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Investor Reaction to Celebrity Analysts: The Case of Earnings Forecast Revisions

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ABSTRACT

We examine the effects of analysts' celebrity on investor reaction to earnings forecast revisions. We measure celebrity as the quantity of media coverage analysts receive in sources included in the Dow Jones Interactive database, and find that media coverage is positively related to investor reaction to forecast revisions. The effect of celebrity on the reaction to forecast revisions remains significant after controlling for forecast performance variables examined in prior studies (ex post forecast accuracy, ex ante accuracy, award status, and other variables shown to be related to forecast accuracy). While these results are consistent with the familiarity of the analyst's name affecting the market reaction, we cannot rule out that our measure of celebrity is correlated with error in the performance measures we examine and/or correlated with other unexamined dimensions of forecast performance. A content analysis of a random subsample of the media coverage of our sample analysts suggests that our

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findings likely are not due to the increased availability of forecast revisions. Finally, an investigation of the excess returns around the quarterly earnings announcement date suggests that market participants react too strongly to forecast revisions issued by analysts with high levels of media coverage. Taken together, these findings suggest that an analyst's level of media coverage can affect the initial market reaction to his forecast revisions.

1. *Introduction*

We examine the effect of the celebrity of sell-side analysts on investor reaction to analysts' quarterly earnings forecast revisions. We predict that the celebrity status of analysts will affect investor reaction to forecast revisions not merely because celebrity analysts are superior performers, but also because these analysts' names are more familiar (e.g., Barber and Odean [2005], Barberis and Thaler [2003], Hogarth [1987]). We define a celebrity as a famous or well-publicized person, known for being well-known in addition to his performance-related qualities (Boorstin [1987]). Thus, our study extends prior work in this area that documents cross-sectional variation in the market reaction to forecast revisions conditional on characteristics of the analyst, his employer, and the forecast (e.g., Bonner, Walther, and Young [2003], Clement and Tse [2003], Gleason and Lee [2003], Park and Stice [2000], Stickel [1992]).

We use the quantity of media coverage analysts receive as our empirical proxy for celebrity since prior work has contended that there is a strong association between media coverage and celebrity status (Rojek [2001], Giles [2000], Rein, Kotler, and Stoller [1997], Boorstin [1987]). We measure total media coverage as the number of appearances of an analyst's name associated with his brokerage house employer in all media sources included in the Dow Jones Interactive (DJI) database. Additionally, we measure media coverage in the following individual categories: newspapers and magazines, press wires, *The Wall Street Journal,* and television and radio. We collect these data for a sample of randomly selected sell-side analysts appearing in the Zacks database for the period 1997–1999. We investigate whether the market reacts more strongly to forecast revisions issued by analysts with more media coverage than it does to forecast revisions issued by analysts with less media coverage by relating measures of media coverage to cumulative abnormal returns in the five-day (−2, +2) window surrounding the forecast revisions.

Consistent with our prediction, we find that market participants condition their short-window reaction to the release of forecast revisions on analysts' media coverage. We investigate two alternative explanations for our findings. First, we examine whether our results are due to media coverage proxying for forecast performance. Although we find that media coverage is related to measures of forecast performance used in previous studies (ex post (realized) and ex ante (predicted) forecast accuracy), as well as to measures of perceived performance (membership on the *Institutional Investor*

All-American list and membership on the *Wall Street Journal* All-Star list),[1] the effect of media coverage on the abnormal return surrounding forecast revisions continues to be significant after controlling for these measures. This finding is consistent with our prediction that market participants react more strongly to forecast revisions issued by analysts whose names are more familiar. However, we are unable to rule out that our measure of celebrity (media coverage) is either correlated with the measurement error in our proxies for forecast performance and/or correlated with other unexamined dimensions of forecast performance. Thus, an alternative interpretation of our results is that, consistent with prior work in the area, market participants react more strongly to forecast revisions issued by analysts with superior performance.

The second alternative explanation we investigate is whether media coverage captures the increased availability of the revised forecast. We perform a content analysis of a randomly selected subsample of media coverage for our sample analysts and find that only 4% of the content items examined contained an earnings forecast. This result suggests that the effect of analysts' media coverage likely is not due to the increased availability of forecast revisions. This conclusion is reinforced by our finding of a significant association between the market reaction to the revised forecast and analysts' media coverage in the *prior* year.

Together, these results suggest that analysts' celebrity, as measured by media coverage, can affect the initial market reaction to forecast revisions. To investigate if the initial market reaction to forecast revisions issued by analysts with higher levels of celebrity can be characterized as too strong, we evaluate the returns around the subsequent earnings announcement date. Specifically, we examine the market reaction at the earnings announcement date conditioned on the magnitude of the revision and the analysts' levels of media coverage. We find that, after controlling for unexpected earnings and other determinants of the market reaction at the earnings announcement date, the earnings announcement date returns are significantly associated with analysts' media coverage. This evidence is consistent with the notion that investors do not fully incorporate the information in forecast revisions conditional on media coverage.

Our findings are important for two reasons. First, we contribute to the literature examining the effects of media coverage on stock prices. In recent years, there has been a dramatic increase in the amount of media attention devoted to the stock market and key players therein, such as analysts (Brown and Caylor [2005], Carlson [2000]). Mass media communications can have substantial effects on beliefs and behavior in general (Jamieson and Campbell [2001], DeFleur and Ball-Rokeach [1989], Iyengar and Kinder [1987]), and recent research shows that the media plays an

[1] The award measures are intended to capture multiple facets of analysts' performance including, but not limited to, forecast accuracy.

important role both in the stock price formation process (e.g., Gadarowski [2004], Busse and Green [2002]) and in accounting settings (e.g., Louis, Joe, and Robinson [2005], Miller [2005], Joe [2003], Mutchler, Hopwood, and McKeown [1997], Dopuch, Holthausen, and Leftwich [1986]). This research, however, focuses only on the amount of media coverage for firms, not individual analysts. Given analysts' role in the incorporation of accounting information into stock prices and the media's role in disseminating information issued by analysts, it is important to understand how the media also may serve as an information intermediary for investors. Understanding the media's effects on investor reaction to analysts' forecasts is particularly interesting given the recent scrutiny media outlets have received regarding their coverage of analysts (e.g., Vrana [2002]).

Second, in contrast to the prior literature, our paper is the first to investigate the effects of analysts' media coverage on the market reaction to analyst earnings forecast revisions. The findings are consistent with an analyst's level of media coverage being associated with the market reaction to forecast revisions. The results of our postrevision returns analysis suggest that market participants react too strongly to forecast revisions issued by analysts with high media coverage. If our measures of actual forecast performance and perceived performance serve as reasonable proxies for their respective constructs, our results are consistent with a factor in addition to performance affecting returns. If, however, our measure of media coverage is correlated with error in these performance measures and/or with other unexamined dimensions of forecast performance, then our results indicate simply that an alternative dimension of performance affects returns.

This paper is organized as follows. The next section reviews the literature and outlines our research question. Section 3 describes the sample, variables, and models we use. Section 4 includes the results of our analyses. Section 5 provides conclusions and directions for future research.

2. *Literature Review and Research Question*

Prior research indicates that sell-side analysts' forecast revisions are important to investors forming expectations about firms' earnings and making investment decisions based on these expectations (Hodge [2003], Williams, Moyes, and Park [1996], SRI International [1987]). This conclusion is supported by the significant stock market reaction to the release of forecast revisions (e.g., Brown, Foster, and Noreen [1985], Givoly and Lakonishok [1979], Gonedes, Dopuch, and Penman [1976]). Prior work also has examined the cross-sectional variation in the reaction to forecast revisions with prior forecast accuracy or factors associated with current forecast accuracy (e.g., Fang and Yasuda [2005a,b], He, Mian, and Sangkaraguruswamy [2005], Bonner, Walther, and Young [2003], Clement and Tse [2003], Gleason and Lee [2003], Park and Stice [2000], Michaely and Womack [1999], Stickel [1992]).

In this study, we examine whether analyst "celebrity" affects how investors react to analysts' forecast revisions. Consistent with Boorstin [1987], we view a celebrity as a famous or well-publicized person, who is known for being well-known in addition to his performance-related qualities. Since prior work suggests that factors related to analysts' forecast performance are important determinants of the reaction to forecast revisions and analysts' forecast performance and celebrity status are likely to be related, we incorporate these factors when evaluating the effect of media coverage within our study. We acknowledge, however, that forecast performance is a multidimensional construct and recognize the difficulty both in determining the key dimensions of performance and in creating suitable proxies for each of these dimensions (see, e.g., Gu and Wu [2003]).

We predict that investors will react more strongly to forecast revisions issued by celebrity analysts not only because of their superior performance, but also because of the familiarity of these analysts' names. Prior work in psychology and finance documents that decision makers (including investors) prefer familiar items and situations to unfamiliar ones (Barberis and Thaler [2003], Bornstein [1989]). This preference for familiar items occurs for two (related) reasons. First, people feel more competent and confident when they are working with familiar information and, thus, are more likely to make decisions that favor the sources of familiar information (Fox and Tversky [1995], Heath and Tversky [1991]).[2] Second, individuals relate the familiarity of information to its validity (e.g., Begg, Anas, and Farinacci [1992], Arkes, Hackett, and Boehm [1991], Hasher, Goldstein, and Topping [1977]). In our situation, these findings suggest that investors believe the forecasts of familiar analysts are more valid than equally competent but less publicized analysts. If an analyst's celebrity captures only an analyst's performance, then relying on the celebrity of the analyst would be an efficient heuristic. However, if celebrity also captures familiarity, a heuristic that uses celebrity to determine performance may lead to suboptimal decisions. The evidence in Barber and Odean [2005] supports this view; they find that individual investors tend to buy stocks that catch their attention, and these stocks subsequently underperform.

To investigate the effects of analyst celebrity on the market reaction to forecast revisions, we use the *quantity* of media coverage (the number of appearances of the analyst in media outlets) as our proxy for celebrity. Mass-media representation is widely acknowledged as the primary measure of celebrity (Rojek [2001], Giles [2000], Rein, Kotler, and Stoller [1997], Boorstin [1987]). We argue that media coverage captures the familiarity induced by celebrity because the number, or frequency, of appearance of items or people in the media is encoded in memory relatively automatically (Zacks and Hasher [2002]). Moreover, the accounting and finance literature to date documents that investors rely on media sources for financial information (Gadarowski [2004], Hodge [2003], Busse and Green [2002]).

