Management Forecasts and Litigation Risk*

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Abstract: We examine the influence of the *ex ante* risk of class-action securities litigation on firms' decisions to issue management earnings forecasts, as well as the characteristics of those forecasts. We model litigation risk using a comprehensive sample of 924 class action lawsuits filed by shareholders in federal courts between 1996 and 2002. We find that litigation risk is positively associated with the likelihood of issuing a forecast for all firms, the association is even stronger for firms with bad news. We further examine the effect of litigation risk on the horizon, precision, and the type of news (good or bad) conveyed by the forecast. After controlling for the self-selection bias associated with the decision to issue a forecast, we find that litigation risk is associated with longer forecast horizons, lower precision, and more bad-news forecasts. In addition, we find that after Regulation FD, firms are more likely to issue a forecast, and their forecasts contain relatively more good news, and occur significantly earlier and quarterly forecasts are more precise.

Key Words: Management Forecasts, Litigation Risk, Regulation FD

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1. Introduction

Litigation risk is an important factor associated with firms' voluntary disclosure decisions because high litigation risk firms potentially benefit from issuing management forecasts in two ways (Skinner [1994]). First, forecasting can reduce the probability of being sued by making it more difficult for plaintiffs to claim that the firm withheld material information it was required to disclose. Second, forecasting can reduce the expected amount of damages conditional on being sued by shortening the class period. Understanding the impact of litigation risk on firms' forecast decisions is important to accounting policymakers and securities regulators who are concerned about the potential impact of legal reforms on firms' voluntary forecast decisions. For example, the expansion of Safe Harbor provisions in the Private Securities Litigation Reform Act of 1995 (PSLRA) was designed to encourage more frequent management forecasts (Skinner [1995]).

In this paper, we undertake a systematic analysis of the relation between the *ex ante* risk of litigation and firms' forecasting behavior. In addition to examining the relation between litigation risk and the probability of issuing a forecast, we examine how litigation risk influences the choice of forecast characteristics, including precision, horizon, and the type of news (good or bad) conveyed by the forecast. Our study focuses on the period after the PSLRA since it represents a major revision in the litigation environment and strengthened Safe Harbor provisions designed to protect forward-looking disclosures (Calderon and Kowal [1996], Cochran and McCoy [1996]). Evidence in Johnson et al. [2001] indicates that the relation between litigation risk and forecast behavior changed after the PLSRA, and suggests that a comprehensive re-examination of forecasting behavior is warranted. We extend our analysis by examining how Regulation FD has affected forecast decisions, because Brown et al. [2003] and

Heflin et al. [2003] document substantial changes in voluntary disclosure practices after the regulation was implemented.

Our empirical analyses are based on over 78,000 firm-quarter observations and roughly 24,000 management forecasts issued between 1996 and 2002. We estimate the *ex ante* litigation risk as the fitted probabilities from a logit model of actual federal securities class action litigation as a function of several firm- and time-specific characteristics expected to affect litigation risk. Our final litigation sample includes 924 class action lawsuits filed in federal court in the post-PSLRA period. Unlike most prior studies that have examined the determinants of forecast characteristics, we explicitly control for the self-selection bias associated with the endogenous decision to issue a management forecast, and include forecasts made by firms with no analyst following.

Firms modify their forecast decisions according to the level of litigation risk because of the reputational and monetary costs associated with class action litigation.[1] Between 1996 and 2003, the annual rate of class action litigation among NYSE, AMEX, and NASDAQ firms has varied between 1.2% in 1996 and 2.4% in 2003, averaging 2.25% over this period. In addition to reputation, job security, and opportunity cost considerations for managers, the direct litigation costs borne by firms are substantial, especially for smaller firms. According to Cornerstone Research, the mean (median) post-PSLRA settlement amount in 2003 dollars is $18.6 million ($6.0 million). In addition to any settlement amounts, legal defense costs are typically between one and three million dollars, and range up to $40 million.

We hypothesize that the relation between the likelihood of issuing a management forecast and litigation risk depends crucially on the level of the expectations gap, which we define as the

[1] Lowry and Shu [2002] find evidence that firms adjust their IPO offer prices in response to litigation risk. Their findings indicate that managers are cognizant of their firm's litigation risk and modify their decisions accordingly.

difference between the market's expectations about the next quarter's earnings and the manager's expectation. We expect that litigation risk will be positively associated with issuing a forecast only when the expectations gap is positive (when the firm has bad news); otherwise, the relation will be negative. We also hypothesize that litigation risk will be negatively associated with forecast precision and the news contained in the forecast (i.e., more bad news), and positively associated with forecast horizon.

Our evidence suggests that litigation risk is a significant determinant of firms' choices about whether or not to issue a forecast and, if so, its type. Our tests indicate that high litigation risk firms are significantly more likely to issue forecasts when those forecasts help to reduce the market's overly optimistic expectations relative to firms with either no news or good news. Unlike the findings in Johnson et al. [2001] study of high technology firms, our results suggest that protecting against litigation is an important reason why firms issue bad-news forecasts. Additionally, the evidence generally supports our hypotheses about the relations between litigation risk and forecast characteristics, especially in regards to forecasts of quarterly earnings.

Contrary to our expectations, we do not find any evidence of a negative relation between litigation risk and the likelihood of issuing a forecast for firms with small or negative expectations gaps. Instead, we find that litigation risk is positively associated with forecast activity for firms with either good news or no news. This finding contrasts with those in the prior literature and the general belief of managers, regulators, and federal legislators that higher litigation risk inhibits the disclosure of forward-looking information (Baginski et al. [2002], House of Representatives [1995], Johnson et al. [2001]). For example, Baginski et al. [2002] find that firms in the U.S. issue forecasts less frequently than do firms in less litigious Canada. Nevertheless, both sets of findings are consistent with firms aiming to minimize the expected costs of litigation. The more litigious U.S. environment deters the issuance of forecasts, which could trigger litigation, so the average disclosure level is lower. However, conditional on being in the U.S. legal environment, firms will choose to issue a forecast if they believe it will reduce their litigation risk (taking into account other costs and benefits).[2]

Regulation FD was intended to end the practice of selective disclosure, whereby firms selectively provided material information to securities analysts and institutional investors before the information was publicly available. To the extent that selective disclosure was commonplace before Regulation FD, it is likely that Regulation FD substantially affected firms' disclosure decisions by eliminating the selective disclosure option.[3] In support of this view, Brown et al. [2003] document an almost 100% increase in the frequency of management forecasts after Regulation FD was implemented in October 2000. We expect that numerous firms will release publicly many of the forecasts that they would have chosen to disclose selectively prior to Regulation FD, thus leading to an increase in the likelihood of a firm making a forecast after Regulation FD. In addition, forecast characteristics will change to the extent that these newly-public forecasts differ systematically from the types of forecasts released publicly beforehand. We predict that after Regulation FD, forecasts will convey relatively more good news, and will have longer horizons, and that forecast precision is likely to change, although its direction is unclear. Our empirical results are generally consistent with these predictions.

While much of the prior literature has focused on how forecasts affect subsequent litigation and/or settlement amounts (Francis et al. [1994], Skinner [1997]), our approach

[2] Evidence in Field et al. [2003] indicates that after controlling for the endogeneity between disclosure and litigation, issuing a management forecast reduces litigation risk in the U.S.

[3] Numerous papers, including Bailey et al. [2003] and Heflin et al. [2003], have analyzed the impact of Regulation FD. Gintschel and Markov [2004] and Mohanram and Sunder [2001] find evidence consistent with analysts being the frequent recipients of private information supplied by managers. Also see Gadarowski and Sinha [2003] and Jorion et al. [2004].

involves examining how the *ex ante* risk of litigation affects firms' subsequent forecast decisions.[4] Unlike many previous management forecast studies, our analyses are based on all NYSE, AMEX, and Nasdaq firms, regardless of analyst coverage, and include all types of management forecasts, such as quarterly and annual, point, range, open range, and qualitative. Our results are important because they illustrate when and how firms alter their disclosure practices in response to litigation risk.[5] It is commonly believed that high litigation risk will only cause firms with bad news to issue a public forecast. Instead, we find that litigation risk is positively associated with the likelihood of issuing a forecast for all firms, regardless of the earnings news, although the relation is stronger for firms with bad news.

