

**Syracuse University**

---

**From the Selected Works of Lihong Liang**

---

January, 2013

# Analyst Pessimism and Forecast Timing

Orie E. Barron, *The Pennsylvania State University*

Donal Byard

Lihong Liang, *Syracuse University*



Available at: [https://works.bepress.com/lihong\\_liang/1/](https://works.bepress.com/lihong_liang/1/)

# **Analyst Pessimism and Forecast Timing**

Orie E. Barron  
The Pennsylvania State University

Donal Byard  
Baruch College—CUNY

Lihong Liang  
Syracuse University

January 2013

We thank Masako Darrough, John Elliott, Dan Givoly, Mike Hyman, Christina Mashruwala, Shamin Mashruwala, John Shon, Jim Vincent, Larry Walther and, especially, Yong Yu. We also thank seminar participants at Baruch College—CUNY, George Mason University, the University of Illinois—Chicago, the Pennsylvania State University, Syracuse University, and conference participants at the AAA Annual Meetings, and the EAA Annual Meetings. Finally, we thank I/B/E/S International Inc. for providing earnings per share forecast data used in this study.

# **Analyst Pessimism and Forecast Timing**

## **Abstract**

In this study, we show that on average relatively pessimistic analysts tend to reveal their earnings forecasts later than other analysts. Further, we find this forecast timing effect explains a substantial proportion of the well-known decrease in consensus analyst forecast optimism over the forecast period prior to earnings announcements, which helps explain why analysts' longer term earnings forecasts are more optimistically biased than their shorter term forecasts. We extend McNichols and O'Brien's (1997) and Hayes' (1998) theory concerning analyst self-selection to argue that analysts with a relatively pessimistic view—compared to other analysts—are more reluctant to issue their earnings forecasts, with the result that they tend to defer revealing their earnings forecasts until later in the forecasting period than other analysts.

**Keywords:** Analysts' forecast timing, Analysts' pessimism, Trading commissions.

## I. Introduction

On average, analysts' longer-term earnings forecasts tend to be more optimistic than analysts' shorter-term earnings forecasts (e.g., see O'Brien 1988; Brown 1997). In this study, we examine one potential explanation for the decrease in analysts' earnings forecast optimism over the period prior to annual earnings announcements: the possibility that individual analysts who hold *relatively* pessimistic views about upcoming earnings—compared to other analysts—choose to reveal their forecasts later in the forecast period.

Our research question is motivated by McNichols and O'Brien's (1997) and Hayes' (1998) research regarding analyst self-selection. They distinguish between *ex ante* and *ex post* optimism. *Ex post* optimism is optimism relative to the actual earnings that are eventually announced, i.e., forecasts that appear in retrospect to have been too optimistic. McNichols and O'Brien (1997) argue that the observed *ex post* optimistic bias in analysts' earnings forecasts (relative to actual earnings) results from analysts' self-selection of the firms they follow: analysts drop coverage of the firms they are *relatively* pessimistic about—*compared* to the other firms they follow. That is to say, analysts decide which firms to follow based upon their *ex ante level of optimism*: analysts compare their beliefs about a firm with their beliefs about other firms and drop coverage of firms they are *relatively pessimistic* about compared to the other firms they follow. In this setting, analysts' coverage decisions depend upon a benchmark that is known to analysts—their beliefs about other firms, and not the (unknown) future actual earnings. We extend McNichols and O'Brien's (1997) theory to explain analysts' forecast timing decisions. We argue that rather than completely dropping coverage of a firm, analysts with a relatively pessimistic outlook—compared to other analysts who follow the same firm—may simply be

more reluctant to issue forecasts, with the result that they tend to reveal their forecasts later than other analysts forecasting the same earnings.

Relatively pessimistic analysts have at least three potential economic incentives which could cause them to wait until later in the forecast period to reveal their forecasts. First, relatively pessimistic forecasts issued later in the forecast period are not as likely to alienate managers because these forecasts help managers beat the average forecast. Analysts may thus have an incentive to issue relatively pessimistic forecasts later in the forecast period in order to please management; pleasing management may provide analysts with better access to management information or may help an analyst's brokerage firm employer win lucrative investment banking business.

Second, according to Hayes (1998) model, optimistic and pessimistic forecasts do not have the same effect on investors' incentives to trade. Both optimistic and pessimistic forecasts reduce investor uncertainty. Viewed in isolation, reductions in uncertainty create an incentive for risk adverse investors to buy. As a result, a relatively optimistic forecast has two effects which both create an incentive for investors to buy — the relatively optimistic signal itself, and the decrease in uncertainty associated with the forecast. On the other hand, if the forecast is relatively pessimistic this creates two countervailing incentives — the information in the forecast itself creates an incentive to sell, whereas the decrease in uncertainty arising from the forecast creates an incentive to buy. Investors' incentive to trade stocks they already own based upon a relatively pessimistic forecast is thus at least partially offset by an incentive to continue to hold because of the reduction in uncertainty.

Third, Hayes (1998) predicts that relatively pessimistic forecasts are also less likely to generate trades and brokerage commissions because of the increased costs and risks of short

selling. Thus, because relatively pessimistic forecasts are less likely to generate trades for analysts' brokerage firms, analysts have less of an incentive to provide relatively pessimistic forecasts than relatively optimistic forecasts. Since the costs of short selling are likely to decrease as investors trading horizons shorten, analysts with relatively pessimistic views may be more likely to issue their relatively pessimistic views when investors trading horizons are shorter. In the case of analysts' earnings estimates, this would result in analysts having a greater incentive to issue relatively pessimistic forecasts closer to the earnings announcement date. Further, some analysts are likely assigned to cover certain firms; in these cases analysts may not have the option to drop or not provide coverage for a stock, even if they are relatively pessimistic about the firm compared to other analysts following the same firm. These analysts are likely to be more reluctant than other analysts to reveal their forecasts.

Our argument is essentially about individual analysts' behavior over the forecast period prior to earnings announcements. We assume that analysts have a sense of whether their beliefs are relatively pessimistic compared to current market prices or expectations (i.e., the expectations of other analysts following the same firm). This is based on the idea that analysts have to periodically update their buy/hold/sell recommendations by comparing their own beliefs with current market prices. Analysts cannot observe the actual earnings number prior to an earnings announcement date, so individual analysts cannot condition their forecast timing decisions on whether *ex post* their forecasts turn out to have been optimistic or pessimistic relative to actual earnings. Consequently, our argument is about analysts' optimism relative to the other analysts who follow the same firm, i.e., analysts' *ex ante* relative optimism, and not analysts' (absolute) *ex post* optimism compared to actual earnings.

< Insert Figure 1 About Here >

Figure 1 illustrates a stylized example of the effect we hypothesize. In this simple example there are six analysts forecasting earnings. The oval on the left hand side of the figure illustrates the distribution of all six analysts' beliefs regarding upcoming annual earnings 12 months prior to the earnings announcement date. We assume that some analysts start with beliefs regarding upcoming annual earnings that are relatively pessimistic (e.g., X5 and X6 in this example) compared to other analysts (e.g., X1 and X2 in this example). Analyst X1 is the most optimistic analyst; she is the first analyst to reveal a forecast. The relatively more pessimistic analysts (X2 to X6) then start forecasting later in the forecast period, on average, with the result that the observed average forecast decreases prior to the earnings announcement date. As relatively pessimistic analysts start issuing forecasts, the average forecast (mean of all forecasts outstanding from analysts) will decrease. Our basic research question is as follows:

Research Question: Do relatively pessimistic analysts start forecasting earnings later, on average, than other analysts?

Our results are consistent with our expectations, indicating that the forecast timing effect we document contributes to the decrease in the observed ex post optimism in analysts' average forecasts. Specifically, we compare the forecasts of annual earnings made by analysts in the last six months of the 12-month forecast period prior to annual earnings announcements. Using these forecasts, we compare the *relative pessimism* of forecasts made by analysts who start forecasting early (more than six months prior to the earnings announcement date), with that of forecasts made by analysts who start forecasting late (issue their first forecast in the last six months prior to the earnings announcement date). We show that the forecasts of late analysts — analysts who issue their first forecast in the last six months prior to the annual earnings announcement date — are relatively more pessimistic compared to concurrent forecasts made by early analysts (those

who start forecasting earlier in the 12-month forecast period). We show that this particular type of analyst self-selection can explain a significant portion of the over-time decrease in the optimistic bias in analysts' average forecasts.

We also find that this forecast timing effect is stronger in the post Regulation Fair Disclosure (Reg FD) period. Over our sample period (1989 to 2010) we find that approximately 40% of the typical decrease in average forecast optimism over the 12-month forecast period prior to annual earnings announcements is due to relatively pessimistic analysts forecasting later in the forecast period; this increases to 50% in the post Reg FD period. While not direct causal evidence regarding the specific incentive(s) which contribute to analysts' relatively pessimistic forecast timing, the fact that the forecast timing effect is stronger in the post-Reg FD period than in the pre- Reg FD period suggests that relatively pessimistic analysts forecast timing is attributable, to some degree, to analysts' incentive to generate trading.

