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Market effects of local media employment reductions on the idiosyncratic risk of nearby firms; returns, valuation, and debt; and firm meet-beat behavior and CEO turnover and compensation

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INTRODUCTION

Media play an important role in capital markets (Tetlock, 2014), but the number of newspaper and broadcast journalists in the United States has fallen by a third in the past 20 years, according to U.S. Bureau of Labor Statistics estimates. Many of those losses have come from cuts at local and regional news organizations, and the role local journalists play in capital markets is less clear.

The reductions in the number of newsgatherers in the United States raise important questions about the effectiveness of democratic institutions (Schulhofer-Wohl and Garrido, 2013; Drago, Nannicini and Sobbrio, 2019) and monitoring of public officials and governments (Nyhan and Reifler, 2015; Gao, Lee and Murphy, 2019). I find that it also has implications in equity investment decisions. A reduction in media employment in the metropolitan statistical area of a company's headquarters is related to an increase in the standard deviation of a firm's stock returns and that increase appears to be driven by increased idiosyncratic risk of the firm. Employing a dynamic difference-in-differences design, I find that reduced potential news coverage of a firm affects its idiosyncratic risk for several years after the media employment reduction is recorded.

I also explore the channels through which a reduction in local media employment might influence firm-level idiosyncratic risk. Attention models (Klein and Bawa, 1977; Merton, 1987) predict that declines in investor awareness will decrease the informativeness of markets. Real effects on product markets can also play a role in idiosyncratic risk (Gaspar and Massa, 2006; Irvine and Pontiff, 2009), with decreasing market power being associated with increased volatility of returns. Finally, the reduced levels of information available to investors may increase their estimation risk, reducing certainty about assessed fair values.

Through a series of empirical tests, I find little support for the hypotheses that investor attention decreases or that competition increases after a firm is affected by a reduction in journalists employed nearby. I find evidence that the increased idiosyncratic risk of affected firms is instead driven by increased estimation risk or information asymmetry among investors.

In addition to identifying a role for local journalists in capital markets, this research contributes in two other primary ways. First, prior research on the media's effect on firm performance has focused on short-term impacts and reversals (Easley and O'Hara, 2004; Fang and Peress, 2009; Fedyk, 2018) rather than the longer-term effect on capital markets and asset pricing. Second, prior research has exploited legacy print publication (Huberman and Regev, 2001; Barber and Odean, 2008; Engelberg and Parsons, 2011) and television broadcast schedules (Busse and Green, 2002; Neumann and Kenny, 2007; Engelberg, Sasseville and Williams, 2012) to identify effects of information release. However, the prevalence of Internet-only outlets, as well as web and social media use by traditional print and broadcast outlets, has made it difficult to determine whether the information landscape has changed in such a way as to diminish the role of traditional news reporters and broadcast analysts in financial markets. My findings provide evidence that journalists continue to play a significant role in the capital markets. My sample considers the period from 2003, the year after many of the oft-cited examples of media effects on financial markets end (Barber and Odean, 2008; Fang and Peress, 2009; Engelberg and Parsons, 2011). I find that a reduction in employees categorized as "reporters and correspondents" and "broadcast news analysts" in a local market continues to affect firms in ways consistent with previous findings, suggesting that the media's role in capital markets persists, perhaps supplemental to changes to technology and theories about those changes effect on investor behavior.

MOTIVATION

Theoretical models

Much of the previous literature on the role media play in capital markets has focused on national or international news outlets. Relatively little has addressed how local or state reporting may affect firm risk and value. However, many of the information models proposed in the literature are potentially affected by reporting at all levels. In a simple model, Neuhierl, Scherbina and Schlusche (2013) consider that news (corporate disclosures, in their setting) may lead investors to revise both their estimate of the fair value, v, of an asset and the precision with which it is estimated, σv. The arrival of value relevant information thus replaces the investor's prior valuation model VFIRM ~ N(v, σv), with the new one, VFIRM ~ N(vNEW, σv,NEW). A reduction in the number of local journalists able to reveal information about a nearby firm does not entail, ex ante, a change in the fair value estimate of the stock, v. However, a change to the information environment around the firm may lead to a change in future estimations about the precision with which that expected value is forecast (i.e., a revision to σv). The change in precision in which the mean return is estimated, σv,NEW - σv, could lead to increase in the volatility of the returns, specifically the firm-specific, or idiosyncratic, risk, particularly if the added estimation risk creates a wider expected variance of values of v. If increased estimation risk is interpreted by investors as increased information asymmetry, increasing estimation risk may also drive market prices lower, even if average estimates of value hold constant. An increased estimation risk may also affect volatility of returns if, as Neuhierl, Scherbina and Schlusche (2013) argue, when valuation model assumptions are weakened, (i.e., σv,NEW > σv), then a security's price reacts more strongly to subsequent information signals.

Klein and Bawa (1977) also consider estimation risk in their attention model, and conclude that the effect of increased estimation risk is, for risk-averse investors, to move more of their investment into a riskless asset. Grossman and Stiglitz (1980) model price informativeness as more uninformed investors move out of the market for a security, and market-makers face an increased risk of adverse selection from a higher concentration of informed investors (Glosten and Milgrom, 1985; Kyle, 1985). Informed investors profit from uninformed investors and market-makers, so information becomes less valuable as the proportion of informed investors increases and the market becomes less profitable to the informed investors who remain in the market. Combined, Klein and Bawa (1977) and Grossman and Stiglitz (1980) suggest that the volatility of prices could increase as remaining informed traders realize smaller profits, market-makers widen the bid-ask spread to hedge against adverse selection, and a lower concentration of uninformed investors participate in the market for a security.

Merton (1987) proposes a model in which investor attention, not estimation risk, is the relevant constraint. Although investors can potentially diversify away all idiosyncratic risk by including an infinite number of assets in their portfolios, their actual portfolios are functionally limited to assets with which they are aware. Merton models "awareness" as both knowing about a firm, i.e., that it exists as an investment opportunity, and being able to properly assess parameters that affect its operation, i.e. estimate its value.[1] If investors are "unaware" of an asset, either because they do not know that an investment opportunity for a firm exists or because they cannot properly estimate its value, then it cannot be included in their portfolios. (Obviously, in reality, the former is a more binding constraint to investment, and limiting an investor's ability to effectively estimate value-drivers could mean increased estimation risk rather than an inability or

[1] In Merton's model, only one level of information exists and incorporates both awareness and equal assessments of fair value.

unwillingness to invest in the asset at all.) Merton's model suggests that firms that fall short of the optimal number of investors will have to deliver higher returns to current investors because of the additional risk those investors carry and that should awareness increase, more investors will buy securities in the firm and the value of the company will increase.

Local media as information provider

For attention and information models to apply to local reductions in media employment, local media must, of course, provide at least the perception of value-relevant information. A variety of research has examined the role of the financial press, at the national and international level, as an information intermediary in capital markets and in security asset pricing. The number of headlines about a firm contribute to momentum effects on asset prices (Chan, 2003). A lack of news stories in major national outlets can mean investors require a higher return from a company (Fang and Peress, 2009). Financial journalism and sell-side financial analysts play complementary roles as information intermediaries (Ahn et al., 2019; Guest and Kim, 2019). Media coverage can help investors interpret information (Huberman and Regev, 2001; Demers and Lewellen, 2003; Bushee et al., 2010; Twedt, 2016; Guest, 2017), but investors can also overreact to "stale" news simply because it is repeated prominently (Carvalho, Klagge and Moench, 2011; Fedyk and Hodson, 2014; Marshall, Visaltanachoti and Cooper, 2014; Tetlock, 2014).