The paper proceeds as follows. Section 2 discusses the expected relations between litigation risk and forecast decisions, including the expected impact of Regulation FD. Section 3 develops our research methodology and describes the sample. The empirical analyses are discussed in Section 4, while Section 5 concludes the paper. Variable definitions are provided in Appendix A and f the litigation risk model and its estimation are discussed in Appendix B.

2. Litigation Risk and Management Forecasts

2.1 Litigation Risk and the Propensity to Forecast

Prior to Regulation FD, a firm had three primary voluntary disclosure options: (1) it could choose to disclose the information publicly; (2) it could choose to disclose the information privately to select analysts and/or institutional investors; or (3) it could choose not to disclosure the information to any outside parties. In this section, we first discuss how litigation risk is

[4] Much of the prior literature examines time periods when forecast frequency was much lower than during our sample, perhaps due to differences in the legal regime. With the exception of Skinner [1997], these studies examine only a few high technology industries and/or rely on small samples, which may limit the generalizability of their results.

[5] Our approach is similar to that in Frankel et al. [2004], who focus on the relation between litigation risk variables and the decision to implement a "quiet period" policy of non-disclosure (done by less than 1.5% of their sample). In contrast, we examine how litigation risk is associated with the primary aspects of firms' forecast decisions.

expected to influence firms' choices about which disclosure option to take. We then discuss how Regulation FD, which banned the selective disclosure of material information, is expected to affect the likelihood of a firm issuing a public earnings forecast.

In the U.S., as in most other jurisdictions, there is an important asymmetry in the way that damages are calculated in a class-action securities litigation (Skinner [1994]). Investors who incur losses due to stock price declines are eligible to collect damages, while opportunity costs are, as a practical matter, not included in damage calculations. The effect of this asymmetry is that, while firms are often sued after large stock price declines, they are rarely, if ever, sued after large stock price increases. The asymmetric treatment of shareholder damages implies that the impact of litigation risk on forecast decisions will vary depending on the sign and magnitude of the expectations gap between the market and the manager. We partition firms into two groups: (1) where the market's expectations substantially exceed those of the manager (bad news or positive gap); and (2) where either the market's and managers' expectations are roughly equal (no news or small gap) or where the market's expectations are substantially below those of managers (good news or negative gap).

As discussed in Francis et al. [1994] and Skinner [1994], releasing a bad-news forecast publicly when there is a positive expectations gap reduces a firm's expected litigation costs. This reduction accrues through two channels. First, issuing a public forecast reduces the probability of litigation because it provides a defense against claims that management withheld material information that they were legally required to disclose (Field et al. [2003], Francis et al. [1994], Skinner [1994]).[6] Second, issuing a forecast reduces the expected settlement amount or damage award if the firm is ultimately sued. This reduction occurs because the information

[6] A firm has a legal duty to disclose prospective material information when it updates or corrects a prior announcement which is still "alive." The precise circumstances under which a firm is required to make a disclosure is the subject of an ongoing legal debate. See Bauman [1979] and Sonsini and Berger [1993].

release often marks the end of the class period, which determines the eligibility of shareholders to recover damages (Skinner [1997]). The same level of benefits is not expected to accrue to firms issuing bad-news forecasts privately since a private forecast is unlikely to provide a defense against charges of withholding material information and is less likely to mark the end of the class period. Accordingly, we predict that litigation risk will be positively associated with the likelihood of issuing a public forecast when the gap is positive.

When the market's expectations roughly equal the firm's expectations, issuing a public forecast is unlikely to lead to a substantial reduction in litigation costs. The reason is that since the forecast does not provide the market with any additional material information, the forecast will be less effective in providing a defense against charges of withholding material misinformation. Since "confirming" forecasts do not lead to large reductions in stock prices (Clement et al. [2003]), the forecast date is unlikely to mark the end of the class period, and therefore will not affect the total amount of damages. In addition, issuing a confirming forecast can actually trigger a future lawsuit if the forecast turns out to be optimistic *ex post*. Even if the forecast were an unbiased assessment of the firm's prospects at the time it was made, it is difficult for firms to defend against *ex post* allegations that the forecast was inaccurate or incomplete, and therefore, misleading (Frankel et al. [2004]). Thus, when the expectations gap is small, we expect that high litigation risk firms are less likely to issue a forecast because the relative net benefits are smaller (more negative).

For similar reasons, issuing a public forecast when there is a negative expectations gap is also unlikely to reduce expected litigation costs since firms generally are not sued for withholding good news. Thus, the release of good news is unlikely to shorten the class period.

In addition, firms may increase the risk of litigation if they issue forecasts that raise the market's expectations of earnings but later fail to meet those expectations.

The above arguments suggest that when a forecast does not contain bad news, firms do not reduce their expected litigation costs, and indeed, may actual increase them when they issue a public forecast. However, these costs are unlikely to be incurred if firms issue good news or confirming forecasts privately to analysts and/or institutional investors. Forecasts disclosed privately are unlikely to result in litigation if the forecast subsequently turns out to have been too optimistic; furthermore, there is no legal obligation to update such forecasts. Therefore, we expect that prior to Regulation FD, high litigation risk firms with either good news or no news are more likely to issue their forecasts privately rather than publicly.

While Regulation FD prohibits the selective disclosure of material information, it does not impose any new disclosure requirements on firms. For this reason, we do not expect Regulation FD to affect disclosure decisions about information that firms would either prefer to disclose publicly or not disclose at all. Thus, to the extent that the regulation affected a firm's decision about whether to issue a public forecast, it will be with respect to a forecast that the firms would have preferred to issue selectively.[7] If the practice of selective disclosure had been relatively widespread and firms generally comply with the regulation, then we expect that firms are more likely to issue public forecasts after Regulation FD. This increase occurs because firms will choose to publicly disclose by means of earnings forecasts at least some of the information that they would have preferred to disclose selectively. The benefits of public disclosure compared to non-disclosure include satisfying the increased demand for information by analysts

[7] It is possible that a firm which preferred public disclosure prior to Regulation FD may prefer non-disclosure afterwards (or vice versa) as other firms in their industry switch from selective disclosure to either public disclosure or non-disclosure. However, we expect such instances to be relatively rare and ignore this possibility in developing our hypotheses.

and reducing the level of information asymmetry (King et al. [1990]). Consistent with these arguments, Brown et al. [2003] and Heflin et al. [2003] find that the number of forecasts increased significantly after Regulation FD. Accordingly, we test the following hypothesis:[8]

H1: The association between the *ex ante* risk of litigation and the propensity to issue a management forecast is positive when the expectations gap between the market and the firm is positive (i.e., when the market is unduly optimistic). Otherwise, the relation will be negative. In addition, firms are more likely to issue forecasts after Regulation FD.

2.2 Litigation Risk and Forecast Characteristics

As discussed in King et al. [1990], once managers have decided to issue a forecast, they must then determine the characteristics of the forecast. In this section, we discuss how litigation risk is expected to influence the choice of whether to release good or bad news, the forecast's precision, and its time horizon.

Prior to Regulation FD, firms had the option of issuing their forecasts publicly or privately. As discussed above, Francis et al. [1994] and Skinner [1994] argue that, for firms with high litigation risk, the benefits associated with public bad-news forecasts are higher than those associated with good-news forecasts, because releasing bad news is (more) effective in reducing litigation costs. Evidence in Kasznik and Lev [1995] is consistent with these arguments.[9] In addition, good-news forecasts disclosed privately to analysts and institutional investors are unlikely to result in higher future litigation costs if the forecast subsequently turns out to have been too optimistic. Therefore, we expect that conditional on the decision to issue a public forecast, the news in forecasts made by high litigation risk firms will contain relatively worse news compared to forecasts made by firms with low litigation risk.

[8] All hypotheses are stated in alternative form.