Recall, there are at least three economic incentives that may cause relatively pessimistic analysts to delay their forecasts: (1) the incentive to please management; (2) Hayes' uncertainty argument; and (3) the incentive to generate trading. While it is not possible to undertake a direct test that differentiates between these different potential causes of relatively pessimistic analysts forecast timing, our evidence that the forecast timing effect is more prevalent in the post-Reg FD period suggests that analysts' incentive to please management in order to gain privileged access to selective disclosures from management (see Francis and Philbrick 1993) or generate investment banking business is unlikely to be the primary incentive driving this phenomenon.<sup>1</sup> Issuing pessimistic forecasts later in the forecast period is one way to cooperate with management to win lucrative investment banking business and gain access to management's

---

<sup>1</sup> As a practical issue, because of concerns regarding potential damage to company relationships, brokerage houses may have a more lengthy and cautious internal review process before issuing pessimistic forecasts.



private information. Such cooperation could turn into an “earnings-guidance” game where managers talk down analysts prior to earnings announcements so that the reported earnings numbers can meet or beat the average forecast at the earnings announcement (e.g., Richardson et al. 2004). However, such selective disclosures are prohibited by Reg FD, and the existing research on the effect of Reg FD confirms that Reg FD has been effective in reducing managers' private communication with selected analysts and investors (see, e.g., Gintchel and Markov 2004; Ke et al. 2008). Thus, in the post-Reg FD period analysts have less of an incentive to please management to gain privileged access of private disclosures. Our finding that the forecast timing effect is stronger in the post-Reg FD period therefore suggests that pessimistic analysts' forecast timing is not likely to be primarily attributable to management selectively “talking down” some analysts.

In addition, the post-Reg FD period also encompasses the period after the adoption of the Global Settlement between U.S. regulators and large investment banks and various other regulations designed to reduce analysts' conflicts of interest arising from investment banking.<sup>2</sup> As a result, analysts' incentive to please management in order to generate investment banking business is also likely to be less important in the post-Reg FD period since these new regulations have curtailed analysts' ability to profit by winning investment banking business. On the other hand, the relative importance of analysts' incentive to generate commissions from trading is likely to have increased in the post-FD period. Indeed, one requirement of the Global Settlement

---

<sup>2</sup> The Sarbanes-Oxley Act of 2002 required both the National Association of Securities Dealers (NASD) and the New York Stock Exchange (NYSE) to adopt new rules designed to curtail analysts' conflicts of interest arising from investment banking. The NASD adopted rule 2711 (Research Analysts Research Reports), and the NYSE amended rules 351 (Reporting Requirements) and rule 472 (Communication with the Public). In addition, the Securities and Exchange Commission (SEC) adopted Regulation Analyst Certification (Reg AC). Finally, the Global Settlement between twelve large investment banks and U.S. regulators also imposed additional requirements specifically designed to curtail analysts' conflicts of interest arising from investment banking.

is that the twelve sanctioned investment banks fund research through trading rather than underwriting (Cowen et al. 2006).

Additionally, we find that the forecast timing effect is stronger in firms with a higher percentage of institutional ownership. Since institutional investors are the primary source of stock lending for short selling, firms with a higher percentage of institutional ownership are likely to be easier and cheaper to short sell. As a result, these should be the firms where analysts have a greater incentive to engage in pessimistic forecast timing because relatively pessimistic forecasts made later for these firms are more likely to yield trades. Our results are consistent with this argument: we find that relatively pessimistic analysts' forecast timing is stronger in firms with a higher percentage of institutional ownership. Taken together, our results suggest that analysts' incentive to generate trading is a factor contributing to explaining relatively pessimistic analysts' forecast timing.

Our findings extend two streams of recent research: prior studies that focus on analysts' self-selection (see McNichols and O'Brien 1997; Hayes 1998) and coverage decisions (see Shon and Young 2011), and recent studies that highlight the effect of the incentive to generate trading on analysts' forecasting behavior (see Irvine 2001; Irvine 2004; Jackson 2005; Cowen et al. 2006). Our results make three contributions. First, our results highlight the potential importance of heterogeneity in the forecasting behavior of individual analysts following the same firm: we show that there is a systematic pattern in the timing of individual analysts' forecasts. Second, our results help explain why the number of analysts forecasting earnings tends to increase over the forecast period prior to earnings announcements (see Brown et al. 1985; O'Brien 1988).<sup>3</sup> Third,

---

<sup>3</sup> This is related to the puzzling observation that a very high proportion of analysts' earnings forecasts have short horizons. If analysts' earnings forecasts are an input in analysts' valuation models (e.g., see Bradshaw 2004; Gu and Chen 2004), then intuitively it seems puzzling that the majority of analysts' earnings forecasts are for such short horizons (i.e., less than one year).

our results extend the literature on the decrease in analysts' forecast optimism prior to earnings announcements (e.g., see Richardson et al. 2004; Hutton 2005). Our results suggest that individual analysts timing of their forecasts plays a significant role in driving the over-time decrease in the optimistic bias in analysts' average forecasts (i.e., the "walk-down" in average forecasts prior to earnings announcements).

The paper is organized as follows. Section II sets out our study design and Section III presents the results for our test for pessimistic analysts forecasting timing. Section IV then outlines the results of additional analyses, including an analysis of the association between pessimistic analysts' forecast timing and levels of institutional ownership and future stock returns. The paper concludes with a discussion in Section V.

## **II. Testing for Pessimistic Analysts' Forecast Timing**

### ***2.1 Study Design: Testing for Pessimistic Analysts Forecast Timing***

As the example in Figure 1 shows, if relatively pessimistic analysts delay their forecasts we cannot observe a forecast from these analysts early in the forecast period. Thus, early in the forecast period we cannot observe these analysts' true beliefs. Because of this data limitation, we base our tests around the timing of analysts' first observed forecast.

We use individual analysts' forecast data from the IBES Detail database to examine if *within a firm-year* analysts with a relatively pessimistic view start forecasting later, on average. Our sample period is from 1989 to 2010. For each annual earnings announcement  $t$ , we select a sample of individual forecasts made in the last six months prior to the annual earnings announcement date. If an analyst issues more than one forecast during this six-month period, we retain only her first forecast. Next, we identify whether this forecast was: (a) late in the sense that

it is the analyst's first forecast for that firm-year; or (b) a revision of an early forecast made by the same analyst where we view this analyst as an "early" analyst.

For this sample of individual analysts' forecasts made (or revised) in the last six months prior to an annual earnings announcement, we create a variable,  $LATE$ , that distinguishes between forecasts made by "early" and "late" analysts. Specifically,  $LATE_{ait}$  is a dummy variable equal to one if analyst  $a$ 's forecast of year  $t$  earnings for firm  $i$  made in the last six months prior to the annual earnings announcement is that analyst's first forecast for that firm-year; otherwise,  $LATE$  is coded zero.  $LATE$  is, thus, coded one for those analysts who start forecasting late in a fiscal year and zero otherwise. For each firm-year in our sample, we require at least one analyst who *only* issued forecast(s) in the last six months of the 12-month forecast period (i.e.,  $LATE_{ait} = 1$ ), and at least one analyst who issued forecast(s) both in the first half of the 12-month forecast period and subsequently revised this forecast in the second half of the 12-month forecast period (i.e.,  $LATE_{ait} = 0$ ).

For example, in the example illustrated in Figure 1 there are six analysts forecasting earnings for a firm-year. The oval on the left hand side of Figure 1 illustrates the distribution of the expectation of these six analysts (X1 to X6) for upcoming annual earnings. In this example analysts X1 and X2 are the relatively most optimistic analysts, they both start forecasting earnings more than six months prior the annual earnings announcement date. Then, in the last six months prior to the earnings announcement date, both X1 and X2 update their (earlier) forecast. As a result, these forecasts which are updates of earlier forecast would be coded  $LATE=0$ . On the other hand, in the example illustrated in Figure 1 analysts X3 to X6 all issue their first forecast in the last six months prior to the earnings announcement date; these analysts' forecasts are thus coded  $LATE=1$ .

Using our sample of individual analysts' forecasts issued in the last six months of the forecast period, we create a second variable, *Rank\_Pessimism*, which compares the relative pessimism (optimism) of analysts' forecasts within each firm-year. *Rank\_Pessimism<sub>ait</sub>* is the rank of analyst *a*'s forecast of year *t*'s annual earnings for firm *i*, relative to all other analysts who also forecast year *t* annual earnings for firm *i*. Higher values of *Rank\_Pessimism* indicate an individual analyst's relative pessimism compared to the other analysts. We scale *Rank\_Pessimism* by the number of forecasts for a firm-year, giving a measure of individual analysts' relative pessimism compared to other analysts covering the same firm-year that is scaled between 1 and 0.