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[2] For example, investors exhibit "home bias," meaning that they invest far more in stocks based in their home countries or in domestic stocks of companies located geographically close to them than is suggested by optimal portfolio allocation (Barberis and Thaler [2003], Lewis [1999]).

Thus, the degree to which an analyst is included in media sources should proxy for the familiarity of that analyst to investors.

One concern with our use of media coverage as our measure of celebrity is that it also likely captures forecast performance, since media firms face incentives to maintain their reputations as credible news outlets (Kovach and Rosenstiel [2001], Shoemaker and Reese [1996]) in order to maintain their existing reader base and to differentiate themselves from their peers with regard to their reputations for good reporting (Miller [2005], Herman [2002], Jamieson and Campbell [2001], Carlson [2000]). These incentives suggest that reporters may attempt to feature the analysts whom they view as superior performers. The media, however, also has substantial incentives that likely lead them away from focusing solely on analysts' performance. In particular, media firms' primary goal is to capture the attention of as large an audience as possible in order to attract advertising, which is their key source of revenues (Herman [2002], Owen [2002], Jamieson and Campbell [2001], DeFleur and Ball-Rokeach [1989]). In order to attract this attention, they may focus, for example, not solely on the most accurate analysts, but also on those analysts who make the story (and the media outlet) most memorable to news consumers (Mullinathan and Shleifer [2002], Jamieson and Campbell [2001], Kovach and Rosenstiel [2001]).[3] Further, media firms have incentives to feature persons employed by firms that provide them a lot of advertising revenues (in our setting, analysts employed by brokerage firms who advertise in the media outlet) (Jamieson and Campbell [2001]). Empirical evidence suggests these incentives exist with regard to mutual funds (Reuter and Zitzewitz [2006]). Finally, due to the severe pressure that media firms face to fill up space and time (Bernstein [2001], Kovach and Rosenstiel [2001]), they may view an analyst's accessibility as an important feature in addition to his performance.

Thus, while it may be necessary for an analyst to have above-average performance to garner media attention, we assume that the amount of media coverage, our proxy for celebrity, captures an element of the analyst's familiarity along with his performance.[4] We examine empirically the appropriateness of our use of media coverage as the proxy for celebrity, and

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[3] Even if the media attempt to focus on the most accurate analysts, they may have difficulty doing so because reporters work under constraints that make it difficult to predict analysts' forecast accuracy. The first constraint is their lack of training; financial news reporters typically have journalism backgrounds, not business backgrounds (Jamieson and Campbell [2001], Kansas [2001]). Thus, they may lack knowledge of the factors related to analyst forecast accuracy when they start their careers. Further, the conditions reporters face—tight deadlines, overwhelming amounts of information, and the complexities of the financial world—may make it difficult for reporters to learn about these factors on the job (Islam [2002], Jamieson and Campbell [2001]). This constraint exacerbates the cognitive biases that most people who are trying to learn about the performance (here, forecast accuracy) of others are subject to (Tvede [1999], Petty and Wegener [1998]).

[4] This assumption is consistent with the work of Posner [2003], who documents an insignificant association between academic performance (as measured by citations to scholarly publications) and media mentions.

provide evidence in section 4 on the association of media coverage with forecast accuracy and award status. While our empirical evidence suggests that media coverage is a reasonable proxy for the construct of celebrity, a caveat to our interpretation of the results is that media coverage may be capturing a factor other than or in addition to familiarity, which is unobservable due to measurement error in our proxies for performance. In the next section, we discuss the data we use to test the effects of celebrity on investor reaction to forecast revisions.

3. *Data and Descriptive Statistics*

We collected data on media coverage for individual analysts appearing on the Zacks database between 1997 and 1999.[5] We focus on these years because they represent a recent time period in which there was a dramatic increase in the media attention devoted to financial analysts. From the total sample of 3,618 unique analysts on Zacks during this period, we randomly selected 25% (905 analysts). Because Zacks provides only the first initial of analysts' first names, for each of these 905 analysts, we searched for the analysts' full first names in Nelson's Directory of Investment Research. We were unable to find first names for 105 analysts. For the remaining 800 analysts, we hand collected media data from the DJI database (now incorporated into the Factiva database) for every year the analyst appeared on the Zacks database between 1997 and 1999. We searched the DJI database using the analysts' first and last names as well as the brokerage house at which they were employed. We include the brokerage firm name in our search to control for the appearances of other individuals, that is, either other analysts or nonanalysts, with the identical name in the DJI database.[6] To measure analyst media coverage in general, we counted the number of appearances of the analyst's name (in combination with his brokerage firm name) in all the databases DJI incorporates (*ALL_MEDIA*).[7] To gather an analyst's media coverage in specific media outlets, we counted the number of appearances of the analyst's name (in combination with his brokerage firm name) during

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[5] Our time period is representative of the 1985-1999 period in terms of analyst forecast accuracy and bias.

[6] The specific search string we used was "first name near3 last name and brokerage house name." To conduct a more comprehensive search, variations on the first name and brokerage house name also were employed. If the analyst changed brokerage houses during a year, we counted the appearances of the analyst at both brokerage houses and aggregated them to derive one media coverage measure for the analyst for the year. Our results are not sensitive to eliminating analysts who changed brokerage houses during the year.

[7] We did not consider whether articles containing analyst appearances were redundant with other articles as to content, but simply counted the number of appearances. For example, a newswire story about an analyst may be reprinted by five newspapers. This would count as six appearances for the analyst in our analysis. While it may not be appropriate for investors to react to media coverage that is completely redundant with previously reported information, research suggests that they do (e.g., Hugon [2006], Ho and Michaely [1988]).

the year in four types of media outlets: newspapers and magazines (*MAG*), newswires (*WIRE*), *The Wall Street Journal* (*WSJ*), and television and radio (*TV*).[8] We examined *The Wall Street Journal* separately because it is arguably the most influential business media outlet we consider. If the analyst was not mentioned in any media outlet, we coded the corresponding media coverage variable as zero. These searches yielded a sample of 1,726 analyst-year observations.

We matched the data on media coverage for the analyst and year to the quarterly forecast revisions the analyst issued during that year, yielding a sample of 75,807 observations. We match media coverage in year t to quarterly forecast revisions in year t to gain the benefit of timeliness. As a robustness check, we match media coverage in year t–1 to quarterly forecast revisions in year t; all empirical results hold under this alternative approach. A second alternative approach is to use a cumulative measure of media coverage rather than the media coverage for a single year, either t or t–1. Again, we find that all empirical results presented below hold if we use a cumulative media coverage measure, defined as the sum of the media coverage in years t and t–1.

To calculate the forecast revision, we require that the analyst issue both a current and a prior quarterly earnings forecast for the same firm and quarter. We choose the analyst's prior forecast to calculate the forecast revision because it is uniformly more informative to the market than the consensus forecast (Gleason and Lee [2003], Stickel [1991], Imhoff and Lobo [1984]). We further require that the current forecast be issued no more than 90 calendar days prior to the quarterly earnings announcement to ensure that our results are not affected by the inclusion of "stale" forecasts. Finally, we require sufficient data from the Center for Research in Security Prices to calculate the cumulative size-adjusted abnormal return surrounding the release of the forecast revision. All results discussed below hold if we use the firm's excess return over the equal-weighted or value-weighted market index. These criteria yielded our final sample of 20,134 observations.

Panel A of table 1 provides descriptive statistics on this sample. We define the forecast revision (*FCREV*) as the current forecast minus the prior forecast, deflated by price 10 trading days before the release of the revised forecast. Consistent with prior findings (Richardson, Teoh, and Wysocki [2004]), analysts revise their forecasts downward throughout the period. The mean forecast revision (*FCREV*) is slightly negative (−0.0018, two-tailed $p < 0.01$), while the median forecast revision is zero. The five-day (−2, +2)

[8] Newspaper and magazine coverage includes the appearances from searching the DJI database's "all U.S. newspapers" and "major news and business publications." Newswire coverage is the number of appearances obtained when searching the DJI database for "press release wires," "Dow Jones newswires," and "world-wide newswires." *Wall Street Journal* coverage is the number of appearances obtained when searching the *Wall Street Journal* database. Television/radio coverage is the number of appearances obtained when searching the DJI "media transcripts" database.

TABLE 1

Sample Descriptive Statistics: Forecast Revision Sample (N = 20,134)

Panel A: Descriptive statistics

Variable	Mean	Median	First Quartile	Third Quartile	Std. Dev.
FCREV	−0.0018	0.0000	−0.0019	0.0000	0.0095
CAR	−0.0059	−0.0058	−0.0467	0.0368	0.1007
ALL_MEDIA	48.6709	20.0000	7.0000	51.0000	80.6041
MAG	35.2631	15.0000	5.0000	38.0000	58.6318
WIRE	12.0614	4.0000	1.0000	13.0000	23.4644
WSJ	2.1385	1.0000	0.0000	2.0000	4.2453
TV	1.7567	0.0000	0.0000	1.0000	6.6122

Panel B: Pearson and Spearman correlations

	FCREV	*CAR*	*ALL_MEDIA*	*MAG*	*WIRE*	*WSJ*	*TV*
FCREV		0.1141	−0.0016	0.0017	−0.0046	0.0103	−0.0162
		(<0.001)	(0.817)	(0.809)	(0.519)	(0.143)	(0.021)
CAR	0.2067		0.0045	0.0029	0.0063	0.0001	−0.0033
	(<0.001)		(0.527)	(0.682)	(0.372)	(0.986)	(0.642)
ALL_MEDIA	−0.0013	−0.0029		0.9624	0.9118	0.7414	0.6711
	(0.860)	(0.680)		(<0.001)	(<0.001)	(<0.001)	(<0.001)
MAG	−0.0015	−0.0040	0.9722		0.8227	0.7848	0.5234
	(0.835)	(0.568)	(<0.001)		(<0.001)	(<0.001)	(<0.001)
WIRE	0.0028	0.0002	0.8957	0.8800		0.5967	0.7502
	(0.690)	(0.973)	(<0.001)	(<0.001)		(<0.001)	(<0.001)
WSJ	0.0130	−0.0053	0.7205	0.7313	0.6457		0.3293
	(0.066)	(0.453)	(<0.001)	(<0.001)	(<0.001)		(<0.001)
TV	−0.0046	−0.0041	0.5543	0.5082	0.5210	0.4508	
	(0.512)	(0.557)	(<0.001)	(<0.001)	(<0.001)	(<0.001)	