[9] Kasznik and Lev [1995] examine a sample of firms with extreme earnings surprises and find that bad-news firms released significantly more voluntary disclosures than did good-news firms.

After Regulation FD, we expect that firms will increase the relative number of good-news forecasts since such forecasts were more likely to have been issued privately prior to Regulation FD. In addition to litigation cost arguments, public good-news forecasts may entail greater proprietary costs than public bad-news forecasts. This asymmetry in proprietary costs occurs if potential competitors are more (less) likely to enter the firm's markets when those markets prove to be more (less) profitable than expected. In this case, proprietary costs associated with good-news forecasts are reduced when the information is released privately and filtered through analysts. Therefore, we expect that the propensity for firms to issue good-news forecasts is higher after Regulation FD, as relatively more good-news forecasts are issued publicly instead of privately. Accordingly, we test the following hypothesis:

H2: Conditional on issuing a forecast, litigation risk is negatively associated with good-news forecasts. Additionally, forecasts will contain more good news after Regulation FD.

We expect litigation risk to be negatively associated with forecast precision for two reasons. First, the expected costs of missing a forecast are increasing with litigation risk. High litigation risk firms are more likely to be sued when they miss their own public forecasts and less precise forecasts are more likely to be met than more precise forecasts (assuming the less precise forecast encompasses the more precise forecast). Second, the prior literature suggests that bad-news forecasts are viewed by the market as being highly credible, regardless of their precision. Since, as we argue above, forecasts by high litigation risk firms are more likely to contain bad news, there are no additional benefits from increasing their precision. Evidence in Hutton et al. [2003] and Hughes and Pae [2004] is consistent with this argument.

Regulation FD will influence forecast precision in two ways. As discussed above, forecasts will convey good news relatively more frequently after Regulation FD. Since such forecasts are less credible than are bad-news forecasts, firms will boost their credibility by increasing their precision (Hutton et al. [2003]). This factor suggests that average forecast precision should increase after Regulation FD. On the other hand, firms with high proprietary costs would have preferred to disclose information selectively. Proprietary costs associated with their forecasts are reduced when the information is released privately and filtered through analysts. In order to mitigate the associated proprietary costs associated with public forecasts, firms will reduce their forecast precision. Bamber and Cheon [1998] find evidence consistent with high proprietary cost firms issuing less precise forecasts.[10] With the prohibition of selective disclosure, these firms will opt to issue low precision public forecasts in order to reduce the associated proprietary costs, although doing so will reduce the forecast's credibility, and hence, market impact. Since these two economic forces predict opposite changes in forecast precision, the impact of Regulation FD on precision is ultimately an empirical question.[11] Therefore, we test the following hypothesis:

H3: Conditional on issuing a forecast, litigation risk is negatively associated with forecast precision. Additionally, the average level of precision will change after Regulation FD.

Conditional on the decision to issue an earnings forecast, we expect that litigation risk will be associated with earlier forecasts for two reasons. First, issuing a forecast earlier rather than later will bolster the firm's defense against charges that they failed to disclose required information promptly. Second, as the release of the forecast often marks the end of the class period, an earlier forecast will reduce the expected amount of damages by shortening the class period. Thus, litigation risk will cause firms to release their forecasts earlier, which results in longer forecast horizons.

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[10] Specifically, they find that firms in highly concentrated industries are significantly more likely to issue less precise forecasts. In the debate surrounding Regulation FD, many analysts feared that firms would reduce the precision of their public disclosures (Business Week Online, Sept. 18, 2000).

[11] In a univariate analysis, Heflin et al. [2003] document an increase in forecast frequency across all precision levels immediately after Regulation FD, with the biggest increase (percentage and absolute) coming for range forecasts. They do not examine whether the average level of precision changed.

Prior to Regulation FD, firms could at least partially satisfy the information demands of analysts and institutional investors through selective disclosure. With this option eliminated, these sophisticated investors will demand more and earlier public disclosure to partially offset the decline in selective disclosure. Firms will respond to this increased demand by releasing their forecasts earlier. Heflin et al. [2003] document that while the mean number of current earnings and future earnings forecasts per firm-quarter both increase significantly after Regulation FD, the increase is substantially larger for earnings forecasts beyond the current quarter. These results are consistent with a post-Regulation FD increase in horizon.[12] Based on the above arguments, we test the following hypothesis:

H4: Conditional on issuing a forecast, litigation risk is positively associated with the horizon of the forecast. Additionally, horizon will increase after Regulation FD.

3. Methodology

In this section, we describe the primary cross-sectional regression model we use to test the hypotheses developed in Section 2. We first discuss the regression model and the primary and control variables. We then describe our sample and provide summary statistics.

In order to examine the influence of litigation risk on firms' forecasting behavior, we construct an *ex ante* measure of a firm's risk of being sued. We model the probability of a firm being sued by its shareholders in a particular calendar quarter as a function of several predictors measured before the quarter begins. A firm-quarter is identified as a lawsuit quarter if the end of the class period of a federal class action securities lawsuit falls in that quarter. Lawsuits and class periods are identified from the Stanford Securities Class Action Clearinghouse and Woodruff-Sawyer databases. We use the predicted values from this model as our *ex ante*

[12] However, Heflin et al. [2003] did not analyze whether horizon changed conditional on the type of forecast. In contrast, Bailey et al. [2003] find weak evidence of a decrease in the average number of forecasts about future earnings using a similar sample. Their findings suggest a decrease in horizon after Regulation FD.

measure of litigation risk, *LitProb*. The specifics of the litigation risk model are described in Appendix B. Unlike Field et al. [2003], we do not exclude lawsuits that are less likely to be affected by firm's forecast decisions, e.g., regarding an aborted takeover attempt. To the extent that forecast decisions are unaffected by these types of lawsuits, the power of our tests should be reduced, but the litigation variables should not be biased.

Creating a single litigation risk variable offers two main advantages over an alternative approach whereby the litigation risk explanatory variables are included as separate predictor variables in the management forecast regressions. First, the predictor variables are likely associated with firms' forecasting decisions for reasons unrelated to litigation risk. For example, while larger firms are more likely to be sued, we expect that size will also influence forecasting behavior because it is associated with differences in the information environment and investors' demand for disclosure. Thus, the coefficient on firm size would be difficult to interpret. Second, a single variable representing litigation risk greatly facilitates the interpretation of the results, since only the sign and significance of a single variable need to be examined. It is difficult to interpret the net effect of litigation risk on forecasting behavior when certain litigation determinants have the predicted associations with forecasting decisions, while other determinants have the wrong associations.[13]

3.1. Primary Regression Specification

Equation (1) models the probability of a firm issuing a management forecast in a particular calendar quarter as a function of explanatory variables observable at the beginning of the quarter. Our sample begins in 1996 and extends through 2002; it includes all NYSE, AMEX, and Nasdaq firms with the necessary data to calculate the variables used in the analyses.

[13] For example, Table 3 in Frankel et al. [2004] shows that while certain litigation risk proxies are significant and have the correct signs, two variables (*QROA* and *QNEGEARN*) are significant but have the incorrect signs.

$$\begin{aligned} Forecast = G(&FD, LitProb,\ ExpGap^{+},\ LitProb{*}ExpGap^{+}, lag(Forecast), \\ &Concentration,\ Size,\ \left|EarnChg\right|,\ FCF,\ Analysts,\ Constant) \end{aligned} \tag{1}$$

where G(·) is the logistic cumulative distribution function, and firm and time subscripts are understood. *Forecast* is an indicator variable that equals one if the firm issued at least one forecast during the quarter, and zero otherwise. *FD* is an indicator variable that equals one if the calendar quarter comes after Regulation FD became effective in October 2000 (including the fourth quarter of 2000), and zero otherwise. *LitProb* is the *ex ante* measure of litigation risk discussed in Appendix B. The expectations gap, *ExpGap*, is a proxy for the level of firm-investor information asymmetry and is measured as the difference between the market's expectations regarding next quarter's earnings (as of the beginning of the calendar quarter) and the firm's expectation. The former is measured using the consensus analyst forecast (when available) or a seasonal random walk model. The latter is based on the actual level of earnings. Based on this measure, we create an indicator variable, $ExpGap^{+}$, that equals one if (*ExpGap*/*Price*) > 0.001. The control variables are discussed below.