Later forecasts — made closer to the earnings announcement date — are more accurate than earlier forecasts (e.g., see Clement 1999). Later forecasts may thus appear to be relatively more pessimistic simply because of new information that made them more accurate. This effect is illustrated in Figure 1 by the fact that forecasts made later in the forecast period are closer to the actual earnings realization. Using the forecast horizon (the number of days between the forecast date and the earnings announcement date) to control for this forecast accuracy effect will mis-specify our test, however. This mis-specification problem is illustrated by the bottom dashed line in Figure 1. This shows that including forecast horizon as a control variable will not only control for the improvement in forecast accuracy through time, but will also extract the effect of increasing relative forecast pessimism through time, i.e., the effect we seek to document.<sup>4</sup> As a result, we include a direct control for forecast accuracy in our tests rather than indirectly controlling for forecast accuracy using forecast horizon: *Rank\_|FError|* is the rank of analysts'

---

<sup>4</sup> In any case, we have found that if you (incorrectly) control for the number of days between the forecast date and the earnings announcement date, then our results still hold.

relative accuracy which is also scaled by the number of forecasts, giving a measure of relative accuracy that is scaled between 0 and 1. We then estimate the following logit model:

$$Prob(Late_{ait} = 1) = \lambda_0 + \lambda_1 Rank\_Pessimism_{ait} + \lambda_2 Rank\_|FError|_{ait} + \varepsilon_{ait}. \quad (1)$$

Our estimate of Equation (1) tests if an analyst's relative pessimism, compared to other analysts forecasting for the same firm-year, is related to whether that analyst initiates forecasting later in the forecast period. If relatively pessimistic analysts — who have higher values of *Rank\_Pessimism*, tend to issue their first forecast later in the forecast period, then higher values of *Rank\_Pessimism* will be positively associated with observations where *LATE* = 1. Thus, we test if  $\lambda_1 > 0$ . Equation (1) controls for analysts' relative accuracy: lower values of *Rank\_|FError|<sub>ait</sub>* indicate relatively more accurate analysts' forecasts.

We also test if this forecast timing effect explains a substantial part of the decrease in the optimism in analysts' average forecasts over the forecast period prior to annual earnings announcements. Using the same forecast data, we estimate two different measures of the decrease in the optimism in analysts' average forecasts between the first and second halves of the 12-month forecast period prior to annual earnings announcement dates. We estimate: (1) the change in the average forecast based upon all available forecasts in both the first and second halves of the 12-month forecast period; and (2) the change in the average forecast based only upon forecasts from the subset of analysts who issue forecasts both in the first and second halves of the 12-month forecast period. Our first measure of the change in average forecasts is a measure of the total decrease in the optimism in average forecasts. Our second measure of the change in average forecasts captures that part of the total decrease in the optimism in average forecasts that is *not* due to relatively pessimistic analysts issuing their forecasts later. The

difference between these two measures thus captures that part of the total decrease in the optimism in average forecasts that is due to relatively pessimistic analysts' forecast timing.

## ***2.2 Sample Selection: Testing for Pessimistic Analysts Forecast Timing***

We use forecasts of one year ahead annual earnings from the IBES Detail database for the sample period 1989 to 2010. Our sample is comprised of all individual forecasts of one year ahead annual earnings made in the last six months prior to annual earnings announcements. We restrict our sample to firm-years where there is at least one analyst who issues her first forecast in the last six months of the 12-month forecast period prior to annual earnings announcements, i.e., at least one analyst with  $LATE_{ait} = 1$ , and at least one analyst who both issues an initial forecast in the first six months of the 12-month forecast period, and subsequently revises this forecast in the last six months of the 12-month forecast period prior to the annual earnings announcement date, i.e., at least one analyst with  $LATE_{ait} = 0$ . This provides a sample of 433,858 analyst-firm-years for 48,955 unique firm-years. These 48,955 firm-years are comprised of observations from 9,512 unique firms and 14,240 different individual analysts.

# **III. Main Results**

## ***3.1 Main Results: Testing for Pessimistic Analysts' Forecast Timing***

The results from estimating Equation (1) are shown in Table 1, which shows that  $\lambda_1$  is significantly positive ( $p < 0.001$ , two-tailed).<sup>5</sup> Higher values of *Rank\_Pessimism* denote analysts with relatively pessimistic views about upcoming annual earnings compared to other analysts who forecast earnings for the same firm-year. The positive coefficient on  $\lambda_1$  confirms that analysts who hold relatively pessimistic views regarding upcoming annual earnings start

---

<sup>5</sup> These results are robust to re-estimating Equation (1) using the Fama-MacBeth (1973) technique.

forecasting later, on average, than other analysts. Our control variable for analysts' relative accuracy,  $Rank\_|FError|$ , is significantly negative ( $p=0.0691$ , two-tailed). Consistent with prior studies (e.g., Brown et al. 1985; O'Brien 1988; Clement 1999), this indicates that later forecasts — that are made closer to the earnings announcement date — are more accurate, on average.

< Insert Figure 2 About Here >

However, if forecasts by “early” and “late” analysts are timed differently within our 6-month sample period, then there could still be an important accuracy difference between early and late analysts which is not appropriately captured by the linear control for forecast accuracy,  $Rank\_|FError|$ .

To test for this possibility, we also undertook a matched pair sample design where we match each “early” (i.e.,  $LATE = 0$ ) forecast with contemporaneous “late” (i.e.,  $LATE = 1$ ) forecasts. This analysis is illustrated in Figure 2. We match each “late” analyst's ( $LATE = 1$ ) forecast with time-matched “early” analyst's ( $LATE = 0$ ) forecast revisions that are made within a five-day window (i.e., date  $t-2$  to date  $t+2$ ). If there is more than one matched “early” forecast, then we use the mean of these forecasts. Of the 118,010 late forecasts, we are able to forecast-horizon-match 38,371 of these forecasts with contemporaneous early forecasts. The median (mean) difference in the forecast horizon, i.e., the number of days between the forecast date and the earnings announcement date, between these horizon-matched forecasts by “early” and “late” analysts is 0 (0.02) days. A comparison of  $Rank\_Pessimism$  across these matched-pairs of “early” and “late” forecasts indicates that the mean forecasts of “late” analysts are significantly more pessimistic than the contemporaneous forecast revisions made by “early” analysts ( $p<0.001$ , two-tailed), i.e., the late analysts are relatively more pessimistic. These results confirm that our results from estimating Equation (1) are due to a change in the beliefs of



individual analysts forecasting over the forecast period and not to an inappropriate control for differences in forecast horizon between early and late analysts.

### ***3.2 Analysis of the Effect of Pessimistic Analysts' Forecast Timing on Average Forecast***

#### ***Optimism***

Next, we examine the contribution of this forecast timing effect to explaining the decrease in average forecast optimism over the forecast period prior to annual earnings announcements. As already discussed, for this analysis we compare two different measures of the increase in average forecast pessimism between the first and the second halves of the 12-month period prior to annual earnings announcements. Specifically, we estimate: (1) the change in the average forecast based upon all available forecasts in both the first and second halves of the 12-month forecast period; and (2) the change in the average forecast based upon only the subset of analysts who issue forecasts both in the first and second halves of the 12-month forecast period. The first measure captures the total change in the average forecast; the second measure captures that part of the total change in the average forecast that is *not* caused by the addition of relatively pessimistic forecasts later in the forecast period. The difference between these two measures captures that part of the change in the average forecast over the 12-month period prior to annual earnings announcements that is attributable to the addition of relatively pessimistic analysts' forecasts later in the forecast period.

< Insert Table 2 Here >

As can be seen in Panel A of Table 2, the typical change in the median forecast based upon all available forecasts (see Column (1)), i.e., the typical change in the average forecast, is -0.1. On the other hand, the typical change in the median forecast based only on the subset of analysts who forecast both early and late in the forecast period is -0.06 (see Column (2)). Thus,

over our entire sample period 40% (-0.04/-0.1) of the typical change in the median forecast is attributable to pessimistic analysts forecast timing. Our analysis, thus, indicates that the forecast timing effect explains a substantial part of the change in average forecasts (and the increase in average forecast pessimism) prior to annual earnings announcements.

We also separately re-estimate the fraction of the decrease in average forecasts that is attributable to forecast timing for both the pre- and post-Reg FD periods. The results are shown in Panels B and C of Table 2. First, as can be seen from the results in Panels B and C of Table 2, it is not unambiguously clear whether Reg FD changed the magnitude of the typical decrease in average forecasts over the forecast period. While the decrease in the median forecast is similar in the pre-Reg FD period as in the post-Reg FD period (-0.1 vs. -0.1), the decrease in the mean forecast is larger in the pre-Reg FD period than in the post-Reg FD period (-0.16 vs. -0.11). Consistent with our expectation, however, there is clear evidence that the composition of the decrease in the average forecast changed between the pre- and post-Reg FD periods. In the pre-Reg FD period, 20% (-0.02/-0.1) of the typical decrease in the average forecast is due to pessimistic analysts' forecast timing; this increases to 50% (-0.05/-0.1) in the post-Reg FD period. Thus, the tendency for analysts with relatively pessimistic views regarding upcoming earnings to start forecasting later than other analysts contributes to explaining a significant proportion of the decrease in average analyst forecasts, especially after Reg FD became effective.

In summary, there are at least three economic incentives that may cause relatively pessimistic analysts to time their forecast to be later, on average, than other analysts: (1) incentive to please management; (2) Hayes' uncertainty argument; and (3) incentives to generate trading. Our analysis using the advent of Reg FD suggests that pessimistic analysts' forecast

timing is not likely to be primarily attributable to management selectively “talking down” some analysts or analysts' incentive to please management to gain access to managerial information or win investment banking business. Analysts' incentive to please management to gain access to information or win investment banking business is likely to have become relatively weaker in the post-Reg FD period. In contrast, analysts' incentive to generate trading is likely to have become relatively more important to analysts in the post-Reg FD period. Our result showing that pessimistic analysts forecast timing has become relatively more important in the post-Reg FD period suggests that analysts' incentive to generate trading contributes at least partially towards explaining this forecast timing effect.

