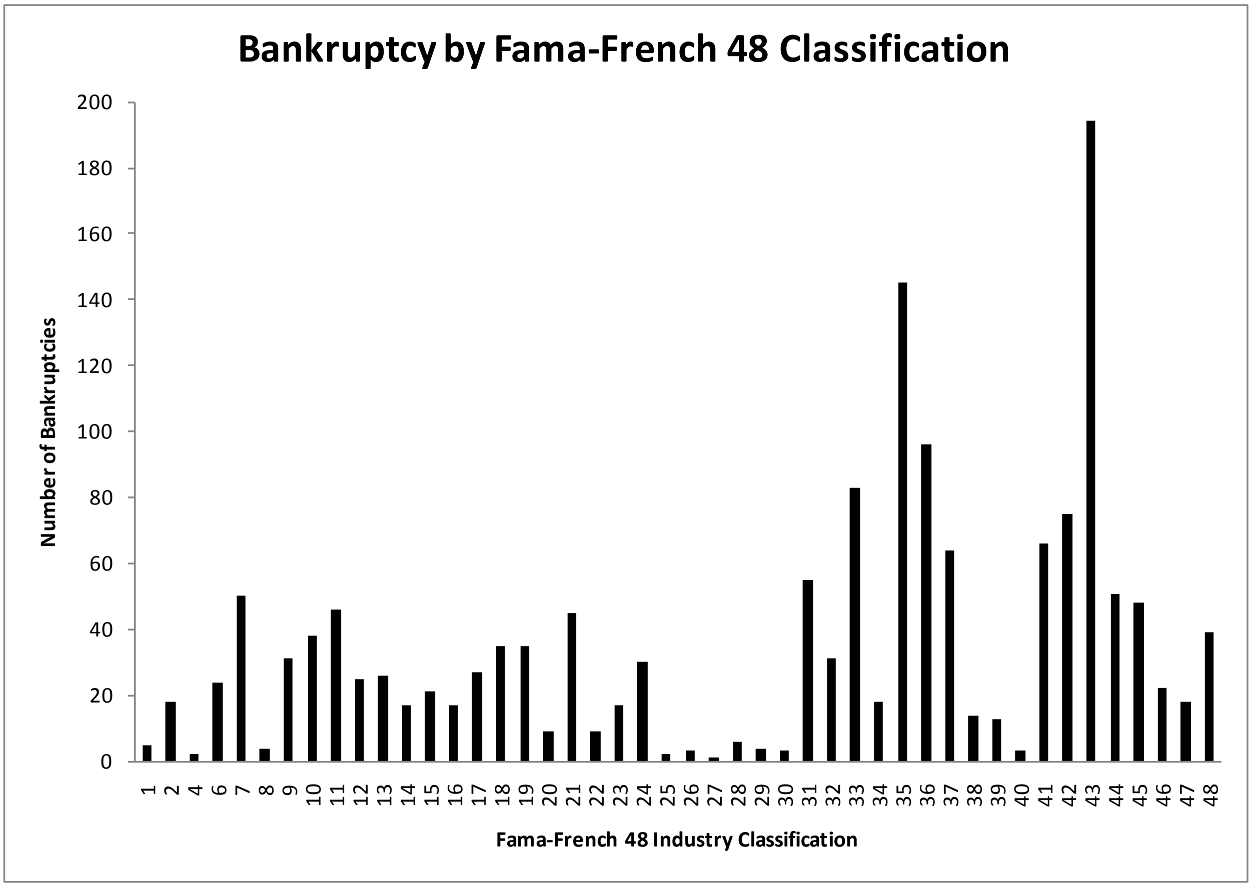


Year	# Bankruptcies	Year	# Bankruptcies	Year	# Bankruptcies
1985*	4	**1993**	43	**2001**	182
1986	73	**1994**	30	**2002**	148
1987	53	**1995**	41	**2003**	113
1988	63	**1996**	36	**2004**	59
1989	88	**1997**	43	**2005**	49
1990	89	**1998**	69	**2006**	35
1991	87	**1999**	93	**2007****	15
1992	53	**2000**	119		

* Bankruptcy filing dates are truncated from below at December 6, 1985.
** Bankruptcy filing dates are truncated from above at June 11, 2007.

Figure 1 presents the calendar-year distribution of the 1,585 Chapter 11 bankruptcy filings underlying the final sample partition of earnings announcements that follow intra-industry bankruptcy. Bankruptcy filing dates range from December 6, 1985 through June 11, 2007. Therefore, data presented here for calendar years 1985 and 2007 are not based on full calendar year periods.