3.2. Management Forecasts

Data on management forecasts is obtained from First Call. In addition to recording the forecast date, First Call collects and categorizes detailed information about each forecast. We use this information in our analyses of the forecast characteristics. Consistent with most previous studies, we exclude all non-earnings forecasts, such as EBITDA and cash flow forecasts.[14]

Table 1 presents descriptive statistics regarding the 23,762 management forecasts made on 18,618 firm-days between 1996 and 2002. Panel A shows that just over half of all forecasts consist of range forecasts, while a quarter are point forecasts. The remaining 23% consist of

[14] Excluded forecasts represent approximately 5% of total forecasts. Untabulated analyses indicate that our results are qualitatively unchanged when we include non-earnings forecasts.

qualitative forecasts and open range forecasts (minimum and maximum). Panel B presents the number of quarterly and annual forecasts in each year of the sample. The results clearly show a rapid increase in the number of forecasts over time. While there were less than 1,000 total forecasts in 1996 and 1997, there were over 7,500 forecasts in 2001 and 6,300 forecasts in 2002. The largest increases in both annual and quarterly forecasts took place in 2001, which is the first post-Regulation FD year. This finding is consistent with our hypothesis that firms are more likely to issue a forecast after Regulation FD. Panel B also shows that while the mean forecast horizon increases over time for quarterly forecasts (from 57 days in 1996 to 76 days in 2002), there is no clear pattern in the mean horizon for annual forecasts (249 days for entire sample period).

3.3 Control Variables

To isolate the effect of litigation risk on forecast decisions, we include several variables in the regression to control for non-litigation-related costs and benefits of issuing management forecast that are also likely to be correlated with litigation risk. Under the Expectations Adjustment Hypothesis (Ajinkya and Gift [1984], King et al. [1990]), firms benefit from issuing a management forecast through a reduction in the level of information asymmetry. Therefore, we expect a positive coefficient on $ExpGap^{+}$.

We include the lagged value of *Forecast*, *lag(Forecast)*, to control for firm-specific factors omitted from eq. (1) that also influence the disclosure decision. Firms are required to update a previous forecast under Rule 10b-5 of the 1934 SEC Act when it is still alive but becomes inaccurate, incomplete, or misleading (Bauman [1979], Loss and Seligman [1991]). Therefore, *lag(Forecast)* also reflects the firm's increased likelihood of issuing a forecast during the current quarter as a result of its duty to update a prior forecast. For these reasons, we expect a positive coefficient on *lag(Forecast)*. By including this variable, we are being conservative and likely biasing against finding a significant association between *LitProb* and *Forecast*, because *LitProb* will be positively correlated over time. Therefore, if high litigation risk induced the firm to issue a forecast during the previous period, the association between *LitProb* and *Forecast* during the current period will be weaker when *lag*(*Forecast)* is included in the regression.

Size and *Analysts* capture the demand for information by investors. As firm size and analyst following increase, we expect investors to demand more disclosure. Bushee et al. [2003] find evidence consistent with firms responding to investors' demand for information. *Size* is defined as the natural log of the firm's market value of equity (in $millions) measured at the end of quarter *t-1* using data from CRSP. *Analysts* is the number of analysts providing earnings forecasts during the calendar quarter for the next closest fiscal quarter end using data from First Call. Consistent with the results in Cox [1985] and Waymire [1985], we expect both variables to be positively associated with the probability of issuing a forecast.

Concentration and |*EarnChg*| capture non-litigation costs of issuing a forecast. *Concentration* is the industry concentration ratio, computed as sales of the largest five firms in the firm's industry divided by total industry sales, based on the industry classification scheme in Fama and French [1997]. *Concentration* proxies for the proprietary costs associated with issuing a forecast. Since proprietary costs are likely to be higher in more concentrated industries (Bamber and Cheon [1998]), we expect a negative coefficient on *Concentration*. Firms with more volatile earnings face a greater risk of inaccurate forecasts, along with the associated reputation costs. Evidence in Waymire [1985] is consistent with managers making forecasts when the volatility of earnings is lower. Therefore, we expect a negative coefficient on

|*EarnChg*|, which is measured as the absolute value of the seasonally differenced EPS deflated by stock price using data from Compustat.[15]

FCF is defined as the sum of operating and investing cash flows scaled by total assets using data from Compustat. This variable is an *ex ante* inverse measure of firm's need to access the external capital markets. Firms benefit from issuing forecasts before raising capital because it reduces the cost of capital via a reduction in information asymmetry. Since firms that are more likely to require external capital will benefit most from the reduction in asymmetry, *FCF* is expected to have a negative coefficient. Evidence in Frankel et al. [1995] and Ruland et al. [1990], which are based on *ex post* measures of raising capital, is consistent with this prediction.

3.4. Sample Description

Descriptive statistics for the primary variables are presented in Table 2. Panel A shows that the estimated level of litigation risk in a given quarter is relatively low, with a mean (median) of 1% and the 95th percentile at 3%. On average, 19% of the firms in our final sample issue a forecast in a given quarter while analysts provide a mean (median) of 3.45 (2) forecasts each quarter. The mean (median) absolute change in earnings is 5% (1%) of share prices, while both the mean and median values of *FCF* (free cash flow) are negative. $ExpGap^{+}$ equals one for 42% of firm-quarter observations, indicating that actual earnings are substantially below the market's expectations almost half the time.

Panel B provides similar descriptive statistics for a sub-sample restricted to firm-quarters with a management forecast. Comparing Panels A and B indicates that forecasting firms tend to have higher levels of litigation risk (in the upper tail), are larger, have higher analyst following, and have higher free cash flows. The latter result is contrary to our expectations but is consistent

[15] Similar to Baginski et al. [2002], our definition of earnings change captures the current level of volatility and minimizes the amount of survivorship bias associated with using a longer time-series to measure earnings volatility.

with our multivariate results discussed below. In addition, forecast firms are more likely to have substantial positive expectation gaps (bad-news) compared to other firms. The distributions for *Concentration* and |*EarnChg*| are similar between the two samples, except in the right tails, where forecasting firms are in more concentrated industries and have less volatile earnings. In addition, forecasting firms experience negative cumulative abnormal returns around the forecast date (mean = -2.28%, median = -0.68%).

Panel C provides Pearson and Spearman correlations based on the entire sample of 121,336 firm-quarters. The results indicate that litigation risk is positively correlated with issuing a forecast. In addition, *LitProb* is positively correlated with *Size,* |*EarnChg*|, and *Analysts*, and is negatively correlated with *Concentration* and *FCF*. *Forecast* is positively correlated with *Size*, *FCF*, *Analysts,* and $ExpGap^+$ and negatively correlated with |*EarnChg*|. These results are all consistent with our expectations, with the exception of the positive correlation between *Forecast* and *FCF*.

4. Empirical Results

In this section, we present the results from our cross-sectional empirical analyses. We first examine how our measure of litigation risk is related to the probability of issuing a forecast. We then explore the relation between litigation risk and forecast characteristics, including the type of news (good or bad), precision, and horizon. These analyses incorporate an adjustment to correct for the self-selection bias that occurs because the decision to issue a forecast in the first place is endogenous (Heckman [1979]).

4.1. Probability of a Forecast

The results of estimating eq. (1) are presented in Table 3. The analyses are based on 78,375 firm-quarter observations. The psuedo-R^2 is 17%, which indicates that the model has reasonably good explanatory power. All reported *z*-statistics (and *t*-statistics) are based on Huber-White standard errors that allow for the lack of independence in errors on individual firm observations in different quarters.

As anticipated, firms with positive expectations gaps are significantly more likely to issue a forecast (*z*-statistic = 21.2), compared with firms with small or negative gaps (i.e., when *ExpGap* < 0.1% of stock price). These results partially support the Expectations Adjustment Hypothesis: when the market's earnings expectations are well above the firm's expectations, firms are significantly more likely to issue a forecast to reduce expectations. These results are consistent with those in Kasznik and Lev [1995].