### ***3.3 Sensitivity Analysis Using the Number of Days in the Forecast Horizon***

As a robustness test, instead of using the dummy variable *LATE* that indicates whether an analyst initiates coverage in the first or the second half of the one year period prior to annual earnings announcements, we use the number of days between analysts starting coverage for a firm-year, i.e., coverage initiation, and the earnings announcement. Specifically, we use the following model to test our prediction that relatively pessimistic analysts start forecasting later in the forecast period:

$$\# DAYS_{ait} = \lambda_0 + \lambda_1 Rank\_Pessimism_{ait} + \lambda_2 Rank\_|FError|_{ait} + \varepsilon_{ait} . \quad (2)$$

Where *#DAYS* is the number of days between the issue date of analyst *a*'s first forecast of year *t*'s annual earnings for firm *i* and the eventual earnings announcement date. We would expect that:  $\lambda_1 < 0$ .

< Insert Table 3 Here >

The results from estimating Equation (2) are shown in Table 3, which shows that, consistent with our expectation,  $\lambda_1$  is significantly negative ( $p < 0.001$ , two-tailed).<sup>6</sup> Higher values of *Rank\_Pessimism* denote analysts with relatively pessimistic views about upcoming annual earnings compared to other analysts. Smaller values of *#DAYS* indicate that analysts initiate coverage later in the period. The negative coefficient on  $\lambda_1$  confirms that analysts who hold relatively pessimistic views regarding upcoming annual earnings start forecasting later, on average, than other analysts. Consistent with the results of prior studies (e.g., Brown et al. 1985; O'Brien 1988; Clement 1999), the control variable for analysts' relative accuracy,  $\text{Rank\_}|FError|$ , is significantly positive ( $p < 0.001$ , two-tailed), indicating that forecasts issued later are more accurate, on average.

## IV. Further Analysis

### *4.1 Is Relatively Pessimistic Analysts' Forecast Timing an Analyst-Specific Effect*

Our main results document that relatively pessimistic analysts tend to start issuing their forecasts later on average than other analysts following the same firm-year. A natural question that arises is whether this result is attributable to an analyst-specific effect, i.e., whether the results are driven by a particular type of analyst. We undertake three sets of analyses to shed more light on this question.

First, we examine whether our main results are driven by "new" analysts. That is to say, we test if our results are attributable to firm-years where an analyst is issuing her first ever forecast for a firm. In untabulated analysis, we identified all observations in our sample that are the first forecast by an analyst for a particular firm as forecasts by a new analyst. Then, we

---

<sup>6</sup> These results are robust to re-estimating Equation (2) using the Fama-MacBeth (1973) technique.

dropped these forecasts made by new analysts from our sample and re-ran our analysis. The results from this analysis are consistent with our main results and our inferences are unchanged. This indicated that our main results are not driven by the subset of forecasts issued by “new analysts,” i.e., forecasts that are the first ever forecast by an analyst for a particular firm.

Second, we tested if our results are attributable to persistent relative pessimism of some analysts with respect to some firm, i.e., we examine the persistence of individual analysts' *Rank\_Pessimism* scores for the same firm across years. Specifically, for each analyst-firm we calculated the correlation between analyst *a*'s value of *Rank\_Pessimism* for firm *i* in year *t*, with analyst *a*'s *Rank\_Pessimism* for firm *i* in year *t-1*. Then, to gauge if individual analysts tend to be persistently relatively pessimistic for certain firms, we calculate the average values of these across-time correlations for each analyst-firm pair. The mean (median) value of these correlations is -0.2388 (-0.2355); further, the mean value is significantly less than zero ( $p < 0.001$ ), confirming that, on average, analysts' values of *Rank\_Pessimism* for individual firms tend to be mean-reverting. Thus, our main results do not seem to be driven by persistent pessimism by certain analysts for certain firm, on average.

Third, we tested if our results are attributable to persistent relative pessimism by some analysts for all firms the analyst follows, i.e., whether an individual analyst *a*'s average level of *Rank\_Pessimism* across all firms she forecasts for tends to be correlated across time. If analyst *a* tends to issue relatively pessimistic forecasts for all the firms she follows, then we would expect to observe that the average level of *Rank\_Pessimism* in year *t* for that analyst would be positively correlated with the average level of *Rank\_Pessimism* for that analyst in year *t-1*. We find that, on average, these correlations are negative: the mean (median) correlation is -0.1691 (-0.1592). Again, these correlations are significantly less than zero ( $p < 0.0001$ ), suggesting that analysts

average *Rank\_Pessimism* scores are mean-reverting on average. Taken together, these results suggest that the forecast timing effect we document is not driven by new analysts or a persistent tendency towards relative pessimism on behalf of certain analysts, or a persistent tendency towards relative pessimism in the forecasts certain analysts for certain firms.

#### ***4.2 Exploring one Possible Explanation for Pessimistic Analyst Forecast Timing: Analysts' Incentive to Generate Trading***

To recap, we extend McNichols and O'Brien's (1997) and Hayes' (1998) research regarding analyst self-selection. We argue that analysts with relatively pessimistic views regarding upcoming annual earnings are more reluctant to provide forecasts. That is, prior to earnings announcements, we assume that analysts have a sense of whether their views are relatively pessimistic compared to current market prices or expectations. This is based on the idea that analysts have to continuously update their recommendations based upon a comparison of their own views with current market prices. If this is the case, then we argue that relatively pessimistic analysts will be reluctant to forecast earlier in the forecast period. In this section we explore one of the possible incentives that may contribute to this forecast timing effect: analysts' incentive to generate trading commissions.

Our earlier analysis of the pre- versus post-Reg FD periods suggests that analysts' incentive to generate trading commissions contributes to pessimistic analysts' forecast timing. Hayes (1998) shows analytically that the analyst self-selection effect McNichols and O'Brien (1998) document can, in part, be partially attributed to analysts' incentive to generate trading commissions: because buy recommendations are likely to generate more trading than sell recommendations, analysts' incentive to generate trading commissions helps explain analysts' self-selection in their coverage decisions. We argue that the dropped coverage decisions

documented by McNichols and O'Brien (1997) is an extreme form of analyst censorship; also, in certain cases analysts may not be able to completely drop coverage of certain firms that they have been assigned to cover. Since the analyst forecast timing effect we document, whereby relatively pessimistic analysts are more reluctant to issue forecasts, is also motivated by Hayes' (1998) self-selection argument, a natural question to ask is whether this forecast timing effect is related to analysts' incentive to generate trading commissions. We explore this issue further in this section.

To examine this issue, we assume that both the costs and risks of short selling a stock based on an analyst's earnings forecast decreases as the earnings announcement date approaches — because the window for trading on this earnings news shortens (see D'Avolio 2002; Lamont 2004; Boehmer, Jones, and Zhang 2007).<sup>7</sup> In fact, Diether (2008) examines short-selling contract data from 1999 to 2005 and finds that contracts last on average 38 trading days and the median contract lasts only 11 trading days. *Ceteris paribus*, relatively pessimistic forecasts issued closer to an earnings announcement date can generate more trading commissions because they are more likely to generate short sale transactions.<sup>8</sup> Thus, analysts with a relatively pessimistic view regarding upcoming earnings have a greater incentive to issue their forecasts

---

<sup>7</sup> To short sell a stock, one must first be able to borrow the stock. Financial institutions, such as mutual funds, trusts, or asset managers, provide much of this stock lending for which they receive a daily fee (see D'Avolio 2002; Cohen, Haushalter, and Reed 2004). Stock lenders retain an option to recall the stock at any time. As a result, once a short seller has initiated a short position by borrowing stock, the borrowed stock may be recalled at any time by the lender. If a short position is recalled, then in order to continue to maintain the short position, the short seller needs to find another stock lender. This can be expensive if the new stock lender charges a substantially higher fee. If the short seller is unable to find another lender, he is forced to close his position. This possibility leads to recall risk, one of many risks that short sellers face (see Cohen et al. 2004). These risks decrease as the trading horizon shortens (see D'Avolio 2002). Thus, both the costs and risks of undertaking a short position based on an earnings forecast are likely to decrease as the earnings announcement date approaches.

<sup>8</sup> In Hayes (1998), short sales constraints are one of multiple reasons why optimistic forecasts generate more trading commissions than pessimistic forecasts.

closer to earnings announcement dates because their forecasts are more likely to trigger short sales transactions — and trading commissions — when they are timed in this fashion.<sup>9</sup>

In sum, we expect that the relatively pessimistic analysts' forecast timing effect we document in Section III varies across firms. Specifically, we expect that the pessimistic analysts forecast timing effect will be more evident in firms that are easier to short sell.