This table provides descriptive statistics for the forecast revision sample. Panel A provides the mean, median, first and third quartiles, and the standard deviation for the forecast revision, market reaction, and media coverage variables. *FCREV* is the price-deflated quarterly earnings forecast revision; *CAR* is the five-day size-adjusted excess return centered on the forecast revision date; *ALL_MEDIA* is the analyst's annual media coverage in all databases DJI incorporates; *MAG* is the analyst's annual media coverage in the newspapers and magazines database on DJI; *WIRE* is the analyst's annual media coverage in the newswires database on DJI; *WSJ* is the analyst's annual media coverage in *The Wall Street Journal* database on DJI; and *TV* is the analyst's annual media coverage in the television and radio database on DJI. Panel B provides the Pearson (Spearman rank) correlation above (below) the diagonal; the two-tailed *p*-value is provided in parentheses below the correlation.

cumulative abnormal return (*CAR*), defined as the firm's compounded return minus the compounded return on the size decile portfolio to which the firm belonged at the beginning of the year, also is slightly negative (mean = −0.59%, median = −0.58%, two-tailed *p*-values < 0.01). The sample analysts have a mean (median) of 48.67 (20) media appearances (*ALL_MEDIA*) during the year; 8.5% of our sample corresponds to forecast revisions issued by analysts with no media coverage (not tabulated). The majority of the media appearances are in newspapers and magazines (*MAG*) (mean = 35.26, median = 15), followed by newswires (*WIRE*) (mean = 12.06, median = 4). By contrast, the frequency of mentions in the *Wall Street Journal* (*WSJ*) (mean = 2.14, median = 1) or on television (*TV*) (mean =1.76, median = 0) is very low.[9]

[9] *ALL_MEDIA* includes appearances in sources that are not captured in our *MAG*, *WIRE*, *WSJ*, or *TV* categories. Thus, the total for *ALL_MEDIA* is not merely the sum of these four components.

Panel B of table 1 provides the correlations between these variables. As expected, there is a positive and significant correlation between *FCREV* and *CAR* (Pearson correlation = 0.1141; Spearman rank correlation = 0.2067, two-tailed p-values $<$ 0.001), indicating that larger forecast revisions are associated with larger market reactions at the revision date. We find little evidence that media coverage is associated with either the level of the forecast revision or the level of the market reaction at the revision date. With the exception of a small, negative association between *FCREV* and *TV* using a Pearson correlation (−0.0162, two-tailed $p = 0.021$) and a small, positive association between *FCREV* and *WSJ* using a Spearman rank correlation (0.0130, two-tailed $p = 0.066$), all correlations between *FCREV* or *CAR* and our media measures are not significant. These univariate correlations, however, do not indicate whether investors react differentially to a given magnitude of forecast revision conditional on the media coverage of the analyst. Finally, and not surprisingly, all of our media measures are positively correlated with each other. For example, the Pearson correlations between *ALL_MEDIA* and the components of media coverage range from 0.6711 to 0.9624 (two-tailed p-values $<$ 0.001).

4. *Empirical Results*

In this section, we first examine if the market reacts more strongly to forecast revisions of a given magnitude if they are issued by celebrity analysts. We then perform several analyses that incorporate empirical proxies of forecasting performance from the prior literature, and also evaluate the increased availability of the revised forecast as an alternative explanation for our findings. Finally, we investigate the returns around the subsequent earnings announcement date conditional on analysts' levels of media coverage.

4.1 INVESTOR REACTION TO FORECAST REVISIONS CONDITIONAL ON MEDIA COVERAGE

To examine whether investors react more strongly to forecast revisions issued by analysts with more media coverage, we estimate the following regression:

$$CAR_{i,j,t} = \alpha_0 + \alpha_1 FCREV_{i,j,t} + \alpha_2 MEDIA_{i,t} + \alpha_3 FCREV_{i,j,t} * MEDIA_{i,t} + \varepsilon_{i,j,t} \quad (1)$$

where:

$CAR_{i,j,t}$ = The five-day size-adjusted excess return for firm j centered on the forecast revision issued by analyst i at time t;[10]

[10] Our conclusions hold if we use as the dependent variable either the three-day size-adjusted return beginning on the forecast revision date (0, +2) or the one-day size-adjusted excess return on the forecast revision date (0, 0).

$FCREV_{i,j,t}$ = The price-deflated quarterly earnings forecast revision issued by analyst i for firm j at time t;
$MEDIA_{i,t}$ = The annual media coverage of analyst i at time t (as measured by *ALL_MEDIA*, *MAG*, *WIRE*, *WSJ*, or *TV*); and
$\varepsilon_{i,j,t}$ = Error term.

Because our sample contains multiple observations for each analyst, we estimate equation (1) using four approaches to control for the resulting cross-sectional dependence (see Gleason and Lee [2003], Keane and Runkle [1998]). The first approach (full sample – A) uses all 20,134 observations in the estimation, but calculates the *t*-statistics using the Huber-White estimator (White [1980], Huber [1967]). In this computation, cross-sectional dependence is addressed by combining individual revisions for the same firm in estimating the variance–covariance matrix. The second approach (full sample – B) also uses all 20,134 observations in the estimation, but uses the generalized method-of-moments (GMM) estimator developed by Keane and Runkle [1998] to adjust the standard errors for both aggregate and firm-specific shocks. The third approach (nonoverlapping sample) uses only those 15,461 observations that have no other forecast revisions for the firm occurring during the return accumulation period, thereby eliminating cross-sectional dependence arising from overlapping return windows. The fourth approach (monthly regressions) estimates equation (1) separately for each month using only one observation for each analyst (see Fama and MacBeth [1973]). We examine the time-series averages of the individual coefficient estimates from the 36 monthly cross-sectional regressions. The *t*-statistics are based on time-series standard errors using the Newey and West [1987] approach to address potential serial correlation in the coefficient estimates from the individual monthly regressions. For brevity, we present the results only for the nonoverlapping sample; the results are unchanged if we use any of the other three approaches to deal with potential cross-sectional dependence. We tabulate the results after observations with studentized residuals greater than two in absolute value or Cook's D greater than one are eliminated; our results are unchanged if outliers are retained.

Table 2 reports the results from estimating equation (1). As expected, the estimated coefficient on *FCREV* is positive and statistically significant regardless of the media measure used, indicating that the market reaction around the release of the revised forecast is associated with the signed magnitude of the forecast revision. Consistent with our predictions, the estimated coefficient on $FCREV * MEDIA$ is positive and statistically significant in all estimations.[11] Untabulated results indicate that this significant coefficient on $FCREV * MEDIA$ holds in the subsamples of positive and negative revisions,

[11] To ensure that observations with a small value of our scaling variable, price, are not driving these results, we re-estimated equation (1) after eliminating 588 observations with a stock price below \$5. The estimated coefficient on $FCREV * MEDIA$ remains significantly positive in this estimation (nonoverlapping sample: 0.0024, $t = 2.60$, two-tailed $p < 0.01$).

TABLE 2

Market Reaction to Forecast Revisions Conditional on Media Coverage: Nonoverlapping Forecast Revision Sample (N = 15,461)

Equation (1):

$$CAR_{i,j,t} = \alpha_0 + \alpha_1\ FCREV_{i,j,t} + \alpha_2\ MEDIA_{i,t} + \alpha_3\ FCREV_{i,j,t} * MEDIA_{i,t} + \varepsilon_{i,j,t}$$

Variable	*ALL_MEDIA*	*MAG*	*WIRE*	*WSJ*	*TV*
Intercept	−0.0020***	−0.0019***	−0.0021***	−0.0020***	−0.0021***
	(−3.14)	(−2.99)	(−3.37)	(−3.28)	(−3.74)
FCREV	0.8598***	0.9347***	0.8818***	0.8476***	0.9798***
	(11.98)	(12.57)	(13.29)	(12.55)	(14.60)
MEDIA	−0.0000	−0.0000	−0.0000	−0.0002	−0.0001
	(−1.14)	(−1.17)	(−1.33)	(−1.30)	(−1.38)
FCREV ∗ *MEDIA*	0.0029***	0.0039***	0.0084***	0.0746***	0.0220***
	(4.03)	(3.46)	(3.90)	(3.99)	(3.32)
R^2	1.76%***	1.87%***	1.85%***	1.75%***	1.79%***
	(86.96)	(92.52)	(91.60)	(86.64)	(88.67)

This table provides results from estimating equation (1) on the forecast revision sample. For each variable included in equation (1), the coefficient estimate is presented; the *t*-statistic is provided in parentheses below the estimated coefficient. The *F*-statistic is provided in parentheses below the R^2. The results are provided for the nonoverlapping sample, which only includes those observations that have no other forecast revisions occurring during the return accumulation window. Variables are defined in the notes to table 1.

*** indicates significance at the 1% level, two-tailed.

indicating that the direction of the revision does not change the significance of the effect of media coverage.

The specification in equation (1) assumes a linear relation between the amount of media coverage and the market reaction. To ensure that this assumption does not drive our results, we perform three sensitivity analyses. First, we replace *MEDIA* in equation (1) with the log of the media coverage of the analyst (*LOG_MEDIA*).[12] In this specification, the coefficient estimate on *FCREV* ∗ *LOG_MEDIA* remains positive and statistically significant (*ALL_MEDIA* in the nonoverlapping sample: 0.1767, $t = 3.46$, two-tailed $p < 0.01$). Second, we replace *ALL_MEDIA* in equation (1) with an indicator variable (*HIGH_MEDIA*) that equals one if the media coverage of the analyst is greater than 20, the median value in our sample, or zero otherwise. In this specification, the coefficient estimate on *FCREV*∗ *HIGH_MEDIA* remains significantly positive (nonoverlapping sample: 0.3634, $t = 2.72$, two-tailed $p < 0.01$).[13]

Third, we regress *CAR* on an intercept and *FCREV* separately for subsamples formed on the overall media coverage of the analyst, and compare the coefficient estimate on *FCREV* across these estimations using seemingly unrelated regression techniques. The five subsamples are: *MEDIA_CAT0*,

[12] Our inferences are unchanged if we estimate equation (1) in ranks or in logs.