Hypothesis H1 predicts that the role of litigation risk in forecasting decisions will depend on the sign of the expectations gap. Specifically, we expect that *Forecast* will be positively (negatively) associated with *LitProb* when $ExpGap^{+}$ equals one (zero). The results are not fully consistent with our hypothesis. The fact that $LitProb \times ExpGap^{+}$ is positive and significant (*z*-statistic = 2.0) supports our hypothesis that litigation risk is a more important factor in managers' decisions to issue a forecast when the market's expectations are unduly optimistic. However, contrary to our expectations, even when the market's expectations are pessimistic, litigation risk is positively associated with the likelihood of issuing a forecast, as indicated by the significant *LitProb* coefficient (*z*-statistic = 2.0). In summary, firms respond to high litigation risk by increasing their level of voluntary disclosure, rather than curtailing it; furthermore, the need to correct optimistic expectations increases the magnitude of the association.[16]

[16] We note that firms are more significantly more likely to issue forecasts after Regulation FD (*z*-statistic = 33.4), supporting and strengthening a similar finding in Heflin et al. [2003], whose sample period only extends through the second quarter of 2001. This finding provides further support for hypothesis H1.

Baginski et al. [2002] and Johnson et al. [2001] both find that firms in legal environments where the average level of litigation risk is lower are more likely to issue a forecast.[17] While these results imply that firms with higher litigation risk are less likely to issue a forecast, we find the opposite is true, as firms with higher litigation risk are more likely to issue a forecast. Together, these results suggest that (1) a change in the legal environment that increases the baseline level litigation risk is likely to result in an overall decrease in number and frequency of management forecasts, and (2) holding the legal environment constant, firms with higher litigation risk are more likely to issue a forecast irrespective of the type of earnings news, although firms are even more likely to issue a forecast when the news is bad. These findings imply that researchers must carefully control for the legal environment in studies of management forecasts, and by extension, other forms of voluntary disclosures.

Consistent with our expectations, firms are more likely to issue forecasts when the demand for information is higher (proxied by *Size* and *Analysts*). The coefficient on |*EarnChg*| is negative and marginally significant (*z*-statistic = -1.8), indicating that firms with more volatile earnings are less likely to forecast, consistent with Cox [1985] and Waymire [1985]. Of particular importance in the model is the control for forecast behavior during the prior quarter, which is positively associated with issuing a forecast during the current quarter (*z*-statistic = 51.5).

Forecasts of firms in more highly concentrated industries are expected to incur higher proprietary costs because their forecasts may lead to additional entry into the industry. This relation suggests that *Concentration* should be negatively associated with the propensity to issue a forecast; Bamber and Cheon [1998] find support for this prediction. However, we find no

[17] Baginski et al. [2002] find that firms in Canada are more likely to issue forecasts compared to firms in the highly litigious United States. Johnson et al. [2001] find that after the 1995 PSLRA, there was an increase in forecasting frequency and this increase was larger for firms with higher litigation risk in the pre-PSLRA period.

significant association between the decision to issue a forecast and *Concentration* (*z*-statistic = 1.2). Finally, the highly significant and positive *FCF* coefficient (*z*-statistic = 10.9) indicates that firms with higher free cash flow are more likely to provide forecasts. Since these firms are less likely to require external capital (Dechow et al. [1996]), these findings are inconsistent with the results in Frankel et al. [1995] and Ruland et al. [1990].[18] One explanation for this unexpected result is that free cash flows reflect good operating performance, and firms with good operating performance are more likely to issue forecasts. Another possibility is that the difference in results is driven by our use of an *ex ante* measure of the need to raise external capital versus their use of an *ex post* measure.

4.2. Forecast News

In this section, we examine the relationship between litigation risk and the type of news contained in the forecast using the following equation (with firm, forecast, and time subscripts understood):

$$\begin{aligned} AbnRet = \beta_0 + \beta_1 FD + \beta_2 LitProb + \beta_3 Concentration + \beta_4 Size \\ + \beta_5 \left| EarnChg \right| + \beta_6 Analysts + \beta_7 Annual + \beta_8 Mills + \varepsilon \end{aligned} \quad (2)$$

AbnRet is the cumulative size-adjusted return in the three days surrounding the management forecast (-1, 0, +1). In addition to *Concentration, Size,* |*EarnChg*|, and *Analysts,* we include two additional control variables. *Annual* is an indicator variable that equals one if the forecast is for annual rather than quarterly earnings to control for any systematic differences in information content. *Mills* is the inverse Mills ratio based on the model in eq. (1) and is included to account

[18] Frankel et al. [1995] find that when firms raise external capital, they are significantly more likely to issue a forecast, except in the periods up to nine months before the capital is raised.

for the self-selection bias inherent in analyzing data that is conditional on a forecast being made.[19]

We first estimate eq. (2) using all forecasts for which we have available data. However, since almost 40% of the forecasts in our sample are released concurrently with the quarterly earnings announcement, our measure of the news in the forecast (*AbnRet*) is potentially contaminated by the news contained in the earnings announcement. Therefore, we also estimate eq. (2) on a restricted sample that only contains forecasts that do not coincide with earnings announcements (within three days). The qualitative results are identical across the two samples.

Table 4 presents the results of these analyses. As predicted, we find that, on average, forecasts made by firms with higher litigation risk convey more negative news to the market (*t*-statistic = -5.8). This finding supports the arguments in Skinner [1994] that high litigation risk firms benefit by issuing bad-news management forecasts. In addition, we find that forecasts in the period subsequent to Regulation FD contain relatively more good news (or less bad news). The *FD* coefficients are significantly positive with *t*-statistics of 15.6 and 10.5, respectively. Both findings provide support for hypothesis H2. However, we cannot rule out the possibility that the greater propensity for good news forecasts was not the result of general improvements in economic conditions in the post-Regulation FD period. We also note that *Mills* is significantly positive, indicating that it is important to control for self-selection bias when examining management forecast characteristics.

The analysis reported in Table 4 uses abnormal returns around the time of the management forecast as a proxy for news contained in the forecast. However, managers cannot perfectly predict how the stock market will react to a given forecast. To the extent that managers

[19] To be consistent with Heckman [1979], *Mills* is based on a probit specification of equation (1); the results in Table 3 are based on a logit specification.

cannot perfectly predict the market's reaction, using *AbnRet* may introduce measurement error into the dependent variable. From this perspective, it is more appropriate to use a measure of news that is observable to management at the time of their forecast. Managers readily observe analysts' earnings forecasts, so the difference between the management forecast and the consensus analyst forecast is potentially a more accurate measure of the news management intends to convey to the market through the forecast. Accordingly, we re-estimate eq. (2), replacing *AbnRet* with *RevExp* as the dependent variable, where *RevExp* is the difference between the management forecast and the consensus analyst forecast (deflated by stock price and multiplied by 100). Eliminating observations with no analyst forecasts reduces the sample to 12,551 firm-quarter observations. The untabulated results are qualitatively similar to those reported in Table 4. In particular, *LitProb* has a significantly negative coefficient and the *FD* coefficient is significantly positive.

4.3. Forecast Precision

In this section, we examine how forecast precision is related to litigation risk and how it has been affected by Regulation FD. We measure *Precision* as (–(range of forecast)/stock price) × 100 and winsorize it at the 1% level for point and range forecasts. For open-ended and qualitative forecasts, we set *Precision* equal to the 1% winsorized level of range and point forecasts. Hypothesis H3 predicts that high litigation risk firms provide less precise forecasts in order to reduce the potential costs of litigation and that forecast precision may either increase or decrease after Regulation FD, depending on whether the effect of more good-news forecasts (which have higher precision) or the effect of proprietary costs (which leads to lower precision) dominates.