Anecdotally, local news sources have sometimes been in a unique position to provide investors information relevant to firm valuation. In January 2005, shortly after being named CEO of RadioShack, David Edmondson was arrested for driving under the influence in Fort Worth, Texas, where the company is headquartered. It was his third drunken driving arrest, and the local newspaper, the *Fort Worth Star-Telegram,* began looking into Edmondson's past. About a year later, Star-Telegram retail reporter Heather Landy revealed that Edmondson had lied about two of the degrees listed on his resume. The story was picked up by national news outlets across the country, including *Bloomberg News Service,* the *Wall Street Journal,* and the *New York Times.* A week after Landy's story was published, RadioShack's board announced his resignation, in February 2006. The day Landy's story first appeared, RadioShack's board issued a statement saying, it was "aware of the matters raised in the Fort Worth Star-Telegram article and has previously given due consideration to them." However, in announcing Edmondson's resignation a week later, RadioShack's executive chairman, Leonard H. Roberts, admitted, the board knew "some, but definitely not all" of the issues raised in Landy's reporting.[2] In the week between Landy's story being published and Edmondson's resignation, RadioShack shares fell more than 10 percent.

While some research exists on the aggregate effects of local media reporting on trading, it is unclear that local journalists generally provide information that is valuable to price discovery. Barber and Odean (2008) and Engelberg and Parsons (2011) find that coverage by local media outlets encourages trading activity within their coverage areas and that this increased volume of trading is associated with increased prices. Gurun and Butler (2012) find that local media are subject to "hype" when covering firms headquartered nearby. They determine that local news outlets are more likely to cover local firms with a positive slant and suggest that media outlets are, in a sense, captured because of their advertising relationships. Gurun and Butler (2012) find that security prices increase temporarily as a result of this hype, but that the prices reverse shortly after the positive coverage. In a survey of journalists, which includes several local news outlets, Call et al. (2018) find that journalists acknowledge pressure from management to avoid unfavorable

[2] New York Times, Floyd Norris, "RadioShack Chief Resigns After Lying," Feb. 21, 2006.

stories, but that they consider monitoring companies one of journalism's most important functions. Call et al. (2018) also find that journalists have incentives to produce high-quality articles with exclusive content and that negative articles have the most impact.

Overall, prior research has not determined conclusively whether journalists provide value-relevant information or simply encourage noise trading. Furthermore, the explosion of Internet-only outlets and social media sites (Antweiler and Frank, 2004) may have made the possible mechanisms by which local news media produce and disseminate value-relevant information, well, irrelevant. However, the FCC notes that "an abundance of media outlets does not translate into an abundance of reporting" (Waldman, 2011). In fact, studies by the Pew Center for Excellence in Journalism and others suggest that the vast majority, sometimes as much as 95%, of stories collected and shared by all media originates with what are called "legacy" media, i.e., a newspaper or local broadcast station. Just as investors face attention constraints, national and international news outlets and agencies may also lack resources to uncover many stories at the local and state level and rely on local media outlets to find stories they can aggregate or report more thoroughly. Therefore, despite the increasing number of outlets online, "TV stations and newspapers have emerged as the largest providers of local news online" (Waldman, 2011).

Cage, Herve and Viaud (2019) also examine the sources of online news, initially using a "transmedia" approach agnostic to which type of media company originally published new information and then tracking back the original source. They find that almost two-thirds of articles contain at least some copied material, and that original stories are disseminated by other media outlets in under 3 hours on average, and sometimes in as little as 4 minutes. Almost three-quarters of the original content that did not originate with news agencies, such as AFP and Reuters, came from newspapers, while 11.5% was from television news stations. Radio and online-only news sources account for about 10% and 7% of original content, respectively.

HYPOTHESIS DEVELOPMENT

Idiosyncratic risk

Reduction in local media coverage could increase the idiosyncratic risk of firms within their coverage area through declining attention, increased product market competition, or increased estimation risk. It is also possible that local media, i.e. reporters and broadcast journalists located within a metropolitan statistical area or other geographic region, provide no new incremental information or awareness over national media outlets and electronic sources. Idiosyncratic risk is not directly observable (Xu and Malkiel, 2003), as a security's standard deviation captures both the systematic risk, as typically measured by the market beta3, as well as the firm-specific, or idiosyncratic risk. To isolate the effect of media employment reductions on idiosyncratic risk specifically, I employ two indirect methods. First, I match treated firms with a control sample by size (log of market value) and market beta, as computed by regressions of daily returns each year for each firm with the Fama-French Three-Factor Model (Fama and French, 1993, 1997). The calculated beta from these regressions finds the correlation of returns for individual securities to the overall market. The small-minus-big (SMB) factor measures the difference between the excess return on a portfolio of small versus big capitalization stocks, and the high-minus-low (HML) factor is the difference between the excess return on a portfolio of high book-to-market stocks

[3] As in the capital asset pricing model of Sharpe (1964) and Lintner (1965), the Fama-French Three Factor Model in Fama and French (1992 and 1993), the Carhart Four Factor Model in Carhart (1997) and others.

versus low book-to-market stocks. Fama and French suggest that these factors incorporate elements of systematic risk distinct from the firm's relationship to the overall market.[4]

Because treated firms are matched with control firms on market beta, any differences due to systematic market risk should be minimized. However, I also hypothesize that a more direct test of a firm's idiosyncratic risk, the firm-specific residuals from the Fama-French Three Factor regressions, will also demonstrate significant increases among treated firms when compared with control firms.

Hypothesis 1: A decrease in local media reporters and broadcast analysts leads to increased standard deviations of returns (and increased root mean squared errors from Fama-French Three Factor regressions) for firms headquartered in those geographic areas.

If, as demonstrated in the Results section, the idiosyncratic risk of firms located within geographic areas that experience a decrease in local media employment increases significantly, that increased risk may be the result of decreased awareness about the firm's securities as an investment option. It is also possible that reducing the number of reporters or broadcasters available to share firm information has product market consequences for companies (Grullon, Kanatas and Weston, 2004). Increased idiosyncratic risk may also be due to higher levels of estimation risk of value for a firm in which investors now perceive themselves to be more at risk of adverse selection from informed traders and firm insiders.

[4] In unreported results, I also employ the market model and the Four Factor model of Carhart (1997) to match on market beta and collect residuals. Results are consistent regardless of which method is employed; however, because the Carhart momentum factor may be correlated with attention, I use Three Factor matching and results in my reported results.

I form affirmative hypotheses for each possible channel of increased idiosyncratic risk, decreased attention, reduced market power (i.e., increased competition), and estimation risk.

Attention

Although it is unlikely that previously aware investors become unaware of a stock simply because it is not mentioned as frequently in the media they consume, Tversky and Kahneman (1973, 1974) and Shiller (1980) suggest that individuals suffer from recency bias. And Yuan (2015) finds that attention-grabbing events predict trading behavior and market returns. If investors overweight the most recent information they have received, they may 1) shift their investments from securities that have previously received coverage to stocks that have received attention more recently or 2) focus new investment activity on securities in which they have more recent information. Whether these investors are considered informed or uninformed (Grossman and Stiglitz, 1980) could affect whether they realize gains or losses as they benefit or suffer from adverse selection . Regardless, however, if attention is declining, it may be reflected in at least two other independent settings. Da, Engelberg and Gao (2011) find that Google search traffic for a firm's ticker symbol is related to overall investor attention in the security, and my first attention-related hypothesis compares reductions in local media coverage with the Da, Engelberg and Gao (2011) measure of investor attention:

Hypothesis 2: A decline in local media employment in a firm's geographic region is accompanied by a decline in overall investor attention for the firm as measured by Google search traffic.