Figure 2: Distribution of bankruptcy filings by Fama-French 48 industry category



Code	Industry Name	Code	Industry Name	Code	Industry Name
1	Agriculture	19	Steel Works, Etc.	34	Personal Services
2	Food Products	20	Fabricated Products	35	Business Services
4	Alcoholic Beverages	21	Machinery	36	Computers
6	Recreational Products	22	Electrical Equipment	37	Electronic Equipment
7	Entertainment	23	Miscellaneous	38	Measuring and Control Equip
8	Printing and Publishing	24	Automobiles and Trucks	39	Business Supplies
9	Consumer Goods	25	Aircraft	40	Shipping Containers
10	Apparel	26	Shipbuilding, Railroad Eq	41	Transportation
11	Healthcare	27	Defense	42	Wholesale
12	Medical Equipment	28	Precious Metals	43	Retail
13	Pharmaceutical Products	29	Nonmetallic Mining	44	Restaurants, Hotel, Motel
14	Chemicals	30	Coal	45	Banking
15	Rubber and Plastic Products	31	Petroleum and Natural Gas	46	Insurance
16	Textiles	32	Utilities	47	Real Estate
17	Construction Materials	33	Telecommunications	48	Trading
18	Construction				

Figure 2 presents the Fama-French 48 industry distribution of the 1,585 Chapter 11 bankruptcy filings underlying the final sample partition of earnings announcements that follow intra-industry bankruptcy.

Figure 3: Baseline asymmetry in the market response to earnings news over time

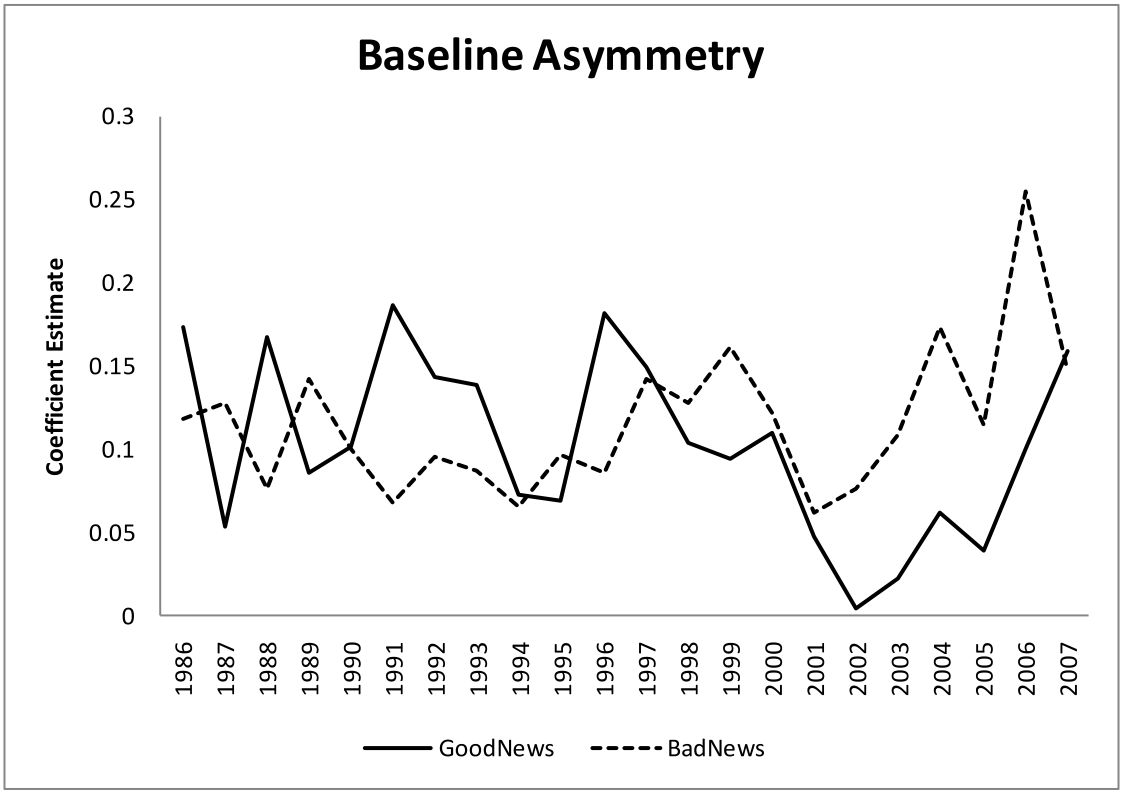


Figure 3 plots the total coefficients on good (β_2) and bad (β_2+β_3) news earnings from the following regression, estimated by calendar year of earnings announcement date, using the full sample of 274,658 quarterly earnings announcements:

$$CAR_{i,t} = \beta_0 + \beta_1 BAD_{i,t} + \beta_2 NEWS_{i,t} + \beta_3 BAD * NEWS_{i,t} + \beta_4 SIZE_{i,t} + \beta_5 LEV_{i,t}$$
$$+\beta_6 BADPRE_{i,t} + \beta_7 PRECAR_{i,t} + \beta_8 BADPRE * PRECAR_{i,t} + \varepsilon_{i,t}.$$

CAR is firm *i*'s three-day cumulative abnormal return centered on its quarter *t* earnings announcement date. *BAD* is an indicator variable that = 1 if *NEWS* < 0 and = 0 otherwise, where *NEWS* is firm *i*'s seasonally differenced quarterly earnings per share scaled by stock price two trading days prior to its quarter *t* earnings announcement date. *SIZE* is the natural log of firm *i*'s market value of common equity at the end of quarter *t*. *LEV* is firm *i*'s total liabilities divided by book equity at the end of quarter *t*. *BADPRE* is an indicator variable that = 1 if *PRECAR* < 0 and = 0 otherwise, where PRECAR is the cumulative abnormal return from two trading days after earnings announcement for quarter *t*-1 through two trading days before earnings announcement for quarter *t*.