Using a proprietary dataset from a leading stock lender, D'Avolio (2002) finds that the cost and risk of short selling are significantly lower for firms with relatively higher institutional ownership (see also Lamont 2004). Institutional holdings have been associated with both analyst following and the incidence of short-sales constraints. Using institutional ownership as a proxy for short-sale constraints, Asquith, Pathak, and Ritter (2005) document that stocks where short-sales are constrained underperform. Firms with larger institutional holdings have more trading so analysts have more capacity to generate trades in such firms. Additionally, with larger institutional holdings, these firms are easier to short-sell. Greater levels of institutional ownership are, thus, likely associated with firms where short-selling constraints are less binding (because of a likely greater supply of stock for lending). As a result, analysts should have a stronger incentive to time relatively pessimistic forecasts in firms with greater levels of institutional ownership as relatively pessimistic forecasts are more likely to generate trades for such firms. Thus, we test if the analyst timing effect we document is more prevalent in firms with larger institutional ownership using the following model:

$$\begin{aligned}
 Prob(Late_{ait} = 1) = & \lambda_0 + \lambda_1 Rank\_Pessimism_{ait} + \lambda_2 Rank\_|FError|_{ait} + \lambda_3 IO_{it} \\
 & + \lambda_4 (Rank\_Pessimism_{ait} * IO_{it}) + \lambda_5 (Rank\_|FError|_{ait} * IO_{it}) + \varepsilon_{ait}, \quad (3)
 \end{aligned}$$

---

<sup>9</sup> Analysts have an incentive to generate trading commissions because trading commission are used to fund sell-side research and brokerage firms tie analysts' compensation, in part, to the trading commission they generate (see Irvine 2001; Cowen et al. 2006). Consistent with the importance of analysts' incentive to generate trading for their brokerage-firm employers, Irvine (2001) provides evidence that analysts' coverage decisions are related to the extent to which analysts can generate trading in a stock.



where  $IO_{it}$  is the percentage of institutional ownership.  $IO_{it}$  is the number of shares held by institutional investors divided by the total number of shares outstanding for firm  $i$  at the beginning of fiscal year  $t$ . Since shares held by institutions are only reported quarterly, we use the institutional ownership data from the beginning of the quarter for all the months in the quarter. Larger values of  $IO_{it}$  denote firms where the costs and risk of short selling are lower. We expect to see a more pronounced forecast timing effect in firms that are easier and cheaper to short sell; these firms are easier and cheaper to short sell because they have a larger level of institutional ownership, on average. Thus, we expect that  $\lambda_4 > 0$ .

< Insert Table 4 Here >

The results from our estimates of Equation (3) are shown in Table 4. As can be seen in Table 4, the results indicate that the forecast timing effect is more evident in firms with higher institutional ownership:  $\lambda_4$  is significantly positive ( $p < 0.001$ , two-tailed).<sup>10</sup> In summary, the results using institutional ownership as a proxy for the marginal costs and risks of short selling stock ( $IO_{it}$ ) are consistent with a greater tendency for relatively pessimistic analysts to time their forecasts to be later in the forecast period for those stocks that are cheaper and less risky to short sell.

### ***4.3 Testing an Implication of Pessimistic Analysts' Forecast Timing***

If analysts who hold relatively pessimistic views about upcoming earnings tend to start forecasting later than other analysts then this suggests that changes in the number of analysts forecasting over the period prior to earnings announcements will be related to firms' subsequent returns, i.e., cumulative abnormal returns after earnings announcement dates. It is now widely accepted that if short selling is costly and there are heterogeneous investor beliefs, a stock can be

---

<sup>10</sup> These results are robust to re-estimating Equation (3) using the Fama-MacBeth (1973) technique.

overvalued and generate low subsequent returns. Desai et al. (2002) report, for example, that the negative abnormal performance of stocks with high short interest persists for up to 12 months. Thus, we can expect that the change in the number of analysts forecasting over the twelve months prior to annual earnings announcements (which captures the overall increase in pessimistic forecasts) will be negatively related to firms' future performance, i.e., cumulative abnormal returns after earnings announcement dates. We test this expectation.

Further, we examine whether any association between the change in the number of analysts forecasting over the period prior to earnings announcements and firms' subsequent returns is attenuated for firms with less short sale constraints. The literature on short sales and stock returns primarily relies on the institutional restrictions governing short sales and on heterogeneous beliefs among investors. With heterogeneous beliefs and no short-sale constraints, pessimistic investors who sell short counterbalance optimistic investors who buy long and they jointly set equilibrium stock prices and, as a consequence, subsequent returns. With short-sale constraints, pessimistic investors are unable to short the stock to the extent they desire, and the equilibrium price will reflect a positive bias and subsequent returns will be low. For any given amount of divergence in expectations, the greater the constraint on short sales, the greater the price and return bias, therefore, the lower the subsequent returns. More divergence in forecasts will be caused by combining the more pessimistic forecasts with the more optimistic forecasts that have already been issued by other analysts (see Figures 1 and 2 for an illustration of this).

Using institutional ownership as a proxy for short-sale constraints, Asquith, Pathak, and Ritter (2005) document that portfolios of stocks with high short interest generally underperform the market and, the lower the level of institutional ownership, the more negative are the

portfolio's abnormal returns. They argue that less constrained stocks (i.e., stocks that are easier to short sell) are more likely owned by institutions since most lendable shares are from institutional owners. Thus, we test whether the effect of increases in the number of forecasts (which increases the amount of pessimism) over the period prior to earnings announcements on firms' subsequent performance is attenuated for firms with higher institutional ownership.

We use average forecasts from the IBES Summary file for the 63,735 firm-years in our sample. The IBES Summary database provides monthly data for the number of analysts with outstanding forecasts. We use this data to measure the change in the number of analysts forecasting over the 12 months prior to annual earnings announcements:  $\Delta Cov_{1-12,it} = Cov_{1,it} - Cov_{12,it}$ , where  $Cov_{1,it}$  and  $Cov_{12,it}$  are the number of forecasts outstanding 1 and 12 months prior to the earnings announcement, respectively. We calculate the cumulative market model adjusted abnormal return ( $CAR_{it}$ ) after the earnings announcement for firm  $i$  in year  $t$  over the 90-trading day interval (1, 90), where day 1 is the day after the earnings announcement. The daily abnormal return for firm  $i$  is computed as the difference between the daily return of firm  $i$  and the value-weighted market return adjusted using the market model. The market model is estimated using a 255 trading-day estimation period ending 60 days before the earnings announcement date. We delete an observation if the stock has fewer than 3 days of return data in the estimation period.  $IO_{it}$  is the percentage of institutional ownership, measured as shares held by institutional investors divided by shares outstanding for firm  $i$  three months before earnings announcement date for year  $t$ . Since institutional ownership is only reported quarterly, we use the institutional ownership data from the beginning of the quarter for all the months in the quarter. We examine if the change in the number of analysts forecasting is related to the cumulative abnormal returns after earnings announcements. More importantly, we use the following regression to test if the

effect of the change in the number of analysts forecasting over the 12 months prior to annual earnings announcements on firms' subsequent performance ( $CAR_{it}$ ) is attenuated for firms with high institutional ownership:

$$CAR_{it} = \alpha_0 + \alpha_1 \Delta Cov_{1-12,it} + \alpha_2 IO_{it} + \alpha_3 IO_{it} * \Delta Cov_{1-12,it} + \varepsilon_{sit} \quad (4)$$

The results from estimating Equation (4) are shown in Table 5.  $\Delta Cov_{1-12,it}$  is significant negatively related to  $CAR_{it}$  ( $p < 0.001$ , two-tailed), indicating that the increased number of analysts forecasting later in the period is associated with lower subsequent abnormal returns, consistent with the idea that analysts' incentive to generate trading from short sales helps explain the pessimistic analysts' forecast timing effect. The significant positive  $\alpha_3$  coefficient ( $p < 0.001$ , two-tailed) indicates that the underperformance associated with larger increases in the number of analysts forecasting over the twelve months prior to annual earnings announcements is attenuated for firms with higher levels of institutional ownership (less constraints for short sales).<sup>11</sup> The results are consistent using the percentage change in the number of analysts forecasting. Results are also robust if we measure  $CAR_{it}$  using alternative windows as (1, 30), (1, 60), (1,180) and (1, 365) and equally-weighted market returns.

< Insert Table 5 Here >

## V. Discussion and Conclusions

We extend McNichols and O'Brien's (1997) and Hayes' (1998) research on analyst self-selection. We argue that dropping coverage may be an extreme form of analyst censorship, and that analysts who hold relatively pessimistic views about future earnings may also choose to forecast later than other analysts. That is, we assume that analysts have a sense of whether their

---

<sup>11</sup> These results are also consistent using the Fama-MacBeth (1973) technique. In addition, we also confirm that these results are robust to estimating Equation (4) using rank regressions, decile rank regressions (see Bradshaw et al. 2001), and to alternative cutoffs for winsorizing.

views are relatively pessimistic compared to current market prices or expectations, and, if analysts hold relatively pessimistic views, they are more likely to choose to forecast earnings later than other analysts.

We first show that individual analysts who hold relatively pessimistic views about future earnings start issuing earnings forecasts later than other analysts forecasting for the same firm-year. Second, consistent with Hayes' (1998) argument, we also show that this forecast timing effect is more prevalent in firms that are cheaper and less risky to short sell. We also show that analysts' forecast timing is stronger in the post-Reg FD period than the pre-Reg FD period. Finally, we show that this analysts' forecast timing effect contributes to explaining the decrease in the average forecast optimism over the forecast period before annual earnings announcements. Our estimates indicate that over our entire sample period (1989 to 2010), 40% of the typical decrease in the average (median) forecast is due to analysts timing of their forecasts; this increases to 50% in the post-Reg FD period.

There are a number of possible reasons why relatively pessimistic analysts may engage in forecast timing. This behavior may be motivated by analysts' incentive to please management or their incentive to generate trading. Since the forecast timing effect is more evident in the post-Reg FD during which managers are prohibited from communicating material private information with select analysts, we conclude that this forecast timing effect is unlikely to be primarily driven by management "talking down" some analysts later in the forecast period.<sup>12</sup> In other words, if we assume that selective disclosures by management were the primary cause of the forecast timing effect we document, then we would expect the forecast timing effect to be weaker in the post-Reg FD period. We find that this is not the case. While we cannot completely rule out the

---

<sup>12</sup> The stronger results for the post-Reg FD period also suggest that analysts' incentive to please management in order to maintain preferential access to managerial information is unlikely to be the primary factor driving this forecast timing effect.

possibility that analysts' incentive to please management contributes in some way to explaining relatively pessimistic analysts' forecast timing, our analysis suggests that analysts' incentive to generate trading contributes to explaining relatively pessimistic analysts' forecast timing.