I also examine levels of overall trading activity and whether a reduction in media employment is correlated with decreased activity. Barber and Odean (2008) and Engelberg and Parsons (2011) suggest that recent news about a firm, regardless of whether it reveals new information, drives an increase in local trading. Barber and Odean (2008) find that retail investors, who are likely to be relatively less informed traders, buy attention-grabbing stocks rather than sell on news.

A complication arises in testing trade volume. The SEC estimates that more than half of all trades are now conducted by high-speed computer algorithms, also known as "high-frequency trading" (Gerig, 2015). Menkveld (2013), however, finds that HFTs operate much like traditional market-makers, incurring losses on their inventory and profiting on the bid-ask spread as compensation for the risk of adverse selection. In the dynamic difference-in-differences setting, overall trends in number of trades should be absorbed so only the group-specific number of abnormal trades associated with the treated firms are detected. My second hypothesis posits that investor attention declines after a reduction in local media employment:

Hypothesis 3: A decline in local media employment in a firm's geographic region is accompanied by a decline in overall trading activity in the corporation's stock.

Competition

Even if attention levels remain constant, a firm's idiosyncratic risk can be affected by increasingly competitive environments (Gaspar and Massa, 2006; Irvine and Pontiff, 2009; Abdoh and Varela, 2017). Gaspar and Massa (2006) find that firms can use their market power to pass along a larger proportion of idiosyncratic cost shocks to their consumers. They also find that an ability to avoid competition decreases uncertainty about firms' future performance. Irvine and

Pontiff (2009) find that increased idiosyncratic risk is tied directly to increased volatility of cash flows. Abdoh and Varela (2017) find that higher levels of competition are associated with higher proportions of idiosyncratic risk. Reductions in media employment could have product market or cash flow implications for firms if affected companies face higher levels of competition because they lose a channel for disclosing firm and product news or a medium for advertising is eliminated or less effective. Gurun and Butler (2012) find that local media outlets essentially reward advertising dollars with positive coverage of advertisers. If, as Gurun and Butler (2012) find, local media outlets provide another outlet for firm advertising or marketing in news stories because of their advertising relationships, then a reduction in the number of reporters or broadcasters able to produce favorable stories could affect firms' product markets (and also potentially their investor relations). If firms face increased competition because of the loss of this "hype," then increased idiosyncratic risk may be the result of effects on cash flow volatility or market power. My next hypothesis relates reductions in local media employment with increased levels of competition.

Hypothesis 4: A decline in local media employment in a firm's geographic region reduces a firm's market power (concentration) and increases competition within an industry.

Estimation risk

Next, I consider another possible channel through which a decline in local media coverage may be linked to a firm's increased idiosyncratic risk. In addition to being reflected in idiosyncratic risk, increased levels of estimation risk will be reflected in increasing levels of perceived information asymmetry. Whether local news stories reveal genuinely new information about a firm, any coverage may give retail investors a sense of firm value through assurances of monitoring or disclosure about cash flow information. If the information environment is understood by investors to be worse after a reduction in local media, then the perceived probability of facing adverse selection increases. The information models of Klein and Bawa (1977), Grossman and Stiglitz (1980), and Glosten and Milgrom (1985) suggest that bid-ask spreads will increase as information asymmetry between informed and uninformed investors increases and as informed investors make up larger percentages of the market for a security. Most information models consider "informed" and "uninformed" as a binary; however, real markets reflect varying levels of information. Investors who become aware of a security through local, or even national, media coverage could potentially be classified either as informed or uniformed investors. However, most prior research suggests that the majority of retail stock-pickers acting on information relayed from the media will, in aggregate, function as uninformed traders (Barber and Odean, 2008; Engelberg and Parsons, 2011).

Chan (2003) finds that the number of headlines about a firm contribute to momentum effects on asset prices. Tetlock (2011); Carvalho, Lagge, and Moench (2011); and Fedyk and Hodson (2019) find that investors overreact to news that has already been revealed simply because it is repeated. Barber and Odean (2008) and Engelberg and Parsons (2011) find that coverage by local media outlets encourages trading activity within their coverage areas, and Huberman (2001) finds that regional investment reflects overinvestment in familiar firms, beyond levels that would be reflected in a rationally diversified portfolio. Fedyk (2018) finds that even finance professionals' trading decisions can be affected by placement of headlines on Bloomberg terminals independent of those headlines' news value.

Gurun and Butler (2012) find that local media tend to "hype" local firms that advertise with the news outlet and price increases are followed by quick reversals. Grullon, Kanatas and Weston (2004) find that product market advertising can also serve as investment advertising and can increase the number of retail shareholders of the firm. Given these findings, I treat investors influenced by local media coverage as relatively "uninformed" and subject to adverse selection as in Grossman and Stiglitz (1980) and Glosten and Milgrom (1985).

Evaluating estimation risk using a common proxy for information asymmetry is problematic. The effects, as in Merton (1987), of declining attention and increased estimation risk would be identical when evaluating information asymmetry, which is commonly proxied by the bid-ask spread (Glosten and Milgrom, 1985; McInish and Wood, 1992; Lee, Mucklow and Ready, 1993; Chan, Christie and Schultz, 1995; Grullon, Kanatas and Weston, 2004; Carlin, Longstaff and Matoba, 2014). If uninformed investors move to decrease or eliminate a security from their portfolios because they are (or become, with recency bias) relatively unaware of a company, fewer uninformed investors are available to provide liquidity to the market for the asset, and bid-ask spreads increase. If fewer uninformed investors participate in a security's trading because they no longer believe their assessments of the value of the security are reliable then the effect is the same: fewer uninformed investors in the market for the security, reduced liquidity, and increased bid-ask spreads. If however, as I demonstrate in the Results section, overall levels of attention and trading appear unaffected by reductions in local media employment, then statistically significant increases in the bid-ask spread may imply that (uninformed) retail investors, specifically, decline to invest in securities with which they are familiar but unable to form reliable valuation assessments.

Hypothesis 5: Even contingent on attention levels of securities being similar, the bid-ask spread of firms headquartered in locations where local media employment declines will increase relative to firms unaffected by such a local media employment shock.

Hypothesis 5 examines a possible effect realized when uninformed, or noise, traders are no longer able to arrive at a fair valuation of an asset because of increased information asymmetry. To confirm the effect on retail investors, I explore another hypothesis. Institutional investors are generally classified as "informed" investors (Gompers and Metrick, 2001; Chan and Lakonishok, 2004). As large investors, institutional investors frequently hold large blocks of a security, which means firms held by institutional investors have relatively fewer overall shareholders. Therefore, if decreasing media coverage increases estimation risk among retail investors, the effect on bid-ask spread will be more pronounced when the number of shareholders is relatively higher.

Hypothesis 6: Firms headquartered in locations where local media employment declines will have relatively higher bid-ask spreads as the number of shareholders increases.