[13] The coefficient on *FCREV* ∗ *HIGH_MEDIA* also remains significantly positive at a two-tailed $p < 0.01$ if *HIGH_MEDIA* equals 1 if the analyst's media coverage is greater than 51, the 75th percentile value in our sample, 0 otherwise.

which contains those observations where the analyst's media coverage is 0; *MEDIA_CAT1*, which contains those observations where the analyst's media coverage is between one and nine; *MEDIA_CAT2*, which contains those observations where the analyst's media coverage is between 10 and 25; *MEDIA_CAT3*, which contains those observations where the analyst's media coverage is between 26 and 56; and *MEDIA_CAT4*, which contains those observations where the analyst's media coverage is greater than 56.[14] We reject the null hypothesis that the coefficient estimate on *FCREV* is identical across the five categories at $p < 0.01$ (nonoverlapping sample: $\chi^2(4) = 28.75$). Rather, the coefficient estimate on *FCREV* is lowest in the category of observations where the analyst has zero media coverage (nonoverlapping sample: 0.4057, $t = 3.77$, two-tailed $p < 0.01$), and increases dramatically in the subsample where the analyst has between one and nine media hits (nonoverlapping sample: 1.0004, $t = 7.11$, two-tailed $p < 0.01$).[15] We find no significant differences in the coefficient estimate on *FCREV* between *MEDIA_CAT1* and *MEDIA_CAT2* (nonoverlapping sample: $\chi^2(1) = 0.34$, $p > 0.10$). The coefficient on *FCREV* increases again in the sample of observations where the analyst has between 26 and 56 media hits (nonoverlapping sample: 1.9742, $t = 10.20$, two-tailed $p < 0.01$), but then remains roughly constant in the subsample of observations where the analyst has greater than 56 media hits.[16] These results suggest that the effect of media coverage on the market reaction is nonlinear; specifically, there is a substantial effect from obtaining just a few media mentions and also increasing from a medium level to a high level of media coverage. However, even allowing for potential nonlinearities, we find that the market reacts more strongly to forecast revisions issued by analysts with high media coverage.

4.2 ALTERNATIVE EXPLANATIONS

In this section, we perform several analyses to investigate alternative explanations for our findings. First, we investigate whether our measure of celebrity—media coverage—is associated with actual forecast performance (measured by various forecast accuracy metrics) and perceived performance (measured by the award status of the analyst). We also examine whether the effect of media coverage on investor reaction to forecast revisions holds after controlling for forecast accuracy and award status. Second, we investigate

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[14] We obtained these categories by first placing all observations where the analyst had zero media coverage in the first category, and then dividing the remaining observations into quartiles.

[15] The null hypothesis that the coefficient estimate on *FCREV* is identical across *MEDIA_CAT0* and *MEDIA_CAT1* is rejected at $p < 0.03$ (nonoverlapping sample: $\chi^2(1) = 4.75$).

[16] The null hypothesis that the coefficient estimate on *FCREV* is identical across *MEDIA_CAT2* and *MEDIA_CAT3* is rejected at $p < 0.03$ (nonoverlapping sample: $\chi^2(1) = 10.77$). The null hypothesis that the coefficient estimate on *FCREV* is identical across *MEDIA_CAT3* and *MEDIA_CAT4* cannot be rejected at conventional levels (nonoverlapping sample: $\chi^2(1) = 2.13$, $p > 0.10$).

whether media coverage of the analyst captures increased availability of the forecast revision.

4.2.1. Media Coverage, Forecast Accuracy, and Award Status. In this section, we examine if the relation between media coverage and investor reaction to forecast revisions is due to differential media coverage reflecting differential actual forecast performance (measured by forecast accuracy) or differential perceived performance (measured by inclusion on the *Institutional Investor* All-American list or *Wall Street Journal* All-Star list). We use award status as our proxy for perceived performance because these rankings attempt to capture multiple dimensions of performance, not just forecasting performance. To investigate this issue, we first examine the univariate correlations between either ex post (realized) or ex ante (predicted) accuracy, award status, and the various media coverage variables. We then examine if our findings in table 2 hold if we allow the market reaction to analyst forecast revisions to vary with both forecast accuracy and award status.

Panel A of table 3 provides the univariate correlations between ex post forecast accuracy and the media coverage variables. We measure ex post accuracy as the absolute percentage forecast error (*APE*), defined as the absolute value of actual quarterly earnings minus the quarterly earnings forecast, deflated by price 10 trading days before the release of the revised forecast.[17] Although the Pearson correlations between ex post accuracy and the media coverage variables are not significantly different from zero, the Spearman rank correlations between *APE* and either *MAG* or *WSJ* are significantly negative. These correlations indicate that, as an analyst's ex post accuracy increases (i.e., *APE* decreases), his coverage in magazines and in the *Wall Street Journal* increases.[18]

We also examine the correlation between media coverage and ex ante forecast accuracy. To calculate ex ante accuracy, we first regress the accuracy of the revised forecast on variables shown in prior research to be associated with forecast accuracy (Bonner, Walther, and Young [2003], Clement and Tse [2003], Brown [2001], Clement [1999], Jacob, Lys, and Neale [1999], Mikhail, Walther, and Willis [1997, 1999], Stickel [1992], O'Brien [1988]). Following Clement [1999], we control for time-period and firm effects by mean-adjusting each explanatory variable. Because we require a minimum of two analysts for each firm and quarter to make these adjustments as well as data to calculate the other determinants of forecast accuracy, the sample

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[17] The use of an absolute percentage error metric, instead of a squared percentage error metric, to measure forecast accuracy is supported by Gu and Wu's [2003] finding of a positive association between forecast bias and earnings skewness.

[18] These findings hold if we examine the relation between ex post accuracy and the analyst's media coverage in the prior year. Specifically, the Pearson (Spearman rank) correlation between *APE* and *WSJ* for the prior year is –0.0121, two-tailed $p < 0.10$ (–0.0382, two-tailed $p < 0.01$). The Spearman rank correlation between *APE* and *MAG* for the prior year is –0.0145 (two-tailed $p < 0.05$). All other correlations are not significant. We also find no association between ex post accuracy and media coverage when using the relative accuracy measure employed by Hong, Kubik, and Solomon [2000] (results not tabulated).

TABLE 3

Univariate Correlations between Media Coverage, Forecast Accuracy, and Award Status

	ALL_MEDIA	*MAG*	*WIRE*	*WSJ*	*TV*
Panel A: Correlations of media coverage with ex post accuracy ($N = 20{,}134$)					
Pearson	0.0030	−0.0049	0.0016	−0.0035	0.0115
	(0.676)	(0.487)	(0.821)	(0.616)	(0.103)
Spearman rank	−0.0107	−0.0177	−0.0050	−0.0157	−0.0001
	(0.128)	(0.012)	(0.474)	(0.026)	(0.993)
Panel B: Correlations of media coverage with ex ante accuracy ($N = 7{,}474$)					
Pearson	−0.0055	−0.0183	−0.0013	−0.0433	0.0029
	(0.670)	(0.159)	(0.921)	(<0.001)	(0.821)
Spearman rank	−0.0165	−0.0235	−0.0229	−0.0428	−0.0042
	(0.203)	(0.070)	(0.078)	(0.001)	(0.748)
Panel C: Correlations of media coverage with *Institutional Investor* All-American list ($N = 20{,}134$)					
Pearson	0.3384	0.3209	0.3277	0.3961	0.2091
	(<0.001)	(<0.001)	(<0.001)	(<0.001)	(<0.001)
Spearman rank	0.2477	0.2331	0.2564	0.2995	0.1878
	(<0.001)	(<0.001)	(<0.001)	(<0.001)	(<0.001)
Panel D: Correlations of media coverage with *Wall Street Journal* All-Star list ($N = 20{,}134$)					
Pearson	0.1744	0.1468	0.1546	0.2093	0.1666
	(<0.001)	(<0.001)	(<0.001)	(<0.001)	(<0.001)
Spearman rank	0.1445	0.1527	0.1369	0.2511	0.1172
	(<0.001)	(<0.001)	(<0.001)	(<0.001)	(<0.001)

This table provides Pearson and Spearman rank correlations between media coverage and ex post accuracy (*APE*, panel A), ex ante accuracy (*PRED_ACC*, panel B), membership on the *Institutional Investor* All-American list (*IIAWARD*, panel C), and membership on the *Wall Street Journal* All-Star list (*WSJAWARD*, panel D). In all panels, the two-tailed *p*-value is provided in parentheses below the correlation. *APE* is defined as the absolute value of actual quarterly earnings minus the quarterly earnings forecast, deflated by share price 10 trading days before the release of the revised forecast. *PRED_ACC* is defined as the fitted value from equation (2). *IIAWARD* is an indicator variable that equals 1 if the analyst was named to the *Institutional Investor* All-American list, 0 otherwise. *WSJAWARD* is an indicator variable that equals 1 if the analyst was named to the *Wall Street Journal* All-Star list, 0 otherwise; all other variables are defined in the notes to table 1.

used in these tests is reduced to 7,474 observations. Descriptive statistics on this reduced sample (not tabulated) are consistent with those found in prior research (e.g., Bonner, Walther, and Young [2003], Clement and Tse [2003]).