Our study increases understanding of analysts' forecasting behavior. We argue empirically that such strategic forecast timing behaviors is associated with analysts' incentives to generate trading commissions through short sales. This helps explain why the number of analysts' forecasting tends to increase over the forecast period prior to earnings announcements (see Brown et al. 1985; O'Brien 1988).

Our analyses add to two streams of research. First, our findings extend prior studies that focus on analysts' self-selection (see McNichols and O'Brien 1997; Hayes 1998) and coverage decisions (see Shon and Young 2011). Second, our findings also add to recent studies that highlight the importance of analysts' incentive to generate trading commissions and the potential impact of this incentive on analysts' behavior (see Irvine 2001; Irvine 2004; Jackson 2005; Cowen et al. 2006). Our results extend the findings of these studies by suggesting that analysts' incentive to generate trading commissions is also likely to influence analysts' timing of their forecasts.

## REFERENCES

- Asquith, P., P. Pathak, and J. Ritter (2005), 'Short Interest, Institutional Ownership, and Stock Returns', *Journal of Financial Economics*, Vol. 78, No. 2 (November), pp. 243-276.
- Boehmer, E., C. Jones, and X. Zhang (2008), 'Which Shorts are Informed?', *Journal of Finance*, Vol. 63, No. 2 (April), pp. 491-527.
- Bradshaw, M., (2004), 'How do Analysts Use Their Earnings Forecasts in Generating Stock Recommendations?', *The Accounting Review*, Vol. 79, No. 1 (January), pp. 25-50.
- Bradshaw, M., S. Richardson, and R. Sloan (2001), 'Do Analysts and Auditors Use Information in Accruals?', *Journal of Accounting Research*, Vol. 39, No. 1 (June), pp. 45-74.
- Brown, L. (1997) 'Analyst Forecast Errors: Additional Evidence', *Financial Analyst Journal*, Vol. 53, No. 6, pp. 81-88.
- Brown, P., G. Foster, and E. Noreen (1985) 'Security Analyst Multi-Year Earnings Forecasts and the Capital Market', *Studies in Accounting Research #21* (American Accounting Association).
- Clement, M., (1999), 'Analyst Forecast Accuracy: Do Ability, Resources, and Portfolio Complexity Matter?' *Journal of Accounting and Economics*, Vol. 27, No. 3 (July), pp. 285-303.
- Cohen, J., D. Haushalter, and A. Reed (2004), 'Mechanics of the Equity Lending Market', *Short Selling: Strategies, Risks and Rewards*, Fabozzi, F. (ed.), (Wiley Publishing, New York).
- Cowen, A., B. Groyberg, and P. Healy (2006), 'Which Types of Analyst Firms are More Optimistic?' *Journal of Accounting and Economics*, Vol. 41, Nos. 1-2 (April), pp. 119-146.
- D'Avolio, G. (2002), 'The market for borrowed stock', *Journal of Financial Economics*, Vol. 66, Nos. 2-3 (November-December), pp. 271-306.
- Desai, H., Ramesh, K., Thiagarajan, S., Balachandran, B. (2002), 'An Investigation of the Informational role of Short Interest in the Nasdaq Market', *Journal of Finance*, Vol. 57, No. 5 (October), pp. 2263-2287.
- Diether, K. 2008. Short Selling, Timing, and Profitability. Working paper (Ohio State University).
- Fama, E., and J. MacBeth. (1973), 'Risk, Return, and Equilibrium: Empirical Tests', *Journal of Political Economy*, Vol. 81, No. 3 (May-June), pp. 607-636.

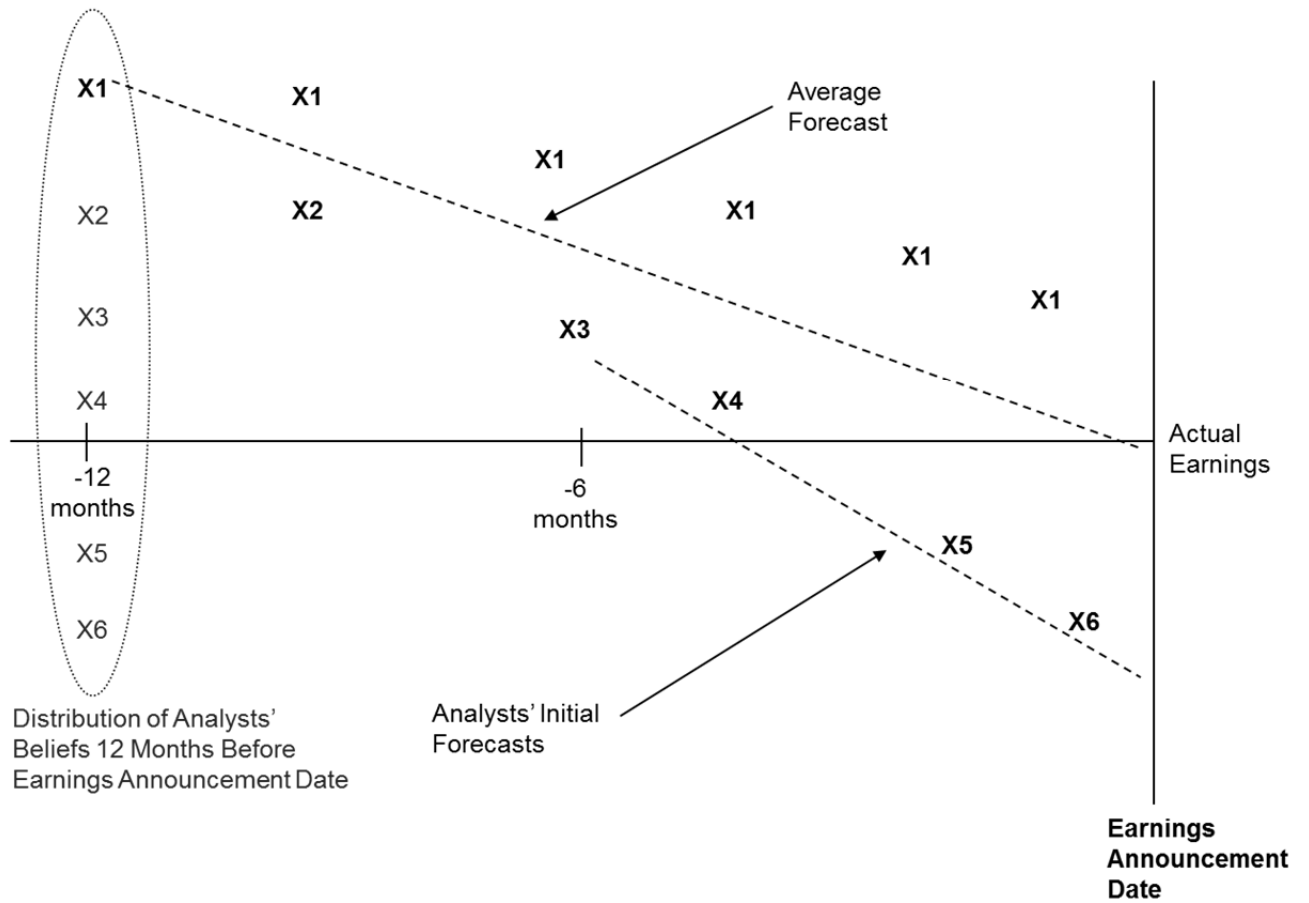
- Francis, J., and D. Philbrick. (1993), 'Analysts' Decisions as a Product of a Multi-Task Environment', *Journal of Accounting Research*, Vol. 31, No. 2 (Autumn), pp. 216-230.
- Gintschel, A. and Markov, S. (2004), 'The Effectiveness of Regulation FD', *Journal of Accounting and Economics*, Vol. 37, No. 3 (September), pp. 293-314.
- Gu, Z., and T. Chen. (2004), 'Analysts' Treatment of Nonrecurring Items in Street Earnings', *Journal of Accounting and Economics*, Vol. 38, Nos. 1-3 (December), pp. 129-170.
- Hayes, R. (1998), 'The Impact of Trading Commission Incentives on Analysts' Stock Coverage Decisions and Earnings Forecasts', *Journal of Accounting Research*, Vol. 36, No. 2 (Autumn), pp. 299-320.
- Hutton, A. (2005), 'Determinants of Managerial Earnings Guidance Prior to Regulation Fair Disclosure and Bias in Analysts' Earnings Forecasts', *Contemporary Accounting Research*, Vol. 22, No. 4 (Winter), pp. 867-914.
- Irvine, P. (2001), 'Do Analysts Generate Trade for their Firms? Evidence from the Toronto Stock Exchange', *Journal of Accounting and Economics*, Vol. 30, No. 2 (October), pp. 209-226.
- Irvine, P. (2004), 'Analysts' Forecasts and Brokerage-Firm Trading', *The Accounting Review*, Vol. 79, No. 1 (January), pp. 125-149.
- Jackson, A. R. (2005), 'Trade Generation, Reputation, and Sell-Side Analysts', *The Journal of Finance*, Vol. 60, No. 2 (March), pp. 719-755.
- Ke, B., K. Petroni and Y. Yu. (2008), 'The Effect of Regulation FD on Transient Institutional Investors' Trading Behavior', *Journal of Accounting Research*, Vol. 46, No. 2 (September), pp. 853-883.
- Lamont, W. (2004), 'Short Sale Constraints and Overpricing', *Short selling: Strategies, Risks and Rewards*, Fabozzi, F. (ed.), (Wiley Publishing, New York).
- McNichols, M., and P. O'Brien. (1997), 'Self-selection and Analyst Coverage', *Journal of Accounting Research*, Vol. 35 (Supplement), pp. 167-199.
- O'Brien, P. (1988), 'Analysts' Forecasts as Earnings Expectations', *Journal of Accounting and Economics*, Vol. 10, No. 1 (January), pp. 53-83.
- Richardson, S., S. Teoh, and P. Wysocki. (2004), 'The Walk-dDwn to Beatable Analyst Forecasts: The Role of Equity Issuance and Insider Trading Incentives', *Contemporary Accounting Research*, Vol. 21, No. 4 (Winter), pp. 885-925.