DATA AND DESCRIPTIVE STATISTICS

Bureau of Labor Statistics data

Information on the number of reporters and correspondents in an area is taken from the Bureau of Labor Statistics' Occupational Employment Statistics program. The survey produces annual estimates of employment for 810 specific occupations in more than 580 areas, including metropolitan statistical areas (MSAs) and nonmetropolitan areas throughout the U.S. states, the District of Columbia, and U.S. territories. Because of the way the survey is conducted, its results cannot be used for time-series analysis. The OES "surveys approximately 180,000 to 200,000 establishments per panel (every six months), taking three years to fully collect the sample of 1.2 million establishments. To reduce respondent burden, the collection is on a three-year survey cycle that ensures that establishments are surveyed at most once every three years." The data are collected for the reference months of May and November by state workforce agencies, with BLS-provided surveys and instructions. Responses are collected by mail, email, Internet or other electronic means, telephone, or in-person visit. The statistics are compiled annually after each year's May survey and are released in March of the following year.

I use the total numbers of employees in the Reporter and Correspondent categories, which include both the Publishing Industries, such as newspapers and periodicals, and the Radio and Television Broadcasting Industries. These two industries account for almost 90 percent of the employees counted as Reporters and Correspondents in the May 2018 estimates. The remainder were listed as employees of Other Information Services; Motion Picture and Video Industries; and Colleges, Universities, and Professional Schools.

The employment total used in my analysis also includes Broadcast News Analysts, which are listed in one category with Reporters and Correspondents in some years. The Office of Management and Budget (OMB) Standard Occupational Classification (SOC) System describes Reporters and Correspondents as employees who "collect and analyze facts about newsworthy events by interview, investigation, or observation" and "report and write stories for newspaper, news magazine, radio, or television." Broadcast News Analysts are employees who "analyze, interpret, and broadcast news received from various sources." The definition suggests Broadcast News Analysts may also reveal or convey salient information to or generate attention from investors. I include the category in my analysis both because of the possibility that they reveal relevant information and to retain comparability throughout the sample period.

I calculate the percentage change to employees in the sum of these two categories across two-year windows to account for the rolling nature of the survey. Years in which the number of employees in the Reporters and Correspondents and Broadcast News Analyst categories falls 25 percent or more from the OES survey released two years prior are identified as negative shocks to coverage. The OES has released results of the metropolitan area-level survey since 1997; however, the category descriptions have only been standardized since 2001, when the OES survey began using the North American Industry Classification System (NAICS). Prior data are based on the Standard Industrial Classification (SIC) system. Although the classification change is unlikely to affect the particulars of this analysis, I limit observations of shocks to employment to 2003, as measured by the difference from 2001, and beyond.

Sample description

The first, and largest number, of firms with negative shocks are recorded in 2003, as shown in Figure 1, with more than 250 firms headquartered in geographic areas that were affected. Reductions in media employment affected more than 100 firms in 2004, 2005, 2007, 2008, and 2009. As Gao, Lee, and Murphy (2019) show, shocks to employment are not uniform across time. In their sample, newspapers closed within counties at different periods, and they use the introduction of Craigslist to a market as an instrument for a newspaper closure, because of the ensuing loss of classified advertising revenue. Craigslist was founded in 1995 in San Francisco and entered the largest metropolitan areas first, which suggests that cities with the headquarters of a relatively larger number of publicly traded firms may have been affected earlier than other areas. The last affected year considered in this analysis is 2015 to allow at least three years of post-shock returns and accounting information. Figure 2 shows where affected firms are located, by state. As is to be expected because of the number of firms headquartered in these states, the heaviest concentrations are in California, New York, and Texas, but the firms included in my analysis span 41 states and Washington, D.C.

The Metropolitan Statistical Areas used by the BLS are designated by the U.S. Census Bureau. I link ZIP codes to MSAs using Census Bureau designations for 2010, and identify the ZIP code of a firm's headquarters using Compustat listings for its mailing address, because many companies incorporate outside of the state where their operational headquarters may be located for tax and governance reasons. To ensure that total changes in overall employment are not affecting the results, I drop observations in which a 25% reduction in the sum of Reporters and Correspondents and Broadcast News Analysts is accompanied by a reduction in overall employment as recorded by the OES. By requiring that overall employment has not fallen by more than 5 percent, 196 firms are eliminated from the analysis.

Bhojraj, Lee and Oler (2003) find that the Global Industry Classifications Standard (GICS) explains stock return comovements better than Standard Industry Classification (SIC) and the North American Industry Classification System (NAICS) classification systems, and Levi and Welch (2017) find that firm market value strongly correlates with a firm's beta and other market model factor similarities when compared with other firms. I find exact matches for each sample firm by year and by 6-digit G industry code, and then match, without replacement, by market value within 15% and, because I am attempting to isolate the effects of idiosyncratic risk, within 15% of the market beta of the treated firm's calculated beta in the Fama-French Three Factor model regressions. The Fama-French Three Factor model is:

$$R_{i,t} - R_{f,t} = \alpha_i + \beta_i^M \left(R_t^M - R_{f,t}\right) + \beta_t^{SMB} R_t^{SMB} + \beta_i^{HML} R_t^{HML} + \varepsilon_{i,t}$$

Where $R_{i,t}$ is the return to the firm *i* at time *t*, $R_{f,t}$ is the risk-free rate, R_t^M is the return to a value-weighted market portfolio, R_t^{SMB} is the difference of the return to small market value stocks minus larger firms, and R_t^{HML} is the difference between the return to high book-to-market stocks and low book-to-market firms. The error term, ε_i, is the firm specific residual from the regression, which measures the idiosyncratic risk of the firm as separate from the systematic risk factors measured by β_i^M, β_t^{SMB}, and β_i^{HML}. Regressions use daily returns for firms and the value-weighted market portfolio from the CRSP database, and daily returns to the small-minus-big and the high-minus-low portfolios, R_t^{SMB} and R_t^{HML}, are from the North American factor file on Kenneth French's website.

I eliminate firms with relevant missing control variable data from Compustat or CRSP and firms with common stock share prices that closed the year trading under \$2 to mitigate small changes in share price driving high variation in returns. I also eliminate firms with fewer than 160 days of trading data available each year. My main findings are robust to a number of matching procedures, including matching with one, five, or as many as possible control firms with replacement and matching firms by their headquarter regions' overall employment levels, rather than beta, after matching by market value. I require that matching firms have not experienced a reduction in media employment in the area in which they are headquartered in any firm-year during the 13-year period for which each firm is analyzed. In the final sample used, it is not possible to find a matching firm for 1,291 companies in my treated sample[5]. I drop them from my analysis to ensure that results are not driven by unusually small or large firms or those with unusual market risk, as measured by market beta. When those firms are included in the analysis, results are stronger[6] than reported here.

My final sample consists of 1,373 firms headquartered in locations that experienced media employment reductions and their 1,373 matching control firms. Summary statistics for the treated

[5] A similar number of firms are unmatched using other procedures.
[6] Positive significant relationship in all years for overall standard deviation ($p < 0.01$ in five of seven years and $p < 0.05$ in two), and highly significant for all years in test of idiosyncratic risk ($p < 0.01$).

firms and the control sample at the year in which the shocks to media employment are recorded are presented in Table 1, Panel A and Panel B, respectively. By design, mean market values and firm beta are nearly identical. Other statistics, including total assets and standard deviation of returns, are also similar at year 0. The firms also appear well matched in respect to the mean and median number of analysts following treated and control firms, and the R^2 values from the Fama-French Three Factor regressions is similar between both groups of firms.