We define ex ante accuracy as the predicted value of accuracy from estimating the following regression after eliminating observations with studentized residuals greater than two in absolute value or Cook's D greater than one:

$$\begin{aligned} PRMAPE_{i,j,q} = \beta_0 &+ \beta_1 PRMAPE_{i,j,q-1} + \beta_2 RFCAGE_{i,j,q} + \beta_3 RFIRMEXP_{i,j,q} \\ &+ \beta_4 RGENEXP_{i,j,q} + \beta_5 RTURNOVER_{i,j,q} + \beta_6 RFCFREQ_{i,j,q} \\ &+ \beta_7 RNOFIRM_{i,j,q} + \beta_8 RNOIND_{i,j,q} + \beta_9 RBROKSIZE_{i,j,q} \\ &+ \beta_{10} RIIAWARD_{i,j,q} + \xi_{i,j,q} \end{aligned} \quad (2)$$

where:

$PRMAPE_{i,j,q}$ = The absolute value of analyst i's forecast error for firm j and quarter q minus the mean absolute forecast error for firm j and quarter q, deflated by the mean absolute forecast error for firm j and quarter q;[19]

$RFCAGE_{i,j,q}$ = The number of calendar days between the issuance of analyst i's forecast for firm j and quarter q minus the mean forecast age for firm j and quarter q;

$RFIRMEXP_{i,j,q}$ = The firm experience of analyst i for firm j at quarter q minus the mean firm experience for firm j and quarter q;

$RGENEXP_{i,j,q}$ = The general experience of analyst i at quarter q minus the mean general experience for firm j and quarter q;

$RTURNOVER_{i,j,q}$ = An indicator variable that equals 1 if analyst i changed brokerage houses during the year in which quarter q falls (0 otherwise) minus the mean of the indicator variable for firm j and quarter q;

$RFCFREQ_{i,j,q}$ = The number of forecasts issued by analyst i for firm j and quarter q minus the mean forecast frequency for firm j and quarter q;

$RNOFIRM_{i,j,q}$ = The number of firms followed by analyst i at quarter q minus the mean number of firms followed for firm j and quarter q;

$RNOIND_{i,j,q}$ = The number of industries followed by analyst i at quarter q minus the mean number of industries followed for firm j and quarter q;

$RBROKSIZE_{i,j,q}$ = The size of the brokerage house at which analyst i is employed at quarter q minus the mean brokerage house size for firm j and quarter q;

$RIIAWARD_{i,j,q}$ = An indicator variable that equals 1 if analyst i was named to the *Institutional Investor* All-American list during the year prior to that in which quarter q falls (0 otherwise) minus the mean of the indicator variable for firm j and quarter q; and

$\xi_{i,j,q}$ = Error term.

Inferences from estimating equation (2) are unchanged if outliers are retained in the estimation and are consistent with prior work (results not tabulated).[20]

[19] When the analyst issues more than one forecast for the firm and quarter, we calculate *PRMAPE* using the one issued closest in time to the quarterly earnings announcement date.

[20] Specifically, the estimated coefficients on *RFCAGE*, $PRMAPE_{q-1}$, *RTURNOVER*, *RNOFIRM*, and *RNOIND* are statistically positive, indicating that forecasts that are issued earlier

As shown in panel B of table 3, the Pearson correlation between ex ante (predicted) accuracy and *WSJ* is significantly negative (−0.0433, two-tailed $p < 0.001$); thus, as an analyst's predicted accuracy increases, his coverage in the *Wall Street Journal* increases. The Pearson correlations between ex ante accuracy and the remaining media variables are insignificantly different from zero. The Spearman rank correlations between ex ante accuracy and *MAG*, *WIRE*, and *WSJ* are significantly negative, indicating that an analyst's coverage in magazines, press wires, and the *Wall Street Journal* is increasing in his forecast accuracy. The Spearman rank correlations between ex ante accuracy and *ALL_MEDIA* and *TV* are insignificant. To examine the association between media coverage and forecast accuracy in a multivariate setting, we estimate equation (2) including *RALL_MEDIA*, defined as the media coverage of analyst i at quarter q minus the mean analyst media coverage for firm j and quarter q. In this specification, the coefficient on *RALL_MEDIA* is insignificant (0.0001, $t = 0.66$, two-tailed $p > 0.10$).[21] Overall, these results suggest a limited association between celebrity, as measured by media coverage, and actual performance, as measured by forecast accuracy. However, an important caveat to these results is that our empirical proxies of celebrity and performance may contain sufficient measurement error resulting in attenuation of their association.

Panels C and D of table 3 provide the univariate correlations between perceived performance (award status) and media coverage. We find that all measures of media coverage are positively associated with inclusion on the *Institutional Investor* All-American list (*IIAWARD*, panel C) and the *Wall Street Journal* All-Star list (*WSJAWARD*, panel D). The correlations, ranging from 0.1172 to 0.3961, indicate that award-winning analysts have more media coverage. These significant correlations, coupled with the correlations between media coverage and ex post or ex ante accuracy, suggest that media coverage captures dimensions of an analyst's performance in addition to an analyst's familiarity. The magnitude of the correlations is higher with *IIAWARD* than with *WSJAWARD*; this finding may reflect reporters' beliefs that the *Institutional Investor* All-American award is a better measure of performance. Consistent with prior research, we find that award status and forecast accuracy are correlated, indicating that award-winning analysts are more accurate (results not tabulated).

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are less accurate, as are forecasts issued by analysts who are inaccurate in the prior quarter, change brokerage houses, and follow more firms and more industries. The coefficients on *RFCFREQ* and *RIIAWARD* are statistically negative, indicating analysts who forecast more frequently or who are named to the *Institutional Investor* All-American list are more accurate. The coefficient on *RBROKSIZE* is insignificant, as are the coefficients on *RFIRMEXP* and *RGENEXP*. However, if we omit *RFIRMEXP* from the regression, the coefficient on *RGENEXP* becomes significantly negative at one-tailed $p < 0.10$, suggesting that analysts with more general experience issue more accurate forecasts.

[21] The coefficient remains insignificant if we replace *RALL_MEDIA* with the other measures of media coverage (*RMAG*, *RWIRE*, *RWSJ*, or *RTV*).

Overall, these findings indicate an association between our measure of celebrity (media coverage) and dimensions of forecast performance (accuracy) and award status (perceived performance). This finding raises questions as to whether media coverage captures performance and other related factors in addition to or instead of celebrity. To investigate this issue, we analyze the determinants of analysts' media coverage. We regress the relative media coverage for the analyst on forecast accuracy and award status, as well as variables that capture the extent to which an analyst makes himself "visible," and therefore becomes more likely to be covered by the media. These variables include relative measures of brokerage house size, number of industries followed, number of firms followed, general experience, prior forecast accuracy, forecast frequency, award status, forecast boldness, forecast bias, and affiliation status.

In untabulated analyses, we find that the relative measures of number of firms followed, forecast frequency, and relative forecast bias are positively associated with the analyst's relative media coverage, while relative forecast boldness (defined as the deviation from the consensus forecast) and general experience are negatively associated with the analyst's relative media coverage. The other independent variables, including relative forecast accuracy and relative award status, are not significantly related to media coverage (two-tailed $p > 0.10$) after controlling for relative media coverage in the prior year. Additionally, when relative media coverage in the prior year is added to the regression, the coefficient estimate on relative award status becomes insignificant, while the coefficient estimate on relative general experience becomes significantly negative and the coefficient estimate on relative forecast bias becomes significantly positive. These findings suggest that analysts who have relatively less general experience and who are relatively more pessimistic have more media coverage. Relative media coverage in the prior year is the most important determinant of media coverage in the current year; when prior year media coverage is added to the regression model, the explanatory power increases from 12.72% to 78.93%. Overall, these results suggest that analysts who are more "visible," as captured by the number of firms followed, how frequently they forecast, and their prior media coverage, have higher levels of current media coverage. Although analysts whose relative boldness is lower have higher levels of media coverage, our content analysis (described in section 4.2.2) suggests that this finding may be due to reporters seeking documentation of a mainstream opinion regarding a company's financial prospects.

As an additional test of whether our finding of a greater reaction to forecast revisions issued by analysts with more media coverage is due to the association of media coverage with forecast accuracy or award status, we modify equation (1) to include these variables. As above, we measure forecast accuracy as either ex post accuracy (defined as the absolute percentage error of the revised forecast) or ex ante accuracy (defined as the predicted value from equation (2)). We measure award status as either inclusion on the *Institutional Investor* All-American list or inclusion on the *Wall Street Journal*

All-Star list. In this multivariate analysis, we also control for firm media coverage by including firm size in the model, based on Gadarowski's [2004] finding that firm size is strongly related to firm media coverage. We measure firm size as the log of total assets for the firm; our results are unchanged if we measure firm size as the log of the market value of common equity. We estimate versions of the following regression using the four approaches outlined in section 4.1 to address cross-sectional dependence:

$$\begin{aligned} CAR_{i,j,t} = {} & \delta_0 + \delta_1 FCREV_{i,j,t} + \delta_2 ACCURACY_{i,j,t} + \delta_3 FCREV_{i,j,t} * ACCURACY_{i,j,t} \\ & + \delta_4 AWARD_{i,t} + \delta_5 FCREV_{i,j,t} * AWARD_{i,t} + \delta_6 FIRM\_SIZE_{j,t} \\ & + \delta_7 FCREV_{i,j,t} * FIRM\_SIZE_{j,t} + \delta_8 ALL\_MEDIA_{i,j,t} \\ & + \delta_9 FCREV_{i,j,t} * ALL\_MEDIA_{i,j,t} + \sum_{k=1}^{47} \delta_{9+k} IND_{j,t} + \zeta_{i,j,t} \end{aligned} \tag{3}$$

where the variables not previously defined are:

$ACCURACY_{i,j,t}$ = The ex post or ex ante accuracy of the quarterly earnings forecast revision issued by analyst i for firm j at time t;
$AWARD_{i,t}$ = An indicator variable that equals 1 if analyst i is included on the *Institutional Investor* All-American list or the *Wall Street Journal* All-Star list at time t, 0 otherwise;
$FIRM\_SIZE_{j,t}$ = The log of total assets for firm j at time t;
$IND_{j,t}$ = An indicator variable representing the industry sector for firm j at time t; and
$\zeta_{i,j,t}$ = Error term.

Table 4 provides the results from estimating equation (3). As in our estimation of equation (1), we eliminate observations with studentized residuals greater than two in absolute value or Cook's D greater than one in the tabulated results; our inferences are unchanged if outliers are retained. We present the results only for the nonoverlapping sample using the five-day size-adjusted excess return, *ALL_MEDIA*, and ex post accuracy; the results are unchanged if we use one of the other approaches to deal with potential cross-sectional dependence, replace ex post accuracy with ex ante (predicted) accuracy or Hong, Kubik, and Solomon's [2000] relative accuracy measure, replace *ALL_MEDIA* with *MAG*, *WIRE*, *WSJ*, or *TV*, replace the dependent variable with the size-adjusted return over (0, +2) or on day 0, and/or estimate equation (3) in ranks or in logs.