Shon, J. and S. Young. ( 2011), ‘Determinants of Analysts’ Dropping Coverage Decision: The Role of Analyst Incentives, experience, and Accounting Fundamentals’, *Journal of Business, Finance and Accounting*, Vol. 38, No. 7 (September-October), pp. 861- 886.

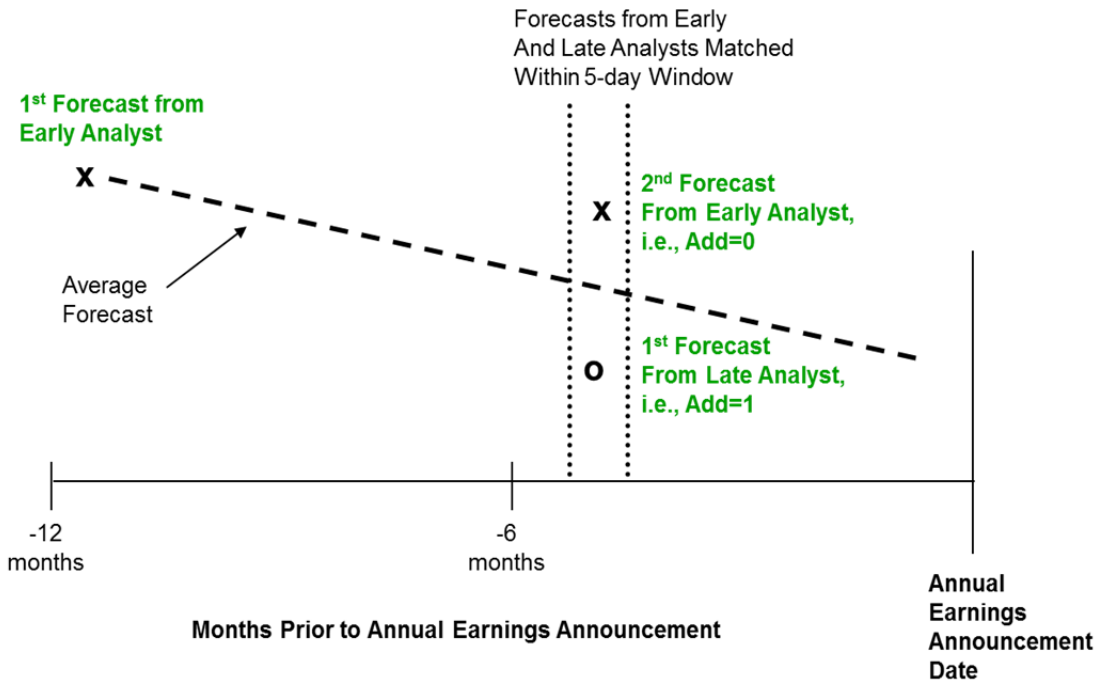
**FIGURE 1**

**An Example to Illustrate Relatively Pessimistic Analysts Forecast Timing**



**FIGURE 2**

**An Alternative Control for Forecast Accuracy:  
Forecast Horizon Matching**



**TABLE 1**  
**Logit Analysis of the Likelihood that Relatively**  
**Pessimistic Analysts Start Forecasting Later**

This analysis tests whether analysts who have a relatively pessimistic outlook for upcoming annual earnings start forecasting later than other analysts. For each annual earnings announcement in this period, we select a sample of individual forecasts that are made in the last six months prior to the earnings announcement date. If an analyst issued more than one annual earnings forecast in this six month sample period, we retain only her first forecast. For this sample of individual analysts' forecasts, we create a variable—*LATE*—that distinguishes between forecasts that are an analyst's first forecast for that firm-year (*LATE*=1), and forecasts that are updates of a previous forecast by the same analyst (*LATE*=0). The latter "updates" are new forecasts that are updates of previous forecasts. Using this sample of individual analysts' forecasts made in the last six months prior to annual earnings announcements, we create a second variable—*Rank\_Pessimism*—to compare the relative pessimism of these individual analysts' forecasts within each firm-year: *Rank\_Pessimism<sub>ait</sub>* is the rank of analyst *a*'s forecast of year *t*'s annual earnings for firm *i*, relative to all other analysts' forecasts of year *t* annual earnings for firm *i*. High values of *Rank\_Pessimism* indicate relative pessimism compared to other analysts forecasting the same annual earnings. We scale *Rank\_Pessimism* by the number of forecasts for a firm-year, giving a measure individual analysts' relative optimism that is scaled between 0 and 1. We use a similar approach to control for analysts' relative accuracy. Using the same individual forecasts made in the last six months prior to the annual earnings announcement date, we calculate *Rank\_|FError|* as the rank of analysts' relative accuracy. *Rank\_|FError|* is also scaled by the number of forecasts, giving a measure of relative accuracy that is scaled between 0 and 1. We use the following logit model to test our prediction that relatively pessimistic analysts start forecasting later in the forecast period:

$$Prob(Late_{ait} = 1) = \lambda_0 + \lambda_1 Rank\_Pessimism_{ait} + \lambda_2 Rank\_|FError|_{ait} + \varepsilon_{ait} . \quad (1)$$

The sample consists of 433,858 analyst-firm-years spread over the period 1989 to 2010, and includes 48,955 firm-years. The results shown below are for a pooled regression that includes all 433,858 analyst-firm-years. These results are robust to using the Fama-MacBeth regression technique.

$\lambda_0$ ( <i>p</i> -value)	$\lambda_1$ ( <i>p</i> -value)	$\lambda_2$ ( <i>p</i> -value)	<i>Likelihood Ratio</i> ( <i>p</i> -value)	<i>No. of Analyst- Firm-Years</i>
-0.5346 (<0.001)	0.1191 (<0.001)	-0.0116 (0.0691)	381.35 (<0.001)	433,858

Higher values of *Rank\_Pessimism<sub>ait</sub>* indicate that analyst *a* is relatively pessimistic compared to the other analysts forecasting year *t* earnings for firm *i*. A significant positive coefficient on  $\lambda_1$  indicates that analysts who start forecasting later (i.e., in the last six months) are relatively pessimistic compared to the other analysts forecasting earnings for the same firm-year. Lower values of *Rank\_|FError|<sub>ait</sub>* indicate relatively more accurate analysts.

**TABLE 2**  
**Does Pessimistic Analysts' Forecast Timing Help Explain the**  
**Decrease in Average Forecasts Prior to Earnings Announcements?**

This table provides univariate statistics for two different measures of the decrease in average forecasts between the first and second 6 months of the 12-month forecast period prior to annual earnings announcements. We estimate two different measures of the decrease in average forecasts between these two 6-month period: (1) the change in the average forecast based upon a sample of all analysts who issue a forecast in either six month period; and (2) the change in the average forecast based upon the sub-sample of only those analysts who issue forecasts in both 6-month periods. Our first measure of the decrease in the average forecast is a measure of the total decrease in average forecasts based upon all available forecasts for that firm-year. Our second measure captures that part of the total decrease in average forecasts that is *not* attributable to pessimistic analysts timing of their forecasts. The difference between these two measures captures that part of the decrease in average forecasts that is attributable to the forecast timing effect where relatively pessimistic analysts issue their forecasts later, on average. Panel A shows the results for our entire sample period which is comprised of 48,955 firm-years for the period 1989 to 2010. Panels B and C show the results for pre- and post Reg FD sub-periods. Panel B shows the results for the pre-Reg FD period (1989 to 1999); Panel C shows the results for the post-Reg FD period (2001 to 2010) periods. Since Reg FD was implemented in 2000, we delete year 2000 in our pre-Reg FD and post-Reg FD analyses. Analyst forecasts are winsorized at the bottom and top 1% of observations.