We test hypothesis H3 using the following equation, with firm, forecast, and time subscripts understood:

$$\begin{aligned} Precision = \phi_0 + \phi_1 FD + \phi_2 LitProb + \phi_3 Concentration + \phi_4 Size \\ + \phi_5 \left|EarnChg\right| + \phi_6 Analysts + \phi_7 Horizon + \phi_8 Mills + \mu \end{aligned} \tag{3}$$

where *Horizon* controls for longer-range forecasts generally being less precise. We examine quarterly and annual forecasts separately to control for other structural differences between the two types of forecasts.

Table 5 presents the results from this analysis. Interestingly, the model's explanatory power for Annual forecasts is over three times as large as that for Quarterly forecasts (psuedo-R^2 = 25.3% and 7.3%, respectively). The coefficient on *LitProb* is significantly negative for both quarterly and annual forecasts (*t*-statistic = -2.6 and -4.3, respectively). This finding supports our prediction that litigation risk is negatively associated with forecast precision.

The strength of the effect of Regulation FD itself differs between quarterly and annual forecasts. For quarterly (annual) forecasts, the *FD* coefficient is 0.50 (0.03) and is highly significant (insignificant) with a *t*-statistic = 12.7 (1.2). These results indicate that forecast precision increased after Regulation FD only for quarterly forecasts. The predicted sign of the *FD* coefficient is ambiguous because of the potentially confounding effects of good-news forecasts and proprietary costs on precision. The finding that precision increased after Regulation FD suggests that the increase in the amount of good-news forecasts dominates, especially for quarterly forecasts. These results are consistent with the results in Table 4, where forecasts after Regulation FD conveyed significantly more good news. Note that including *Concentration* in the regression may partially control for the hypothesized negative effect of proprietary costs on precision. To examine this possibility, we re-estimate eq. (3) excluding *Concentration*. The *FD* coefficient magnitudes and significance levels are virtually unchanged from the results in Table 5. With respect to the disparity in the strength of the quarterly and annual results, one possible explanation is that proprietary costs are higher for annual forecasts due to their longer horizons, and this effect is not fully captured by *Horizon* (Dontoh [1989]).[20]

The results for the control variables show that larger firms provide more precise forecasts, which is consistent with our prediction that such firms have higher demands for earnings-related information. However, the same argument with respect to analyst following is only supported for annual forecasts (*t*-statistic = 2.2), as the coefficient on *Analysts* in the quarterly forecast regressions is insignificant. The negative and generally significant coefficients on |*EarnChg*| and *Horizon* indicate that firms issue less precise forecasts when making accurate earnings forecasts is more difficult.[21] Furthermore, the *Mills* coefficient highly significant in the annual forecast equation, indicating the importance of controlling for the self-selection bias when examining forecast precision.

We find no evidence that *Concentration* is significantly related to forecast precision (*t*-statistic = 1.6 and 1.4). In contrast, Bamber and Cheon [1998] find that, based on a sample of annual forecasts made by 151 NYSE firms between 1981-1991, the industry concentration ratio is significantly and negatively associated with forecast precision. Aside from the substantial differences in sample composition and forecasting behavior between the two sample periods, one possible explanation for the different results could be our use of a continuous measure of precision, whereas they used four precision categories in an ordered logit specification. To investigate this possibility, we replicate the precision analysis using an ordered logit specification. However, *Concentration* is never significantly negative, and is sometimes significantly positive depending on the specification. One possible explanation for this finding is

[20] The results for Table 6 discussed below show that firms in highly concentrated industries, and hence, with high proprietary costs, have shorter annual forecast horizons, which is consistent with Dontoh [1989].

[21] When comparable, our results are generally consistent with those in Baginski and Hassell [1997] with the exception that, while they find size negatively related to precision, we find that larger firms tend to issue more precise forecasts.

that firms in more concentrated industries issue more precise forecasts to forestall entry by potential competitors. In any event, we are unable to reconcile our results with those of Bamber and Cheon [1998].

4.4. Forecast Horizon

Our final analysis of forecast characteristics examines the effect of litigation risk and Regulation FD on the choice of forecast horizon. Hypothesis H4 predicts that high litigation risk firms will issue their forecasts earlier and that forecast horizon will increase after Regulation FD as firms respond to analysts' calls for more timely disclosure. We estimate the following equation to test hypothesis H4, where firm, forecast, and time subscripts are understood:

$$\begin{aligned} Horizon = \lambda_0 + \lambda_1 FD + \lambda_2 LitProb + \lambda_3 Concentration + \lambda_4 Size \\ + \lambda_5 \left| EarnChg \right| + \lambda_6 Analysts + \lambda_7 Mills + \nu \end{aligned} \tag{4}$$

We separately analyze forecasts of quarterly and annual results, since the average horizon differs so dramatically between the two types (68 days for quarterly forecasts versus 249 days for annual forecasts). The results from these analyses are shown in Table 6.

Consistent with hypothesis H4, we find that forecast horizon is positively and significantly associated with litigation risk in both regressions (t-statistic = 2.0 and 5.0, respectively). However, the economic significance of this relation is relatively small, especially for quarterly forecasts. A one percentage point increase in *LitProb* results in an average increase in *Horizon* of 0.55 days for quarterly forecasts and 5.7 days for annual forecasts.

Similar to the Precision results reported in Table 5, the results for the *FD* coefficients also differ for quarterly and annual forecasts. After Regulation FD, the average quarterly forecast horizon increases by 17 days, which is consistent with the descriptive statistics in Table 1. This finding is consistent with the decrease in forecast precision found in Table 5 and represents an economically important change in quarter forecast horizon (relative to the sample mean of 68 days). It is also consistent with more firms adopting regular disclosure policies after Regulation FD, whereby they issue a forecast of the next quarter's earnings at the same time they announce the current quarter's results (Brown et al. [2003]). However, Table 6 unexpectedly shows that the horizon of annual forecasts decreases by almost 9 days after Regulation FD. However, further (untabulated) analyses indicate that this result is driven by a lack of available data on actual earnings for forecasts made in 2002, which artificially curtails the average horizon for firms remaining in the sample. When we exclude observations from the last quarter of 2002, the coefficient on *FD* becomes significantly positive.

Table 6 also shows that while *Concentration* is not significant in the quarterly regressions, it is significantly negative in the annual regressions (*t*-statistic = -3.7). If industry concentration is a positive proxy for the proprietary costs associated with earnings forecasts, then this finding suggests that only longer-term forecasts are associated with significant proprietary costs. This interpretation is consistent with the analysis in Dontoh [1989], where earlier forecasts better enable competitors to adjust their production schedules at the expense of the forecaster.

5. Summary and Conclusions

This study examines how the *ex ante* level of litigation risk influences firms' forecasting decisions. As a part of this analysis, we examine how forecasting behavior changed after Regulation FD. Our analyses are based on almost 80,000 firm-quarter observations, including roughly 24,000 management forecasts made after the Private Securities Litigation Reform Act of 1995.

We find that firms are more likely to issue a management forecast during the quarter when the level of litigation risk measured at the beginning of the quarter is higher. This finding holds regardless of the market's expectations of quarterly earnings vis-à-vis the actual earnings, although the relation is stronger when the market is unduly optimistic. This finding suggests that managers believe the expected net benefits of issuing a forecast are higher when the risk of being subject to class-action securities litigation is higher. These benefits arise as issuing a forecast reduces the probability of litigation (Field et al. [2003]) and/or reduces the expected damages conditional on being sued (Skinner [1997]). In addition, we find that conditional on the decision to issue a forecast, litigation risk is positively associated with longer forecast horizons, increases the probability that the forecast contains bad news, and is negatively associated with forecast precision. Consistent with Heflin et al. [2003], we find that the likelihood of issuing a forecast increased significantly after Regulation FD. In addition, the average news contained in the forecast became significantly less negative, forecast precision and horizon significantly increased after Regulation FD. All of these results are consistent with our hypotheses.