Other variables

Annual beta and standard deviation estimates are calculated using daily price changes from the CRSP database. Daily trading volume data and spread calculations from the daily high ask and low bid are also taken from the CRSP database. Lagged monthly returns are calculated for each of the 12 individual months prior to the year being evaluated. Market value, total assets and other accounting information are taken from the Compustat database. Analyst coverage information is from the IBES database. Google Trends data is from publically available data provided by Alphabet, Inc., and downloaded using a script that queries http://trends.google.com/trends. Google Trends provides a public Search Volume Index (SVI) of searched terms. The SVI for a search term is a normalized number of a randomized sample of searches for that term on Google, scaled by its time-series average against all search traffic for a region.[7] Google search, which accounted for about 77 percent of all search traffic in the United States during the final year considered by Da, Engelberg, and Gao (2011), made up 88.23% of all U.S. search traffic in 2019, according to StatsCounter Global Stats estimates.[8]

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[7] Google Trends allows searches by various geographic regions, including cities, states, countries, and globally. SVI here measures search traffic in the United States.

[8] https://gs.statcounter.com/

As a measure of market power, I calculate a firm's concentration as the percentage of market share for each firm *i* in its six-digit GIC industry using the formula:

$$S_{i,t} = \frac{Sales_{i,t}}{Sales_{IND,t}}$$

Where S_i is firm *i*'s market share calculated as the *Sales* of firm *i* over total sales for all firms available in the Compustat database in the same 6-digit GICS category as firm *i* and with positive sales information.

To examine firms' competitive environment, I measure competition intensity with the widely used Herfindahl-Hirschman index:

$$HHI_{IND} = \sum_{i=1}^{N} S_i^2$$

Where *S* is the percentage of market share for each firm *i* in an industry, and HHI_{IND} equals the sum of the squared market shares of all firms in an industry listed in the Compustat database.

HHI has some limitations, as it only includes publically listed firms and therefore does not reflect the market share of private businesses. It also may imperfectly reflect state or local dynamics, in which a firm has near monopoly power in a region but does not have a significant share of the national market, or cases of collusion among firms. However, it is widely used by government agencies and in the finance literature (Giroud and Mueller, 2011; Hoberg, Phillips and Prabhala, 2014), and Pavic, Galetic and Piplica (2016) show that HHI is largely equivalent to other measures of firm concentration.

METHODOLOGY

Standard deviation and idiosyncratic risk

My primary analysis consists of a dynamic difference-in-differences model that measures the effect of shocks to media coverage on volatility of returns and the measure of idiosyncratic risk. The dynamic DiD model allows effects on the dependent variables to be monitored for several years after the shock, and my analysis examines six years after the shock is recorded. Because the actual reduction in media employment may have happened two years prior to being reported in the BLS data, I include the year before (year t = -1) the shock is recorded in my main reported regression results.[9]

After identifying a negative shock to media employment where a firm's headquarters is located as year t = 0, I collect market information and three-factor market model parameter estimates for as many years t from -6 to +6 as are available. Control firms are matched as described in Section 4 at t = 0 and corresponding information for years from t = -6 to t = +6 are matched to return and accounting information available from CRSP and Compustat.

The model (1) calculates differences between firms that have experienced the media employment shock and those that have not in each of the periods $t \geq -1$, as well as controlling for state- and year-specific fixed effects.

$$Risk_{\,i,t} = \sum_{k\geq 1} \beta_k \, Reduction\ in\ coverage\ has\ been\ in\ effect\ for\ k\ periods_{s,t} + \sum_{s} State\ fixed\ effects_s + \sum_{t} Time\ fixed\ effects_t + \beta_{i,t} Controls + \varepsilon_{i,s,t} \quad (1)$$

[9] In unreported results to confirm the parallel trends assumption of the difference-in-differences design, I test treatment effect at all years t = -6 through t = +6 for each of the tests where the dynamic DiD is employed. The first significant results appear no earlier than year t = -1.

Because the dependent variables are associated with market trading, I follow Levi and Welch (2017a, 2017b) who suggest that market value and lagged returns are the most important determinants in market model specifications. I use the natural log of the firm's market value to normalize the values and calculate the lagged returns of each of the 12 months in the prior year and use each month's total as a control. I also include controls that may affect the information environment of a firm, as compiled in Drake, Guest and Twedt (2014). Jensen and Meckling (1976) argue that analysts play a monitoring role in managerial performance, and Moyer, Chatfield and Sisneros (1989) and Chung and Jo (2016) find that analysts also keep markets informationally efficient by meeting information demands of investors, and Kelly and Ljungqvist (2012); find analysts can reduce information asymmetry. *Analyst Coverage* is a dummy variable equaling 1 if the firm has at least one estimate issued during the year (LaFond, 2005). The *Index dummy* variable equals 1 if a firm is listed on the S&P 1500 because being included on the high-profile index could both raise attention and scrutiny. I also include a firm's book-to-market ratio as a relevant element of an investors' value and estimation risk assessment (Desai, Rajgopal and Venkatachalam, 2004), and lagged beta to further control for the systematic elements of risk realized in the prior year.

Dynamic difference-in-differences regressions

All standard errors, in model (1) and all other regression models employed in this paper, are robust to heteroscedasticity and clustered by OES area number (the level of assignment of treatment) to address possible serial correlation in the error terms (Bertrand, Duflo and Mullainathan, 2004).

There are several threats to validity when using the difference-in-differences method. The first, conditional exogeneity, requires an assessment of the pre-treatment comparability of the treated and control samples. Figures 3 and 4, discussed fully in Section 6.1, demonstrate a similar pattern of standard deviation of returns among the treated and control sample firms prior to recorded media employment reductions, offering some assurance that the parallel trends assumption is reasonable (also, see Footnote 14). In unreported results, I also matched firms on levels of standard deviation in year t = -2 rather than market beta, and results in later differences between the two groups are similar to those in my main findings.

Employing the difference-in-differences design also requires the assumption of no systematic composition changes within nor between the treated and untreated groups (Blundell and Costa Dias, 2009). The assumption of noninterference has been noted in textbooks (Wooldridge, 2010) and numerous studies in economics (Angrist, Imbens and Rubin, 1996; Heckman, Lochner and Taber, 1998; Miguel and Kremer, 2004). This stable unit treatment value assumption, or SUTVA, requires that treatment applied to one unit does not affect the outcome for other units, either in the treated sample or in the control group.

In my setting, affected and unaffected firms are matched by industry, and therefore, are likely competitors. If treatment affects some firms negatively (positively), those effects may simultaneously benefit (hinder) competing firms within the same industry. In an analysis of competition's effect on idiosyncratic risk to firms, I find no evidence of such an effect; however, this is a potential concern of the research design.