Consistent with our findings in table 2, the estimated coefficient on *FCREV* is positive and statistically significant in all specifications. We also find that the estimated coefficient on $FCREV*ACCURACY$ is significantly negative, indicating that market participants react less to revised forecasts that are ex post less accurate. This result is consistent with Clement and Tse [2003] and Bonner, Walther, and Young [2003], who find that the market reaction to forecast revisions is associated with forecast accuracy. Further, the

TABLE 4

Market Reaction to Forecast Revisions Conditional on Media Coverage, Ex Post Accuracy, and Award Status: Nonoverlapping Forecast Revision Sample (N = 15,461)

Equation (3):

$$CAR_{i,j,t} = \delta_0 + \delta_1 FCREV_{i,j,t} + \delta_2 ACCURACY_{i,j,t} + \delta_3 FCREV_{i,j,t} * ACCURACY_{i,j,t} + \delta_4 AWARD_{i,t} + \delta_5 FCREV_{i,j,t} * AWARD_{i,t} + \delta_6 FIRM\_SIZE_{j,t} + \delta_7 FCREV_{i,j,t} * FIRM\_SIZE_{j,t} + \delta_8 ALL\_MEDIA_{i,j,t} + \delta_9 FCREV_{i,j,t} * ALL\_MEDIA_{i,j,t} + \sum_{k=1}^{47} \delta_{9+k} IND_{j,t} + \zeta_{i,j,t}$$

	AWARD = *IIAWARD*		*AWARD* = *WSJAWARD*	
Intercept	−0.0016**	−0.0120	−0.0014**	−0.0116
	(−2.47)	(−0.93)	(−2.07)	(−0.90)
FCREV	0.9334***	0.3685**	0.9124***	0.3185**
	(12.64)	(2.33)	(11.65)	(1.96)
ACCURACY	−0.0476***	−0.0495***	−0.0476***	−0.0502***
	(−4.88)	(−4.59)	(−4.88)	(−4.65)
FCREV∗ ACCURACY	−0.4870***	−0.4690***	−0.4800***	−0.4588***
	(−6.50)	(−5.72)	(−6.37)	(−5.58)
AWARD	0.0016	0.0012	−0.0011	−0.0006
	(0.97)	(0.69)	(−0.69)	(−0.36)
FCREV∗ AWARD	0.1284	0.1088	0.5103**	0.6327***
	(0.48)	(0.40)	(2.51)	(3.08)
FIRM_SIZE		0.0015***		0.0015***
		(4.73)		(4.83)
FCREV∗ FIRM_SIZE		0.0895***		0.0889***
		(2.89)		(2.87)
ALL_MEDIA	−0.0000	−0.0000	−0.0000	−0.0000
	(−1.25)	(−0.96)	(−1.08)	(−0.87)
FCREV∗ ALL_MEDIA	0.0024***	0.0031***	0.0020***	0.0026***
	(2.86)	(3.53)	(2.65)	(3.20)
Industry Effects	No	Yes	No	Yes
R^2	2.09%***	2.68%***	2.10%***	2.67%***
	(44.52)	(6.83)	(44.70)	(6.81)

This table provides results from estimating equation (3) on the forecast revision sample. For each variable included in equation (3), the coefficient estimate is presented; the *t*-statistic is provided in parentheses below the estimated coefficient. The *F*-statistic is provided in parentheses below the R^2. The results are provided for the nonoverlapping sample, which only includes those observations that have no other forecast revisions occurring during the return accumulation window. *ACCURACY* is the ex post (*APE*) accuracy of the revised forecast; *AWARD* indicates whether the analyst is named to the *Institutional Investor* All-American list (*IIAWARD*) or the *Wall Street Journal* All-Star list (*WSJAWARD*); *FIRM_SIZE* is the log of total assets for the firm; *IND* represents a set of industry control variables to indicate to which Fama−French industry sector the firm belongs; all other variables are defined in the notes to tables 1 and 3.

** and *** indicate significance at the 5% and 1% levels, respectively, two-tailed.

estimated coefficient on *FCREV∗ AWARD* is significantly positive for *WSJAWARD*, suggesting that market participants react more to forecast revisions issued by analysts who are historically more accurate. This finding is consistent with Brown [2001], who documents that prior accuracy is the single best predictor of future forecast accuracy. In untabulated analyses, we find that the coefficient on *FCREV* ∗ *IIAWARD* is significantly positive when accuracy is excluded from the regression. This finding is consistent with Stickel [1992], who documents that analysts named to the *Institutional Investor* All-American list have a greater impact on security prices, and Jackson [2005], who finds that award-winning analysts generate greater trade volume. The insignificant coefficient on *FCREV* ∗ *IIAWARD* in table 4 suggests that the effect of *IIAWARD* on the market reaction to forecast revisions is a result of All-American analysts issuing more accurate forecasts (see Stickel [1992]).

Despite the significant effect of forecast accuracy and award status on the market reaction, we continue to find that the estimated coefficient on *FCREV* ∗ *ALL_MEDIA* is positive and significant at $p < 0.01$.[22] The magnitude of the coefficient estimate on *FCREV* ∗ *ALL_MEDIA* in table 4 (0.0024 using IIAWARD; 0.0020 using *WSJAWARD*) is similar to that in table 2 (0.0029 for *ALL_MEDIA*). The coefficient estimate on *FCREV* ∗ *ALL_MEDIA* remains significantly positive if both *IIAWARD* and *WSJAWARD* (and their interactions with *FCREV*) are included in the model (0.0018, $t = 2.12$, two-tailed $p < 0.04$ for the nonoverlapping sample; results not presented). The coefficient estimate on *FCREV* ∗ *ALL_MEDIA* remains significantly positive when we include industry control variables and firm size in equation (3) to control for firm media coverage (see Gadarowski [2004]), and/or include the number of firms the analyst follows and its interaction with *FCREV* in equation (3). Untabulated results indicate that we continue to find a significant effect of media coverage when we replace *ALL_MEDIA* with an indicator variable, *IND_MEDIA*, that equals one if the analyst has one or more media mentions in the year, zero otherwise (*IIAWARD*: 0.4621, two-tailed $p < 0.01$; *WSJAWARD*: 0.3678, two-tailed $p < 0.05$).[23] Using the results for *IIAWARD*, the magnitude of the excess returns to forecast revisions increases by approximately 194%, from −0.16% to −0.47%, when *IND_MEDIA* changes from zero to one, suggesting that the effect of media coverage on the market reaction to forecast revisions is not only statistically significant, but also economically significant.[24]

As a final approach to control for accuracy, we allow the coefficient estimate on *FCREV* to vary with the determinants of analyst accuracy identified by prior research:

[22] As a sensitivity test of whether measurement error in our forecast accuracy metric is driving these results, we re-estimate equation (3) after eliminating 1,169 observations for which the price-deflated special items amount was greater than four standard deviations from the average price-deflated analyst forecast error (see Keane and Runkle [1998]). The estimated coefficient on *FCREV* ∗ *MEDIA* remains significantly positive in this estimation.

[23] In this specification, the coefficient estimate on *FCREV** *IIAWARD* is significantly positive (0.5808, two-tailed $p < 0.05$), in contrast to the results presented in table 4.

[24] To compute these excess return results, we use the coefficient estimates from the regression using *IIAWARD* in conjunction with the mean values for *FCREV*, *ACCURACY*, and *AWARD*.

$$CAR_{i,j,t} = \phi_0 + \phi_1 FCREV_{i,j,t} + \phi_2 FCREV_{i,j,t} * PRMAPE_{i,j,q-1} + \phi_3 FCREV_{i,j,t} * RFCAGE_{i,j,q} + \phi_4 FCREV_{i,j,t} * RFIRMEXP_{i,j,q} + \phi_5 FCREV_{i,j,t} * RGENEXP_{i,j,q} + \phi_6 FCREV_{i,j,t} * RTURNOVER_{i,j,q} + \phi_7 FCREV_{i,j,t} * RFCFREQ_{i,j,q} + \phi_8 FCREV_{i,j,t} * RNOFIRM_{i,j,q} + \phi_9 FCREV_{i,j,t} * RNOIND_{i,j,q} + \phi_{10} FCREV_{i,j,t} * RBROKSIZE_{i,j,q} + \phi_{11} FCREV_{i,j,t} * RIIAWARD_{i,j,q} + \phi_{12} ALL\_MEDIA_{i,j,t} + \phi_{13} FCREV_{i,j,t} * ALL\_MEDIA_{i,j,t} + \upsilon_{i,j,q} \quad (4)$$

where the variable not previously defined is: $\upsilon_{i,j,t}$ = Error term.

Table 5 provides the results from estimating equation (4). As in our estimation of equations (1) through (3), we eliminate observations with studentized residuals greater than two in absolute value or Cook's D greater than one in the tabulated results; our inferences are unchanged if outliers are retained. We present the results only for the nonoverlapping sample with sufficient data to calculate the independent variables; the results are unchanged if we use the other approaches to deal with potential cross-sectional dependence.

In this specification, the coefficient estimate on *FCREV* ∗ *ALL_MEDIA* remains significantly positive (0.0059, $t = 2.34$, two-tailed $p < 0.02$). Thus, the effect of analyst media coverage on the market reaction to forecast revisions remains significant after controlling for performance variables examined in prior work.

The results in tables 2 and 4 indicate that market participants react more strongly to forecast revisions issued by analysts with more media coverage, even after controlling for the accuracy of the revised forecast and the award status of the analyst.