These findings enhance our understanding of the complex role of litigation risk in influencing firms' voluntary disclosures. Baginski et al. [2002] and Johnson et al. [2001] both find evidence indicating that firms in legal environments with a lower baseline level of litigation risk are more likely to issue a forecast. While these results imply that firms with higher litigation risk are less likely to issue a forecast, we find the opposite is true, as firms with higher litigation risk are in fact more likely to issue a forecast. Together, these results suggest that (1) a change in the legal environment that reduces the underlying level of litigation risk is likely to result in more management forecasts, and (2) holding the legal environment constant, firms with higher litigation risk are more likely to issue a forecast, especially when the expectations gap is positive. These findings suggest that researchers must carefully control for the legal environment in studies of management forecasts, and by extension, other forms of voluntary disclosures.

Appendix A: Variable Definitions

Forecast is an indicator variable that equals one if a firm issued one or more forecasts during a calendar quarter and zero otherwise. *Lag(Forecast)* is the value of *Forecast* during the prior calendar quarter.

FD is an indicator variable that equals one if the firm-quarter is after the third quarter of 2000, and zero otherwise.

LitProb is an *ex ante* estimate of the probability the firm will be subject to a class-action securities lawsuit whose class period ends during the following calendar quarter. *LitProb* is the fitted value from the logistic regression model described in Appendix B.

Concentration is the industry sales concentration ratio, computed as the ratio of sales of the top five firms in the industry to the total industry sales during the quarter using the industry classification of Fama and French [1997].

Size is the natural logarithm of the market value of equity at the beginning of the quarter.

|*EarnChg*| is the absolute value of the seasonally differenced earning per share.

FCF is the sum of cash flows from operations plus cash flows from investing scaled by total assets.

Analysts is the number of analysts covering the firm measured as the number of analysts providing earnings forecasts during the calendar quarter.

$ExpGap^{+}$ is an indicator variable that takes on a value of one when [(market expectation – actual EPS)/Price] > 0.001, where the market's expectation of quarterly earnings per share (measured at the beginning of the calendar quarter) is based on the consensus analyst forecast if available and otherwise is based on a seasonal random walk model.

AbnRet is the three day, cumulative size-adjusted abnormal returns around the forecast date.

Annual is an indicator variable that equals one if the management forecast is a forecast of annual earnings, and zero otherwise.

Precision is a measure of the precision of point and range forecasts and is defined as (-(range of forecast)/price)×100. For open-ended and qualitative forecasts, *Precision* is set equal to the 1% winsorized level of range and point forecasts.

Horizon is the number of days between when the management forecast is released and when the actual value of the forecasted item is announced.

Mills is the inverse Mills ratio based on a probit estimatation of the model in Table 3.

Appendix B: Litigation Risk Model

We use the following model of litigation risk in our analyses:

$$\begin{gathered}Prob(Sued = 1) = \mathrm{G}(Constant, Year\ Indicators, Size, Turnover, Skewness, Proceeds, \\ BSM\text{-}Prob, Big\ N, InstOwnT4, EarnChange)\end{gathered} \tag{5}$$

where G(·) is the logistic cumulative distribution function, and firm and time subscripts are understood. Sued is an indicator variable that equals one if the subsequent calendar quarter contains the end of the class period if a securities class action lawsuit is subsequently filed against the firm, and zero otherwise. *Size* is the natural logarithm of equity market value (in $millions) at the beginning of the quarter. *Turnover* is the average daily trading volume (in shares) divided by number of shares outstanding over the proceeding quarter. *Skewness* is the skewness of the firm's stock returns over the preceding six months. *Proceeds* is the natural logarithm of the amount of public capital (debt and equity) raised over the preceding two years. *BSM-Prob* is the risk-neutral probability of bankruptcy derived from the Black-Scholes-Merton option pricing model and estimated according to the method in Hillegeist et al. [2004]. *Big N* is an indicator variable that equals one if the firm's auditor is one of the Big N international audit firms, and zero otherwise. *InstOwnT4* is the fraction of shares outstanding held by the top four institutional investors. *EarnChange* is the seasonally differenced EPS deflated by stock price. We expect *Size, Turnover, Proceeds, BSM-Prob*, and *Big N* to be positively associated with litigation risk and *Skewness, InstOwnT4,* and *EarnChange* to have negative associations.

Most of the explanatory variables in equation (5) are similar to those used in prior work on litigation prediction, including Johnson et al. [2000] and Jones and Weingram [1996]. However, we are unaware of any previous studies that have previously used *BSM-Prob* and *InstOwnT4* to predict litigation, and therefore discuss our motivation for including them. *BSM-Prob* measures the risk of financial distress, defined as when the market value of assets is less than the book value of liabilities. *BSM-Prob* primarily consists of a non-linear combination of a market-based leverage variable and asset volatility. When *BSM-Prob* is high, the decrease in equity value for a given decline in asset value will be proportionately greater. Thus, high *BSM-Prob* firms are more likely to sustain large negative returns, which often prompt litigation. Additionally, *BSM-Prob* proxies for the likelihood of violating debt covenants, as managers have greater incentives to engage in accounting fraud, which frequently results in litigation (Dechow et al. [1996]). The percentage ownership by the top four institutional owners captures the amount and effectiveness of outside monitoring. Institutional ownership concentration is likely a good proxy for monitoring because only institutions with large ownership stakes are willing to engage in costly monitoring activities. More intensive outside monitoring inhibits managers from taking self-interested or short-term actions that may later result in litigation.[22] Thus, *InstOwnT4* will be negatively associated with litigation risk.

We use equation (5) to estimate the probability of a class-action securities lawsuit being initiated against the firm over the calendar quarter. Our sample is based on all NYSE, AMEX, and Nasdaq firms between 1996 and 2002. Our sample consists of 121,238 firm-quarter observations that were subject to 924 lawsuits. Only 25 firms were sued more than once in a 24 month period.

The results from estimating this logit model are presented in Table A. Table A shows that all of the explanatory variables have significant coefficients that have the predicted signs. The pseudo-R^2 is 14.4%, which indicates a reasonably good fit of the model, considering the

[22] Consistent with this reasoning, Hartzell and Starks [2003] find that institutional ownership concentration is positively related to the pay-for-performance sensitivity of executive compensation and negatively associated with total compensation. These results suggest that institutions with large ownership stakes act as effective monitors.

infrequency of shareholder lawsuits.[23] To investigate the accuracy of the litigation risk estimates, we ranked all firm-quarters based on *LitProb* and sorted them into deciles. For each decile, we computed the percentage of all lawsuit quarters. The untabulated results show that 53% (19%) of all lawsuits occur in the decile with the (second) highest level of litigation risk. Furthermore, the lowest (five) litigation risk decile(s) had less than 1% (9%) of all lawsuits. Therefore, we expect that *LitProb* is a relatively good proxy for litigation risk.

[23] As a rough comparison, the litigation risk models in Frankel et al. [2004] and Rogers and Stocken [2003] have psuedo-R^2s of 9.45% and 12.26%, respectively. In addition, the litigation risk model in Johnson et al. [2001] has a model χ^2 of 65.71, compared to 1,531 for our model.

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Table A – Litigation Risk Model

Dependent variable is Sued, a binary variable that equals one if a firm was sued in Federal Court for securities related matters during a particular calendar quarter and zero otherwise. Sued = 1 for 924 firm-quarters and 0 for 120,314 firm-quarters. See Appendix 1 for variable definitions.

Variable	Predicted Sign	Coefficient Estimate	*z* -statistic
Size	+	0.14	6.43
Turnover	+	72.26	29.19
Skewness	–	-0.47	-13.28
Proceeds	+	0.17	11.74
BSM-Prob	+	2.72	14.05
Big N	+	0.40	2.67
InstOwnT4	–	-1.43	-4.31
EarnChange	–	-0.53	-2.31
Y1997		0.33	1.69
Y1998		0.64	3.46
Y1999		0.26	1.34
Y2000		1.24	7.00
Y2001		0.39	2.06
Y2002		0.61	3.22
Constant		-7.49	-33.07
n, pseudo-R^2		121,238	14.41%
number of clusters (firms)		8,478	

Table 1 – Characteristics of management forecast activity

days. Multiple items forecasted by a firm on a single day are treated as separate forecasts.