Another element of the SUTVA concerns treated firms affecting one another. In my research design, treatment is assigned to firms at the metropolitan statistical area or other geographic region used by the OES, and although the model includes controls for state effects, as well as year effects, firm actions may affect other firms in the treated area. For example, one of the potential causes of increased estimation risk may be the reduced levels of monitoring after a reduction in local media employment (see Chapter 3). Parsons, Sulaeman, and Titman (2018) find that city-level norms may be a factor in financial misconduct. If reduced monitoring allows one treated firm to engage in financial misconduct, then norms may change for other firms in the immediate area and potentially increase financial misconduct among all treated firms. If market participants consider this risk as a reason to reassess their fair value estimates of securities, then estimation risk may increase. To address this within element of the SUTVA, I reduce my sample to only those firms with the highest market value in each area and their matching control firms. Only 94 treated firms and their controls are left in this sample, and after matching year t = -6 through +6 information from Compustat and CRSP, a total of 1,677 firm-years are evaluated. The reduced power from the smaller sample size leaves only two years post-treatment with a significant two-way positive correlations (years t = +1 and +2, $p < 0.1$) with standard deviation of returns; however, coefficients are directionally consistent and positive. To illustrate the effects of the reduced power from this analysis, the coefficient on the effect of *Analyst Coverage* remains negative, but is no longer significant, with absolute value of t statistics smaller than for any of the post-treatment year coefficients.

The parallel trends assumption and the Stable Unit Treatment Value Assumption are exclusion restrictions, and therefore cannot be confirmed using data or statistical inference. However, I have little evidence that the assumptions obviously are being violated.

Measuring attention: SVI

Da, Engelberg and Gao (2011) examine the SVI from Google search traffic for stock ticker symbols for all Russell 3000 stocks. They find that (log) SVI is positively correlated with alternative measures of attention such as extreme returns, turnover, and news. They determine that an increase in the weekly (log) SVI often precedes these alternative measures. They determine that SVI captures the attention of individual retail investors, who are likely less-sophisticated individual investors without access to specialized technology such as Reuters or Bloomberg terminals available to professional traders. Da, Engelberg, and Gao (2011) find that abnormal levels of interest often result in temporary increases in valuation. I download SVI ratings for each of my treated and matched control firms, as described in Data and Descriptive Statistics. I then calculate abnormal SVI (ASVI), which is the difference between the median SVI during a year and the average median SVI for the preceding three years; in regression analysis the ASVI is also log-transformed[10]. Da, Engelberg, and Gao (2011) use this measurement to capture unusual amounts of investor attention and they find that it is tied to abnormal returns and trading activity. Barber and Odean (2008) and Engelberg and Parsons (2011) find that increased attention is related to increased trading volume, and Barber and Odean (2008) find that retail investors are more likely to be net buyers of attention-grabbing stocks.

I report both overall levels of attention, as measured by SVI and ASVI, in Figure 6 and Figure 7. In regression analyses, I analyze whether a reduction in media employment at a firm's headquarters is correlated with overall levels of attention (log of 1+ SVI) and abnormal levels of attention (log of 1+ASVI), using the following OLS model:

(2)

$$Attention_{i,t} = \beta_1 Treat + \beta_2 Advertising + \beta_3 Treat * Advertising + \beta_4 No.Analysts + \beta_{i,t} Controls + \sum_s State\ fixed\ effects_s + \sum_t Time\ fixed\ effects_t + \sum_k Industry\ fixed\ effects_t + \varepsilon_{i,t}$$

[10] Da, Engelberg, and Gao (2011) calculate ASVI as the log-transformed difference between a firm's SVI one week and its median SVI for the eight weeks prior. Google only provides monthly SVI in the time frame I examine, and my analysis is annual, so I use a modified calculation of ASVI. I also analyzed ASVI with two-year-prior average median and one-year median. ASVI is essentially indistinguishable using these alternative methods.

Where the dependent variable, *Attention,* is log-transformed SVI or ASVI. *Treat* is a dummy variable equal to 1 if a firm at year t has experienced a reduction in media employment in the area where it is headquartered within the past six years. *Advertising* is advertising expenses reported by a firm scaled by its overall sales[11]. If firms offset the loss or reduction of one channel of information to customers and investors by boosting advertising expenses, then those increased advertising expenses may be positively correlated with attention levels. I include the interaction term *Treat*Advertising* to further test this relationship. In this model, I use the log-transformed number of analysts following a firm, calculated as the natural log of 1 + the number of analysts following the firm, taken from the IBES detail summary. Other controls include the other previously listed elements of a firm's information environment: natural log of market value, a dummy variable equal to 1 if a firm is part of the S&P 1500, firm's book-to-market, and returns for the previous 12 months. The model also includes state-, year-, and industry-fixed effects.

Measuring attention: Volume of shares traded

To further test whether reduced attention drives the increase in idiosyncratic risk for firms that experience a decline in local media, I examine overall trade volume over the pre- and post-treatment period. I return to the dynamic differences-in-differences model to test the relationship between a firm experiencing a shock to its area's media employment and the natural logarithm of total volume of shares traded during the year. The model:

(3)

$$Trades_{i,t} = \sum_{k\geq 1} \beta_k \, Reduction\ in\ coverage\ has\ been\ in\ effect\ for\ k\ periods_{s,t} + \sum_{s} State\ fixed\ effects_s + \sum_{t} Time\ fixed\ effects_t + \beta_{i,t} Controls + \varepsilon_{i,s,t}$$

[11] Missing values for advertising expenses are replaced with 0.

Trades is the number of shares traded annually as recorded in Compustat, and is log-transformed to normalize the data and eliminate the effect of extreme values. *Controls* include previously described variables intended to capture the information environment of the firm: firm size, a dummy variable indicating whether at least one analyst covers the firm, a dummy variable equal to 1 if a firm is part of the S&P 1500 index, and book-to-market. I also include prior 12-month returns. I include the reciprocal of the closing price of the security at year *t* as share price may influence the purchase and volume decision of some potential investors, particularly retail investors. I also include advertising expense scaled by total sales because Grullon, Kanatas, and Weston (2004) find that product market advertising can encourage investor activity.

Competition: Concentration and HHI

To assess whether a reduction in media employment near its headquarters has product market consequences for a firm, I examine whether a firm's market share, S_i, is related to a reduction in media employees in the area where a firm is headquartered. If local media provides another advertising or marketing channel for firms, then firms may use advertising to offset the loss of this channel. The OLS model (4) includes advertising expenses scaled by sales and an interaction term to test directly whether firms that experience a shock to media employment use advertising to offset the loss of that channel. The model is:

$$S_{i,t} = \beta_1 Treat + \beta_2 Advertising_{i,t} + \beta_3 Treat * Advertising + \beta_{i,t} Controls + \sum_s State\ fixed\ effects_s + \sum_t Time\ fixed\ effects_t + \sum_t Industry\ fixed\ effects_t\ \varepsilon_{i,t} \tag{4}$$

Where $S_{i,t}$ is market share as described in Data and Descriptive Statistics, *Treat* is a dummy variable equaling 1 if a firm has experienced a reduction in media employment and time t = -1 through 6, and $Advertising_{i,t}$ is a firm's advertising expenses scaled by the firm's total sales. *Controls* include the HHI of the firm's industry, the total number of companies with the same 6-digit GIC, the natural log of total employment in the firm's geographic area, firm market-to-book value, operating profit (EBIT) scaled by total sales, and the natural logarithm of total assets.

The HHI is included to indicate the potential for inter-firm efforts to maximize profits or conversely for a firm to exert control over pricing by exercising its market power (Rhoades, 1995). I also include the total number of firms within the industry, because for given levels of HHI, the number of firms can vary and therefore the competitive environment may not be entirely reflected in HHI. Area employment figures are included to reflect the available regional customer base, which is most likely to be affected by local advertising or local media coverage as suggested by Huberman (2001); Grullon, Kanatas, and Weston (2004); and Barber and Odean (2008). Market-to-book and operating profit are measures of performance, intended to capture firm-specific elements of managerial efficiency. Total assets and the market-to-book ratio reflect size and current valuation.