4.2.2. Media Coverage and Information Availability. In this section, we investigate whether our findings for the effects of analysts' media coverage on the market reaction to forecast revisions are due to the increased availability of forecast revisions issued by analysts with more media coverage. This increased availability would occur if the analyst's media mention contains the actual forecast revision. To investigate this possibility, we perform a content analysis of the media mentions for a randomly selected subsample of 226 analyst-year observations, representing 402 content items (274 in *MAG*, 59 in *WIRE*, 27 in *WSJ*, and 42 in *TV*).[25]

[25] To construct this random sample, we conduct a two-step sampling routine where we first select randomly a nonzero analyst-year observation and next select randomly a media observation from within the analyst-year observation. To ensure an equal representation of high- and low-celebrity analysts, we dichotomize each media category at the median number of media observations and sample randomly within these partitions. We construct sample sizes in proportion to the number of media observations in each media category (i.e., *MAG*, *WIRE*, *WSJ*, and *TV*), imposing a constraint of at least 20 media observations per category. Each randomly selected analyst-year observation represents a media article containing the analyst's

TABLE 5

Market Reaction to Forecast Revisions Conditional on Media Coverage and the Determinants of Forecast Accuracy: Nonoverlapping Forecast Accuracy Sample (N = 4,955)

Equation (4):

$$CAR_{i,j,t} = \phi_0 + \phi_1 FCREV_{i,j,t} + \phi_2 FCREV_{i,j,t} * PRMAPE_{i,j,q-1} + \phi_3 FCREV_{i,j,t} * RFCAGE_{i,j,q}$$
$$+ \phi_4 FCREV_{i,j,t} * RFIRMEXP_{i,j,q} + \phi_5 FCREV_{i,j,t} * RGENEXP_{i,j,q}$$
$$+ \phi_6 FCREV_{i,j,t} * RTURNOVER_{i,j,q} + \phi_7 FCREV_{i,j,t} * RFCFREQ_{i,j,q}$$
$$+ \phi_8 FCREV_{i,j,t} * RNOFIRM_{i,j,q} + \phi_9 FCREV_{i,j,t} * RNOIND_{i,j,q}$$
$$+ \phi_{10} FCREV_{i,j,t} * RBROKSIZE_{i,j,q} + \phi_{11} FCREV_{i,j,t} * RIIAWARD_{i,j,q}$$
$$+ \phi_{12} ALL\_MEDIA_{i,j,t} + \phi_{13} FCREV_{i,j,t} * ALL\_MEDIA_{i,j,t} + \upsilon_{i,j,q}$$

Variable		
Intercept	0.0008	0.0007
	(0.90)	(0.66)
FCREV	2.0120***	1.8253***
	(9.93)	(7.54)
FCREV PRMAPEq*−1	−0.9670***	−1.0882***
	(−2.69)	(−3.04)
FCREV RFCAGE*	0.0026	0.0017
	(0.24)	(0.15)
FCREV RFIRMEXP*	−0.0182	−0.0203
	(−0.46)	(−0.52)
FCREV RGENEXP*	0.0032**	0.0030**
	(2.13)	(1.96)
FCREV RTURNOVER*	−1.8966**	−1.2963*
	(−2.52)	(−1.67)
FCREV RFCFREQ*	−0.0026	−0.0024
	(−0.02)	(−0.02)
FCREV RNOFIRM*	0.0388	0.0494
	(0.72)	(0.91)
FCREV RNOIND*	−0.0497	−0.1657
	(−0.21)	(−0.72)
FCREV RBROKSIZE*	0.0178*	0.0106
	(1.66)	(0.97)

(*continued*)

Based on a pilot analysis, we generate the following three overall media content categories: analyst commentary concerning a followed company

first name, last name, and brokerage house name. The number of content items (402) differs from the number of analyst-year observations (226) because an observation can contain more than one content item (e.g., when an analyst comments on both a company's management and a company's industry within the same article).

TABLE 5 — *Continued*

Variable		
FCREV * *RIIAWARD*	−3.5850***	−3.4945***
	(−3.65)	(−3.54)
ALL_MEDIA		0.0000
		(0.14)
FCREV * *ALL_MEDIA*		0.0059**
		(2.34)
R^2	3.39%***	3.75%***
	(14.27)	(13.43)

This table provides results from estimating equation (4) on the forecast accuracy sample. For each variable included in equation (4), the coefficient estimate is presented; the *t*-statistic is provided in parentheses below the estimated coefficient. The *F*-statistic is provided in parentheses below the R^2. The results are provided for the nonoverlapping sample, which only includes those observations that have no other forecast revisions occurring during the return accumulation window. *PRMAPE* is the analyst's absolute forecast error for the firm and quarter, minus the mean absolute forecast error of other analysts following the firm for that quarter, deflated by the mean absolute forecast error. *FCAGE* is the number of calendar days between the firm's earnings announcement date and the forecast release date. *FIRMEXP* is the number of prior quarters for which the analyst has issued an earnings forecast for the firm. *GENEXP* is the number of prior firm quarters the analyst has issued a quarterly forecast for any firm on the Zacks database. *TURNOVER* is an indicator variable that equals 1 if the analyst changed brokerage houses during the year, 0 otherwise. *FCFREQ* is the number of quarterly earnings forecasts issued by the analyst for the firm and quarter. *NOFIRM* is the number of firms for which the analyst issues quarterly earnings announcement forecasts for the quarter. *NOIND* is the number of two-digit Standard Industrial Classification codes containing firms for which the analyst issues quarterly earnings announcements. *BROKSIZE* is the number of analysts issuing quarterly earnings forecasts at the brokerage house at which the analyst is employed. *IIAWARD* is an indicator variable that equals 1 if the analyst was named on the *Institutional Investor* All-American team in the previous year. The prefix "R" for the independent variables indicates that the mean value of the variable for the firm and quarter is subtracted from the observation's value. All other variables are defined in the notes to tables 1 and 3.

*, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively, two-tailed.

(*COMPANY*), general analyst mentions (*ANALYST*), and analyst commentary concerning an industry (*GEN_INDUSTRY*). The first category, *COMPANY*, consists of analyst discussions of several company-related issues, including management strategy, stock recommendations, the company's industry, products and services, other business-related comments, sales and sales revenue, completed and potential mergers and acquisitions, stock price in general, management in general, earnings forecasts, profits and losses, expenses, and price targets. The second category, *ANALYST*, consists of the following content subcategories: initiation of coverage on a company, an analyst changing from one brokerage firm to another, other analyst mentions, and analyst improprieties. The third category, *GEN_INDUSTRY*, is comprised of analyst comments about an industry in general rather than about a specific followed company.

Table 6 provides the results from this analysis. Across the three aggregate categories, 89% of the content focuses on a particular company, 4% on the analyst, and 7% on an industry in general. Examining the content pertaining directly to companies reveals that media mentions focus on general information such as management strategy (14%), the company relative to the industry (11%), success and profitability of products (9%), merger and acquisition activity (8%), stock price (5%), and management

TABLE 6

Content Analysis of Media Observations by Media Type

Category	%	TOTAL ($N = 226$)	*MAG* ($N = 140$)	*WIRE* ($N = 46$)	*WSJ* ($N = 20$)	*TV* ($N = 20$)
COMPANY						
STRATEGY	0.14	56	35	8	4	9
STOCK_REC	0.11	44	34	5	5	0
C_INDUSTRY	0.09	38	25	6	3	4
PRODUCT	0.09	37	20	7	1	9
C_OTHER	0.09	36	25	9	1	1
SALES	0.08	32	23	6	1	2
MA	0.08	31	21	5	1	4
PRICE	0.04	18	16	1	0	1
MANAGEMENT	0.04	16	13	0	1	2
FORECAST	0.04	16	16	0	0	0
EARN	0.04	15	8	4	2	1
EXPENSE	0.02	7	6	0	1	0
TARGET	0.01	6	5	0	1	0
ANALYST						
INITIATION	0.02	9	6	3	0	0
CHANGE	0.01	6	2	4	0	0
A_OTHER	0.01	6	1	0	4	1
IMPROPRIETY	0.00	1	0	1	0	0
GEN_INDUSTRY	0.07	28	18	0	2	8
Total	1.00	402	274	59	27	42

This table contains a summary of the content of a randomly selected sample of 226 analyst-year observations representing 402 content items. *COMPANY* represents analyst commentary concerning a followed company; *STRATEGY* represents analyst commentary concerning management strategy; *STOCK_REC* represents analyst commentary concerning stock recommendations; *C_INDUSTRY* represents analyst commentary concerning the company's industry; *PRODUCT* represents analyst commentary concerning products and services; *C_OTHER* represents analyst commentary concerning other business-related comments; *SALES* represents analyst commentary concerning sales and sales revenue; *MA* represents analyst commentary concerning completed and potential mergers and acquisitions; *PRICE* represents analyst commentary concerning stock price in general; *MANAGEMENT* represents analyst commentary concerning management in general; *FORECAST* represents analyst commentary concerning earnings forecasts; *EARN* represents analyst commentary concerning profits and losses; *EXPENSE* represents analyst commentary concerning expenses; *TARGET* represents analyst commentary concerning price targets; *ANALYST* represents general analyst mentions; *INITIATION* represents a mention of initiation of coverage on a company; *CHANGE* represents a mention of an analyst changing from one brokerage firm to another; *A_OTHER* represents other analyst mentions; *IMPROPRIETY* represents a mention of an analyst impropriety; and *GEN_INDUSTRY* represents analyst commentary concerning an industry.

in general (4%). The company content also focused on financial statement information such as sales (8%), earnings (4%), and expenses (2%). With regard to research outputs, the media articles mention stock recommendations most frequently (11%), followed by earnings forecasts (4%) and target prices (2%). Because earnings forecasts amount to only 4% of the total 402 content items provided, our content analysis suggests that it is unlikely that an analyst's higher level of media coverage captures the increased availability of his forecast revisions. This conclusion is reinforced by our finding that the estimated coefficient on $FCREV * MEDIA$ is significantly positive in equations (1) and (3) when we measure media coverage as of the end of the *prior* year.

Overall, we interpret our results as suggesting that the market reacts more strongly to forecast revisions issued by celebrity analysts whose names are more familiar. Our findings, however, are also consistent with the market reacting more strongly to forecast revisions issued by superior performing analysts if media coverage is correlated with measurement error in the proxies we examine for actual and perceived performance. To evaluate if the initial market reaction to analysts with higher levels of media coverage can be characterized as too strong, we next evaluate the returns around the earnings announcement date subsequent to the forecast revision date.

4.3 POSTREVISION RETURNS AROUND THE EARNINGS ANNOUNCEMENT DATE CONDITIONAL ON MEDIA COVERAGE

In this section, we investigate whether market participants appear to react too strongly to forecast revisions issued by analysts with high media coverage due to the celebrity status and resultant familiarity of those analysts. If the initial market reactions are not justified by appropriate proxies for analysts' actual forecast accuracy (as is suggested by the limited association between media coverage and our various measures of accuracy), market prices should reverse at the earnings announcement date, at which time analysts' actual accuracy is known.