Panel A – Tabulation of forecasts by precision

Precision	Frequency	Percent
Point	5,848	24.6
Range	12,443	52.4
Single bound	1,750	7.4
Qualitative	3,721	15.7
Total	23,762	100.0

Panel B – Forecast horizon by periodicity and year

	Annual Forecasts		Quarterly Forecasts	
Year	Number of Forecasts	Mean Horizon	Number of Forecasts	Mean Horizon
1996	102	272	236	57
1997	309	233	633	47
1998	786	263	1,440	55
1999	1,165	275	1,781	61
2000	1,373	267	2,061	66
2001	3,012	262	4,521	74
2002	2,874	215	3,469	76
Total	9,621	249	14,141	68

Table 2 - Descriptive statistics of primary regression variables

See Appendix 1 for variable definitions.

Panel A – Based on 121,336 firm-quarter observations

Variable	Mean	Std. Dev.	5th Pctl	Median	95th Pctl
LitProb	0.01	0.01	0	0.01	0.03
Forecast	0.19	0.40	0	0	1.00
Concentration	0.35	0.13	0.21	0.33	0.58
Size	5.57	1.99	2.47	5.49	9.10
|EarnChg|	0.05	0.14	0	0.01	0.24
FCF	-0.05	0.21	-0.43	-0.01	0.19
Analysts	3.45	4.65	0	2	13
$ExpGap^{+}$	0.42	0.49	0	0	1

Panel B - Based on 15,946 forecasts

Variable	Mean	Std. Dev.	5th Pctl	Median	95th Pctl
LitProb	0.01	0.01	0.00	0.01	0.04
Concentration	0.36	0.14	0.21	0.33	0.68
Size	6.71	1.87	3.69	6.65	10.02
|EarnChg|	0.04	0.11	0.00	0.01	0.16
FCF	-0.003	0.14	-0.25	0.01	0.19
Analysts	6.10	5.95	0.00	4.50	18.00
Precision	-0.70	1.01	-4.17	-0.12	0.00
$ExpGap^{+}$	0.47	0.50	0.00	0.00	1.00
AbnRet	-2.28	12.18	-26.09	-0.68	15.20

Panel C – Correlation Matrix (Pearson correlations on lower diagonal; Spearman on upper)

All correlations are statistically significant at the 0.001 level except for those values in italics.
See Appendix 1 for variable definitions. n = 121,336 firm-quarters.

	LitProb	Forecast	Concentration	Size	|EarnChg|	FCF	Analysts	ExpGap+
LitProb		0.19	-0.08	0.40	*-0.01*	-0.06	0.39	*-0.01*
Forecast	0.13		*0.00*	0.22	-0.02	0.07	0.25	0.07
Concentration	-0.05	0.01		-0.10	0.02	0.03	-0.07	*0.01*
Size	0.24	0.22	-0.07		-0.44	0.12	0.66	-0.19
|EarnChg|	0.09	-0.03	*-0.01*	-0.31		-0.08	-0.24	0.25
FCF	-0.08	0.08	0.03	0.14	-0.08		0.05	-0.08
Analysts	0.31	0.25	-0.05	0.65	-0.11	0.09		-0.08
ExpGap+	*0.01*	0.07	0.02	-0.19	0.12	-0.06	-0.10	

Table 3 – Management forecast decisions and litigation risk

Dependent variable in logit regression is Forecast, an indicator variable that equals 1 if a firm made at least one forecast during a calendar quarter and zero otherwise. Forecast = 1 for 14,859 firm-quarters and 0 for 63,516 firm-quarters. *z*-statistics are based on Huber-White standard errors that allow for the lack of independence in errors on individual firm observations in different quarters. See Appendix 1 for variable definitions.

Variable	Predicted Sign	Coefficient Estimate	*z*-statistic
FD	+	0.77	33.4
LitProb	–	2.38	2.0
$ExpGap^+$	+	0.56	21.2
$LitProb*ExpGap^+$	+	3.04	2.0
lag(Forecast)	+	1.46	51.5
Concentration	–	0.13	1.2
Size	+	0.17	16.3
|EarnChg|	–	-0.15	-1.8
FCF	–	0.63	10.9
Analysts	+	0.05	12.8
Constant	?	-3.63	-51.0
n, pseudo-R^2		78,375	17.0%
number of clusters (firms)		6,442	
χ^2 ($ExpGap^+ = LitProb*ExpGap^+ = 0$), p-value		23.9	<0.0001

Table 4 – Forecast news and litigation risk

Dependent variable is AbnRet, the cumulative size-adjusted abnormal return in the three-day window surrounding the date of the forecast. Non-earnings announcement forecasts exclude forecasts issued during the three-day window surrounding the earnings annoucement date. *t*-statistics are based on Huber-White standard errors that allow for the lack of independence in errors on individual firm observations in different quarters. See Appendix 1 for variable definitions.

		All Forecasts		Non-Earnings Announcement Forecasts	
Variable	Predicted Sign	Coefficient Estimate	*t* - statistic	Coefficient Estimate	*t* - statistic
FD	+	3.32	15.6	2.92	10.5
LitProb	–	-43.37	-5.8	-59.42	-5.8
Concentration	?	-0.22	-0.3	-0.38	-0.4
Size	?	0.94	12.7	1.34	12.4
|EarnChg|	?	-10.4	-8.5	-13.3	-7.6
Analysts	?	-0.04	-1.7	-0.07	-2.3
Annual	?	1.41	10.2	1.45	6.9
Mills	?	1.64	6.1	2.44	6.5
Constant	?	-11.64	-15.8	-15.47	-15.0
n, Adj. R^2		18,150	7.5%	11,081	9.1%
number of clusters (firms)		2,905		2,735	

Table 5 – Forecast precision and litigation risk

Dependent variable is Precision, which is (-(range of forecast)/price)*100 for point and range forecasts and is set equal to the 1% winsorized level of range and point forecasts for open range and qualitative forecasts. Quarterly and Annual forecasts are treated separately. *t* -statistics are based on Huber-White standard errors that allow for the lack of independence in errors on individual firm observations in different quarters. See Appendix 1 for variable definitions.

		Quarterly Forecasts		Annual Forecasts	
Variable	Predicted Sign	Coefficient Estimate	*t* - statistic	Coefficient Estimate	*t* - statistic
FD	+/–	0.50	12.7	0.03	1.2
LitProb	–	-3.12	-2.6	-3.97	-4.3
Concentration	–	0.24	1.6	0.10	1.4
Size	+	0.10	6.8	0.18	16.8
|EarnChg|	–	-1.45	-8.4	-2.90	-12.4
Analysts	+	0.00	-0.8	0.00	2.2
Horizon	–	0.00	-1.4	0.00	-9.3
Mills	?	-0.08	-1.6	0.18	5.6
Constant	?	-1.73	-12.2	-1.70	-15.9
n, Adj. R^2		10,334	7.3%	6,846	25.3%
number of clusters (firms)		2,642		1,782	

Table 6 – Forecast horizon and litigation risk

The dependent variable is Horizon, which is defined as the number of days between when the forecast is issued and when the forecasted item is reported, where annual and quarterly forecasts are treated separately. *t*-statistics) are based on Huber-White standard errors that allow for the lack of independence in errors on individual firm observations in different quarters. See Appendix 1 for variable definitions.

		Quarterly Forecasts		Annual Forecasts	
Variable	Predicted Sign	Coefficient Estimate	*t* - statistic	Coefficient Estimate	*t* - statistic
FD	+	17.32	16.9	-12.67	-3.8
LitProb	+	54.58	2.0	568.7	5.0
Concentration	–	0.01	0.0	-38.0	-3.7
Size	+	1.13	3.2	4.45	3.7
|EarnChg|	–	2.44	0.7	-26.6	-1.2
Analysts	+/–	-0.17	-1.7	-0.40	-1.0
Mills	?	-6.27	-5.1	13.8	3.2
Constant	?	49.09	14.3	212.2	17.8
n, Adj. R^2		10,879	5.4%	7,509	1.2%
number of clusters (firms)		2,689		1,879	