Measure	<u>Change in Average Forecasts Between</u> <u>the First and Second Halves of the 12-Month</u> <u>Period Prior to Annual Earnings</u> <u>Announcements</u>		
	(1) Change in Average Forecast Based on all Available Forecasts	(2) Change in Average Forecast Based on Only Forecasts from Analysts who Forecast in both 6- Month Periods	(3) Fraction of Decrease in the Average Forecasts Explained by Pessimistic Analysts' Forecast Timing
<i>Panel A: All Sample Years—48,955 Firm-Years over the Period 1989 to 2010</i>			
Δ Median Forecast	-0.1	-0.06	-0.04 (40%)
Δ Mean Forecast	-0.14	-0.09	-0.05 (36%)
<i>Panel B: Pre-Reg FD Sample—24,208 Firm-Years Over the Period 1989 to 1999</i>			
Δ Median Forecast	-0.1	-0.08	-0.02 (20%)
Δ Mean Forecast	-0.16	-0.14	-0.02 (13%)
<i>Panel C: Post-Reg FD Sample—22,513 Firm-Years Over the Period 2001 to 2010</i>			
Δ Median Forecast	-0.1	-0.05	-0.05 (50%)
Δ Mean Forecast	-0.11	-0.06	-0.05 (45%)

**TABLE 3**  
**Sensitivity Analysis of the Likelihood that Relatively**  
**Pessimistic Analysts Start Forecasting Later**

This analysis tests whether analysts who have a relatively pessimistic outlook for upcoming annual earnings start forecasting later than other analysts. For each annual earnings announcement in this period, we select a sample of individual forecasts that are made in the last six months prior to the earnings announcement date. If an analyst issued more than one annual earnings forecast in this six month sample period, we retain only her first forecast. Using this sample of individual analysts' forecasts made in the last six months prior to annual earnings announcements, we create a variable—*Rank\_Pessimism*—to compare the relative pessimism of these individual analysts' forecasts within each firm-year:  $Rank\_Pessimism_{ait}$  is the rank of analyst  $a$ 's forecast of year  $t$ 's annual earnings for firm  $i$ , relative to all other analysts' forecasts of year  $t$  annual earnings for firm  $i$ . High values of *Rank\_Pessimism* indicate relative pessimism compared to other analysts forecasting the same annual earnings. We scale *Rank\_Pessimism* by the number of forecasts for a firm-year, giving a measure individual analysts' relative pessimistic that is scaled between 1 and 0. We use a similar approach to control for analysts' relative accuracy. Using the same individual forecasts made in the last six months prior to the annual earnings announcement date, we calculate  $Rank\_|FError|$  as the rank of analysts' relative accuracy.  $Rank\_|FError|$  is also scaled by the number of forecasts, giving a measure of relative accuracy that is scaled between 0 and 1. We use the following model to test our prediction that relatively pessimistic analysts start forecasting later in the forecast period.

$$\#DAYS_{ait} = \lambda_0 + \lambda_1 Rank\_Pessimism_{ait} + \lambda_2 Rank\_|FError|_{ait} + \varepsilon_{ait} \quad (2)$$

Where:  $\#DAYS$  = the number of days between analyst  $a$ 's first forecast issue date of year  $t$ 's annual earnings for firm  $i$  and the earnings announcement date. If relatively pessimistic analysts systematically start forecasting later, then smaller values of  $\#DAYS$  will be systematically associated with lower values of *Rank\_Pessimism*. So, we would expect that:  $\lambda_1 < 0$

The sample consists of 433,858 analyst-firm-years spread over the period 1989 to 2010, and includes 48,955 firm-years. The results shown below are for a pooled regression that includes all 433,858 analyst-firm-years. These results are robust to using the Fama-MacBeth regression technique.

$\lambda_0$ ( <i>p-value</i> )	$\lambda_1$ ( <i>p-value</i> )	$\lambda_2$ ( <i>p-value</i> )	$R^2$	<i>No. of Analyst-Firm-Years</i>
239.0943 (<0.001)	-8.8651 (<0.001)	11.6711 (<0.001)	0.002	433,858

Higher values of  $Rank\_Pessimism_{ait}$  indicate that analyst  $a$  is relatively pessimistic compared to the other analysts forecasting year  $t$  earnings for firm  $i$ . A significant negative coefficient on  $\lambda_1$  indicates that analysts who start forecasting later (i.e., in the last six months) are relatively pessimistic compared to the other analysts forecasting earnings for the same firm-year. Lower values of  $Rank\_|FError|_{ait}$  indicate relatively more accurate analysts.

**TABLE 4**  
**The Association Between Analysts' Forecast Timing and Institutional Ownership**

We test whether analysts' incentives to engage in forecast timing is related to short selling. D'Avolio (2002) reports that the number of shares available to borrow is highly correlated with institutional ownership, therefore, we assume that the supply of shares available to short is correlated with institutional ownership. If the pessimistic analysts who delay their forecasts are motivated by the trading commissions earned from the short sell, then their engagement in forecast timing should be more pronounced when more shares are available to short. We hypothesize that the pessimistic analysts' engagement in forecast timing is stronger for firms with high institutional ownership. We test this possibility using the following model:

$$\begin{aligned}
 Prob(LATE_{ait} = 1) = & \lambda_0 + \lambda_1 Rank\_Pessimism_{ait} + \lambda_2 Rank\_|FError|_{ait} + \lambda_3 IO_{it} \\
 & + \lambda_4 (Rank\_Pessimism_{ait} * IO_{it}) + \lambda_5 (Rank\_|FError|_{ait} * IO_{it}) + \varepsilon_{ait}
 \end{aligned} \tag{3}$$

$\lambda_0$ ( <i>p-value</i> )	$\lambda_1$ ( <i>p-value</i> )	$\lambda_2$ ( <i>p-value</i> )	$\lambda_3$ ( <i>p-value</i> )	$\lambda_4$ ( <i>p-value</i> )	$\lambda_5$ ( <i>p-value</i> )	<i>Likelihood Ratio</i>
-0.3893 (<0.001)	0.034 (0.016)	0.003 (0.825)	-0.2742 (<0.001)	0.1557 (<0.001)	-0.0281 (0.229)	3833.8 (<0.001)

These results are based upon pooled regressions. The sample consists of 416,769 analyst-firm-years spread over the period 1989 to 2010, and includes 46,095 firm-years. These results are robust using the Fama-MacBeth regression technique. A significant positive coefficient on  $\lambda_4$  indicates that analysts who are relatively pessimistic compared to the other analysts forecasting earnings for the same firm-year start forecasting later (i.e.,  $LATE = 1$ ) when firms have higher percentage of institutional ownership.

Variable Definitions:

$LATE_{ait}$ ,  $Rank\_Pessimism_{ait}$ , and  $Rank\_|FError|_{ait}$  are defined in Table 1;

$IO_{it}$  is the percentage of institutional ownership; shares held by institutions divided by shares outstanding for firm  $i$  at the beginning of fiscal year  $t$ . Since institutional ownership is only reported quarterly, we use the institutional ownership data from the beginning of the quarter for all the months in the quarter.

**TABLE 5**  
**Firm Future Performance and Change in Coverage**

We test if the change in analyst coverage ( $\Delta Cov_{1-12,it}$ ) between 12 months prior to an annual earnings announcement ( $Cov_{12,it}$ ) and 1 month prior to an annual earnings announcement ( $Cov_{1,it}$ ) is related to firms' future performance, i.e., abnormal return.  $\Delta Cov_{12-1,it}$  is winsorized at the bottom and top 1% of observations. Asquith et. al (2005) find that short-sale constrained stocks underperform. Using low institutional ownership as a proxy for short-sale constrain, we argue that the increase in coverage from the relatively pessimistic analysts who forecast late contributes to lower future return for firms with low institutional ownership. The sample period is 1989 to 2010. We estimate the following model:

$$CAR_{it} = \alpha_0 + \alpha_1 \Delta Cov_{1-12,it} + \alpha_2 IO_{it} + \alpha_3 IO_{it} * \Delta Cov_{1-12,it} + \varepsilon_{sit} \quad (4)$$

<i>Method</i>	$\alpha_0$ ( <i>p-value</i> )	$\alpha_1$ ( <i>p-value</i> )	$\alpha_2$ ( <i>p-value</i> )	$\alpha_3$ ( <i>p-value</i> )	$R^2$	<i>Number of Firm-Years</i>
Pooled regression	0.0322 (<0.001)	-0.0294 (<0.001)	-0.0288 (<0.001)	0.0188 (<0.001)	0.0145	63,735
Fama-MacBeth Regression	0.0306 (0.140)	-0.0277 (<0.001)	-0.0395 (0.064)	0.0201 (<0.001)	0.0192	63,735

Variable Definitions:

The raw change in coverage is calculated as:  $\Delta Cov_{1-12,it} = Cov_{1,it} - Cov_{12,it}$ ; where  $Cov_{1,it}$  is the number of analysts with outstanding forecasts 1 month before the earnings announcement date, and  $Cov_{12,it}$  is the number of analysts with outstanding forecasts 12 months before the earnings announcement date. Results are robust using the percentage change in coverage.

$CAR_{it}$  = cumulative abnormal returns after earnings announcement for firm  $i$  in year  $t$  over the 90-trading day interval (1, 90), where 1 is one day after earnings announcement. The daily abnormal return for firm  $i$  is computed as the difference between the daily return of firm  $i$  and the value-weighted market return adjusted using the market model. The market model is estimated using a 255 trading-day estimation period ending 60 days before the earnings announcement date. We delete the observation if the stock has fewer than 3 days of return data in the estimation period. Results are robust if we measure  $CAR_{it}$  using alternative windows as (1, 30), (1, 60), (1,180) and (1, 365) and equally-weighted market returns.

$IO_{it}$  is the percentage of institutional ownership; shares held by institutions divided by shares outstanding for firm  $i$  three months before earnings announcement date for year  $t$ . Since institutional ownership is only reported quarterly, we use the institutional ownership data from the beginning of the quarter for all the months in the quarter.