The U.S. Department of Justice considers a market with an HHI of less than 1,500 to be a competitive marketplace, an HHI of 1,500 to 2,500 to be moderately concentrated, and an HHI of 2,500 or greater to be a highly concentrated marketplace. According to the Justice Department, "as a general rule, mergers that increase the HHI by more than 200 points in highly concentrated markets raise antitrust concerns, as they are assumed to enhance market power under section 5.3 of the Horizontal Merger Guidelines jointly issued by the department and the Federal Trade Commission (FTC)." I follow the Justice Department guidelines to construct a subsample analysis of firms in competitive marketplaces, moderately competitive industries, and highly concentrated markets.

Estimation risk: Bid-ask spread

To assess the risk of adverse selection for market-makers (Copeland and Galai, 1983; Kyle, 1985; Lee, Mucklow and Ready, 1993; Easley and O'Hara, 2004), I measure the quoted bid-ask spread as:

$$Spread_{i,t} = \frac{1}{D}\sum_{1}^{D} High\ ask_{i,d} - Low\ bid_{i,d}$$

The spread for each year, *t,* is calculated as the average of the sum of the daily high ask price, $High\ ask_{i,d}$, minus the daily low bid price, $Low\ bid_{i,d}$, divided by the number of trading days, D, each year. The mean bid-ask spreads for treated and control firms from years t = -6 to +6 are shown in Figure 9. The regression uses the dynamic difference-in-differences design employed in other tests with a model of:

(5)

$$Spread_{\ i,t} = \sum_{k\geq 1} \beta_k\ Reduction\ in\ coverage\ has\ been\ in\ effect\ for\ k\ periods_{s,t} + \sum_{s} State\ fixed\ effects_s + \sum_{t} Time\ fixed\ effects_t + \beta_{i,t} Controls + \varepsilon_{i,s,t}$$

Controls include the log transformed price of the security, *Price,* taken at the end of the calendar year reported in Compustat, the information environment dummy variables *Analyst Coverage* and S&P 1500 *Index dummy*, and the average daily volume of shares trading during the year, *Trades.* Per-share price has been shown to be positively correlated with the bid-ask spread (Demsetz, 1968; Tinic, 1972; Tinic and West, 1972; Benston and Hagerman, 1974). In this paper's research setting, price may be a particular concern because higher per-share prices may discourage retail investors from entering the market for the security thereby improving liquidity for market-makers and informed traders. The information environment variables (dummy variables equaling

1 if any analysts cover the firm or if it is listed in the S&P 1500) are included to test the effects of other known information intermediaries. Lee, Mucklow, and Ready (1993) find that market-makers use both spread and depth to control their risk to adverse selection, so as an admittedly coarse control for depth, I include the average daily volume of shares traded as the variable *Trades.* Controlling for depth allows us to isolate the bid-ask spread as the information asymmetry element of liquidity. I expect a positive coefficient on *Price* as relatively informed traders will make up a higher proportion of the market for higher-priced securities and fewer noise traders will be available to improve price information and liquidity and reduce adverse-selection risk for market-makers. I expect negative coefficients on *Analyst Coverage* and *Index Dummy* as both sell-side analyst coverage and being listed on the S&P 1500 raise the profile of a security and imply a higher level of public information processing and monitoring. I make no prediction for *Trades* as both decreased (increased) depth and increased (decreased) spread are elements of decreased (increased) liquidity.

Estimation risk: Number of shareholders

To test the relationship among a reduction in media employment, the number of shareholders and the bid-ask spread, I employ the OLS regression model:

(6)

$$
\begin{aligned}
Spread_{i,t} = {} & Treat_{i} + Holders_{i,t} + Treat_{i} * Holders_{i,t} + \beta_{i,t} Controls \\
& + \sum_{s} State\ fixed\ effects_{s} + \sum_{t} Time\ fixed\ effects_{t} \\
& + \sum_{t} Industry\ fixed\ effects_{t} + \varepsilon_{i,s,t}
\end{aligned}
$$

Where $Spread_{i,t}$ and *Controls* are as described previously; $Treat_{i}$ is a dummy variable equaling 1 if the firm has experienced a reduction in media employment where it is headquartered, and $Holders_{i,t}$ is the number of common shareholders at the end of the year *t* from the CRSP database. My sample includes only firms (treated firms and their controls) from the period t = -1 to +3 to examine the near-term effects on the number of shareholders as it relates to reduced potential media coverage.

RESULTS

Standard deviation of returns and idiosyncratic risk.

Figures 3 and 4 show the mean and median standard deviations of returns for firms that experienced a 25% or more reduction in media employment in the areas where the companies are headquartered. Both figures demonstrate a clear relative increase among firms that experienced the shock to media employment after the reduction was recorded in year $t = 0$. Figures 5 shows the mean values of the root mean squared errors from Fama-French 3 Factor regressions of daily returns. The residuals can be interpreted as a direct measurement of the idiosyncratic risk of the firms, and the separation between shock and control firms again is clear from the figure, although the differences appear to begin in year $t = 0$, consistent with the reduction in media employment actually occurring up to two years before it is reflected in OES data.

The results from the dynamic difference-in-differences regression on standard deviation of returns, shown in the first column of Table 2, show significant positive correlations between a reduction in media employment and the standard deviation of returns beginning in year $t = -1$ ($p < 0.05$) and strongest in year t = +1 ($p < 0.01$). The significant correlations continue until year $t = +4$ ($p < 0.1$). The results from the residuals from the Fama-French Three Factor regressions, shown in the second column of Table 2, appear to affirm the correlation between idiosyncratic risk and a reduction in local media employment, with the strongest correlation at year t = +3 ($p < .01$) and continuing until at least year t = +6 ($p < .1$). Because one of the attributes firms were matched on was market beta, a measure of systematic risk, the results in Table 2 taken together suggest a reduction in media employment creates frictions in attention, the product marketplace, or value estimation that increases the idiosyncratic risk of firms. The signs on the variables intended to capture other information environment factors of firms are all significant ($p < 0.01$) and signed as theory suggests. *Analyst Coverage* and the S&P 500 *Index dummy* suggest reductions in standard deviation of returns and idiosyncratic risk, as does firm size as measured by *Market value.*

Having established a correlation between a reduction in nearby media employment and a firm's idiosyncratic risk, I explore the three channels by which I hypothesize the firms may be affected.

Attention: SVI, ASVI levels

Because Google data only extends back to 2004, the SVI analysis eliminates firms that experienced a shock to local media employment in 2003, so 2,092 firms (treated and matching controls) at year $t = 0$ are included in this analysis. My calculation of abnormal SVI requires three years of lagged SVI information; therefore, firms that experience a shock to media employment before 2007 are excluded from that analysis. That sample includes 1,462 firms at year $t = 0$. I first present figures showing overall median attention levels as measured by SVI (Figure 6) and ASVI (Figure 7) of firms that experienced a shock to media employment compared with control firms from six years before the shock is recorded to six years after. Figure 6 shows that median attention levels for firms that experience a shock stay steady before and after the reduction in media employment. Curiously, control firms appear to build attention, as measured by SVI, over the analyzed period. At six years before their matched firms experience a media employment reduction, control firms have a considerably lower median SVI, about 7 compared with more than

30 on the 100-point scale for "treated" firms[12]. SVI, however, *between* firms is not directly comparable, as an individual search term's SVI can only be compared to its own prior search traffic. However, the control sample firms, in aggregate, appear to be building attention over the first half of the analyzed time period, while treated firms are not. This does not appear to be related to shocks to media employment, however, as the trend among control firms is clear long before any reductions in local media employment occurred, and the control firm series levels off around the time of the shock to the treated sample. The differences in SVI among control firms and treated firms during the year t = -6 to 0 period may be the result of additional media coverage prior to year 0 among the treated sample and therefore reflect more attention earlier. However, because SVI is search-term specific, it may be the control sample that is benefitting from additional media coverage while starting at a similar level of attention in year t = -6. The differences may, of course, be unrelated to media coverage. What is demonstrable from Figure 6 is that there is no obvious *decline* in attention as measured by SVI among the firms that experience a shock to local media employment where they are headquarted.