To investigate this possibility, we examine the association between the excess returns surrounding the quarterly earnings announcement date and analysts' media coverage. To implement this test, we first select the firm-quarter observation corresponding to the analyst with the greatest media coverage for that firm and quarter; our results are unchanged if we use the observation corresponding to the analyst with the least media coverage or a randomly selected observation.[26] This yields a sample consisting of one forecast revision for each firm and quarter in our sample. We measure the excess return as the five-day size-adjusted buy-and-hold excess return $(-2, +2)$ surrounding the quarterly earnings announcement date.

If market participants react too strongly to forecast revisions issued by analysts with higher media coverage, there should be a positive price reaction at the quarterly earnings announcement date for those observations with a negative forecast revision issued by an analyst with higher media coverage. Similarly, there should be a negative price reaction when actual earnings are announced for those observations with a positive forecast revision issued by an analyst with higher media coverage. Thus, there should be a negative association between the excess return around the earnings announcement date and the interaction of the forecast revision and the analyst's media coverage. To test this conjecture, we estimate versions of the following equation:

[26] We select only one analyst forecast revision for each firm and quarter because the returns data are at the firm level rather than at the analyst level.

$$
\begin{aligned}
&EA\_CAR_{j,t} \\
&= \lambda_0 + \lambda_1 UE_{j,t} + \lambda_2 FIRM\_SIZE_{j,t} + \lambda_3 UE_{j,t} * FIRM\_SIZE_{j,t} + \lambda_4 LOSS_{j,t} \\
&\quad + \lambda_5 UE_{j,t} * LOSS_{j,t} + \lambda_6 LEV_{j,t-1} + \lambda_7 UE_{j,t} * LEV_{j,t-1} + \lambda_8 MKTBK_{j,t-1} \\
&\quad + \lambda_9 UE_{j,t} * MKTBK_{j,t-1} + \lambda_{10} BETA_{j,t} + \lambda_{11} UE_{j,t} * BETA_{j,t} \\
&\quad + \lambda_{12} UASALES_{j,t} + \lambda_{13} UE_{j,t} * UASALES_{j,t} + \lambda_{14} FCREV_{i,j,t} \\
&\quad + \lambda_{15} ALL\_MEDIA_{i,t} + \lambda_{16} FCREV_{i,j,t} * ALL\_MEDIA_{i,t} + \omega_{j,t} \qquad (5)
\end{aligned}
$$

where the variables not previously defined are:

$EA\_CAR_{j,t}$ = Five-day (−2, +2) size-adjusted buy-and-hold excess return around firm *j*'s earnings announcement date at time *t*;

$UE_{j,t}$ = Unexpected earnings for firm *j* at time *t* (defined as actual earnings less the revised analyst forecast, deflated by price 10 trading days before the release of the revised forecast);

$LOSS_{j,t}$ = An indicator variable that equals 1 if actual earnings for firm *j* at time *t* are less than zero, 0 otherwise;

$LEV_{j,t-1}$ = Leverage for firm *j* at time *t*−1 (defined as the ratio of long-term debt to total book value of assets);

$MKTBK_{j,t-1}$ = Market-to-book ratio for firm *j* at time *t*−1 (defined as the ratio of price to book value per share);[27]

$BETA_{j,t}$ = Beta for firm *j* at time *t* (obtained from a firm-specific regression of the firm's daily return on the value-weighted market index daily return using the 100 trading days ending 10 days before the quarterly earnings announcement date);[28]

$UASALES_{j,t}$ = The absolute value of unexpected sales for firm *j* at time *t* (defined as sales for the current quarter less sales from the same quarter of the prior year, deflated by market value of equity as of the end of the prior quarter); and

$\omega_{j,t}$ = Error term.

Table 7 provides the results from estimating equation (5). As in our estimation of equations (1) through (4), we eliminate observations with studentized residuals greater than two in absolute value or Cook's D greater than one in the tabulated results; our inferences are unchanged if outliers are retained. We present the results only for *ALL_MEDIA*; the results are unchanged if we replace *ALL_MEDIA* with *MAG*, *WIRE*, *WSJ*, or *TV*.

We find that, after controlling for the news at the earnings announcement date, the estimated coefficient on *FCREV* ∗ *ALL_MEDIA* is significantly

[27] If leverage or market-to-book is missing for the prior quarter, we use the leverage or market-to-book for the current quarter.

[28] We require at least 75 nonmissing observations to estimate a firm's beta.

TABLE 7

Earnings Announcement Date Returns Conditional on Media Coverage

Equation (5):

$$EA\_CAR_{j,t} = \lambda_0 + \lambda_1 UE_{j,t} + \lambda_2 FIRM\_SIZE_{j,t} + \lambda_3 UE_{j,t} * FIRM\_SIZE_{j,t} + \lambda_4 LOSS_{j,t}$$
$$+ \lambda_5 UE_{j,t} * LOSS_{j,t} + \lambda_6 LEV_{j,t-1} + \lambda_7 UE_{j,t} * LEV_{j,t-1} + \lambda_8 MKTBK_{j,t-1}$$
$$+ \lambda_9 UE_{j,t} * MKTBK_{j,t-1} + \lambda_{10} BETA_{j,t} + \lambda_{11} UE_{j,t} * BETA_{j,t}$$
$$+ \lambda_{12} UASALES_{j,t} + \lambda_{13} UE_{j,t} * UASALES_{j,t} + \lambda_{14} FCREV_{i,j,t}$$
$$+ \lambda_{15} ALL\_MEDIA_{i,t} + \lambda_{16} FCREV_{i,j,t} * ALL\_MEDIA_{i,t} + \omega_{j,t}$$

Variable			
Intercept	0.0033***	0.0061***	0.0123***
	(4.17)	(5.75)	(4.08)
UE	0.1886***	0.2004***	0.4881***
	(9.07)	(9.15)	(4.75)
FIRM_SIZE			−0.0006
			(−1.41)
UE * *FIRM_SIZE*			−0.0263**
			(−2.46)
LOSS			−0.0180***
			(−6.84)
UE * *LOSS*			−0.2593***
			(−3.13)
LEV			−0.0066
			(−1.20)
UE * *LEV*			0.1378
			(1.20)
MKTBK			−0.0000
			(−0.19)
UE * *MKTBK*			0.0088
			(1.28)
BETA			0.0020
			(1.25)
UE * *BETA*			−0.0259
			(−0.78)
UASALES			0.0058
			(0.58)
UE * *UASALES*			−0.0118
			(−0.09)
FCREV		0.4203***	0.2873**
		(3.34)	(2.10)
ALL_MEDIA		−0.0000**	−0.0000**
		(−2.07)	(−2.21)
FCREV * *ALL_MEDIA*		−0.0019*	−0.0019*
		(−1.89)	(−1.76)

(*continued*)

TABLE 7 —*Continued*

Variable			
R^2	0.92%***	1.14%***	1.86%***
	(82.24)	(26.08)	(10.89)

This table provides results from estimating equation (5) on the forecast revision sample. For each variable included in equation (5), the coefficient estimate is presented; the *t*-statistic is provided in parentheses below the estimated coefficient. The *F*-statistic is provided in parentheses below the R^2. The sample contains forecast revisions in the nonoverlapping sample issued by the sample analyst with the most media coverage for that firm and quarter. *EA_CAR* is the size-adjusted excess return in the five-day period (−2, +2) surrounding the earnings announcement date; *UE* is the unexpected earnings (defined as actual earnings less the revised analyst forecast, deflated by price 10 days before the release of the revised forecast); *LOSS* is an indictor variable that equals 1 if actual earnings is less than zero; *LEV* is the ratio of long-term debt to total book value of assets; *MKTBK* is the ratio of price-to-book value per share; *BETA* is the stock market beta; *UASALES* is the absolute value of unexpected sales (defined as actual sales less sales from the same quarter of the prior year, deflated by the market value of equity at the end of the prior quarter); all other variables are defined in the notes to tables 1, 3, and 4.

*, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively, two-tailed.

negative (−0.0019, $t = -1.89$, two-tailed $p < 0.06$). This finding holds even when variables previously shown to be associated with the market reaction to unexpected earnings are included in equation (5) as control variables (see, e.g., Francis, Schipper, and Vincent [2002], Hayn [1995], Dhaliwal, Lee, and Fargher [1991], Collins and Kothari [1989]). The significant association between the excess returns at the quarterly earnings announcement date and *FCREV* ∗ *ALL_MEDIA* is consistent with market participants initially reacting too strongly to forecast revisions issued by analysts with high levels of media coverage. This analysis, however, does not indicate whether a trading strategy based on these findings would yield profitable returns after incorporating the costs of implementing the strategy.

5. *Conclusions*

We examine the effect of analyst celebrity, measured by the media coverage received by sell-side analysts, on investor reaction to the analysts' forecast revisions. We use the DJI database to measure media coverage in total and in four specific categories (*The Wall Street Journal*, magazines and newspapers, newswires, and television and radio) for a random sample of sell-side analysts. We find that all measures of media coverage are positively related to investor reaction to the forecast revisions in the short window (−2, +2) surrounding the release. This result holds when we allow the market reaction to vary with forecast accuracy and analyst award status. These findings are consistent with analysts' familiarity affecting investor reaction to forecast revisions. However, we cannot rule out that our measure of celebrity is either correlated with measurement error in our proxies for forecast performance and/or correlated with other unexamined dimensions of forecast performance. Content analysis on a randomly selected subsample of media mentions indicates that it is unlikely that these findings are due to the increased availability of forecast revisions. The results of our postrevision returns analysis are consistent with market participants reacting too strongly to forecast revisions issued by analysts with high media coverage.

Our findings contribute to the literature as follows. First, we provide evidence that the media plays a role in the stock price formation process beyond its effects due to coverage of firms. Specifically, our results suggest that the media's coverage of analysts, key information intermediaries, also affects stock prices. Second, we add to the research on investor reaction to analysts' forecast revisions. This literature indicates that substantial variation in investor reaction to forecast revisions is not accounted for by previously examined performance-related factors. We investigate one property of forecasts—the media coverage of the analysts issuing those forecasts—that appears to account for part of this unexplained variation. Third, we examine an economic consequence of reacting to the media coverage aspect of forecast revisions through our investigation of the postrevision returns around the earnings announcement date. The association between the excess returns around the earnings announcement date and the analyst's media coverage is consistent with investors initially reacting too strongly to forecast revisions issued by celebrity analysts.

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