Figure 7 presents a similar comparison of abnormal search attention (ASVI). The increasing levels of attention experienced by the matching control sample is reflected in relatively higher ASVI when compared with firms that experienced a media employment shock near their headquarters. Years t = -1 through t = +2 look especially striking, which would include the periods in which employment reductions occurred and immediately after. However, the numerical differences between the two group's median ASVI is relatively small, only 5 SVI at its most extreme (with a possible maximum of 100). Importantly in analyzing whether shocks to media

[12] Mean SVI indicates similar trends, with treated firms maintaining a fairly constant SVI and control firms' SVI increasing during the year -6 to year 0 period. Mean SVI, however, is consistently higher than median values, with treated firms around 40.

employment affect attention, the ASVI for firms that have experienced a shock to media employment exceed the control sample ASVI in years 4, 5, and 6.

If increased idiosyncratic risk among firms that have experienced a reduction in local media employment is the result of decreased levels of attention, then one would expect to observe differences in levels of SVI and ASVI after the shock to local media employment occurred. However, Figures 6 and 7 provide little evidence that the null for Hypothesis 2 can be rejected. The figures demonstrate that a dynamic difference-in-differences design is inappropriate because trends are clearly not parallel in the pre-treatment period. Instead, I use an OLS regression to measure the independent variables' correlation with firm levels of (log-transformed of 1+) SVI and ASVI.

Attention: SVI, ASVI regression results

The results of OLS regressions, in Table 3, provide further evidence that increased idiosyncratic risk cannot be attributed to declining levels of attention. The number of analysts, market value and whether a firm is included in the S&P 1500 index are significantly positively correlated with higher levels of attention, as measured by overall levels of SVI. This is consistent with firms that receive more information channels having higher-attention stocks. Somewhat surprisingly, advertising expenses appear to have a negative correlation with SVI, although the relationship is not significant. Using both SVI and ASVI as the dependent variable, a dummy variable indicating whether a firm has experienced a decline in local media employment within the past six years is insignificant for SVI and significantly *positive* for ASVI ($p < 0.01$). In other words, declines in local media employment are correlated with *increased* search traffic for firms headquartered in the same area. The abnormal SVI may be explained by investors turning to online sources for additional information about a firm as a substitute for local media coverage. And, as noted, absolute values of ASVI are relatively small. However, the result provides clear evidence that investors are not *less* aware, in the sense of the first condition of the Merton (1987) model, of the investment opportunities. Therefore, I cannot reject the null hypothesis that attention levels are unchanged after a shock to local media employment. Advertising expenses are positively correlated with ASVI ($p < 0.1$), and the number of analysts is significantly negatively correlated with abnormal attention ($p < 0.01$). Analyst coverage may provide a relatively steady information environment for investors and make firms less likely to suffer from the attention shocks measured by ASVI or a significant number of investors may use analyst forecasts as a substitute for general search queries. The S&P 1500 *Index dummy* is significant and positively correlated with ASVI, perhaps because changes to the index of large, popular and heavily monitored stocks encourage investors to seek out information about the index's component securities.

Attention: Trade volume

Figure 8 shows mean trade volume in years t = -6 to t = +6 for firms that experience a shock to media employment and their control firms. Both lines increase at similar slopes over the period, suggesting no apparent differences in the patterns of trading volume between the two sets of companies. The observed pattern also suggests the parallel trends assumption is not apparently violated for the period from year t = -6 to t = 0, providing support for use of the dynamic difference-in-differences model. The results from that regression test, in Table 4, suggest that the market value, information environment, price, and return controls explain a significant portion of the variance in trading volume (adjusted R^2 = 72.68%). I find only lightly significant relationships to some of the years after a reduction in local media employment affects the area where a firm is headquartered, and that relationship is positive. To the extent that a reduction in media employment affects trading volume, it appears to be correlated with an increase, although not a significant one. I again fail to reject the null hypothesis that a reduction in media employment reduces attention as measured by trading volume.

As mentioned in Hypothesis Development, HFTs make up an increasingly important source of trading activity. Despite the dynamic DiD model's assumption that only the differences between the two samples post-treatment and the pre- and post-treatment differences in the treated sample are being captured, it is possible increased algorithmic trading creates a confounding trend at a regional level unspecified in the model despite state- and year-fixed effects. For instance, HFTs may begin to target stocks in certain regions persistently over several years, thus inflating trading volume heterogeneously. To ameliorate this concern, I also control for linear state-specific time trends by allowing an interaction between each state dummy variable and each identification of *year t* from year *t* = -1 to *t* = +6. The regression model becomes:

(7)

$$Trades_{i,t} = \sum_{k \geq 1} \beta_k \, Reduction\ in\ coverage\ has\ been\ in\ effect\ for\ k\ periods_{s,t} + \sum_{s} State\ fixed\ effects_s + \sum_{t} Time\ fixed\ effects_t + [\sum_{s,t} State_s * Time_t] + \beta_{i,t} Controls + \varepsilon_{i,s,t}$$

Where all variables are the same as in model 3, with the addition of the *State*Time t* interaction that creates an explanatory linear trend.

I employ a third model examining trading volume that allows both linear state-time trends and quadratic state-specific time trends to capture the state-specific effects of HFT volume that may be nonlinear. The regression model is:

(8)

$$Trades_{\,i,t} = \sum_{k\geq 1} \beta_k\, Reduction\ in\ coverage\ has\ been\ in\ effect\ for\ k\ periods_{s,t} + \sum_{s} State\ fixed\ effects_s + \sum_{y} Year\ fixed\ effects_y + [\sum_{s} State_s * Time_t] + [\sum_{s,t} State_s * Time_t^{\,2}] + \beta_{i,t} Controls + \varepsilon_{i,s,t}$$

Where all variables are the same as in model 3, the *State*Time$_t$* interaction captures a potential linear trend and the *State*Time$_t^2$* allows for the effects of a quadratic trend.

The results for model 4 find a significantly positive relationship between trading and treatment at year $t = +2$ ($p < 0.1$) and the coefficients are positive until year 4, where it is negative but insignificant. The R^2 in the linear-trend model improves less than half a percentage point. Including the linear and nonlinear trends in model 8 yields an R^2 of 78.12%, an improvement of more than 5 percentage points, and while coefficients remain positive for years $t = -1$ to $t = +3$, as in model 4, no significant relationships are detected in model 8.

Increased competition: Market share

Next, I consider firms' competitive environment. If the loss of a channel of advertising and marketing is eliminated or reduced, firms' increased idiosyncratic risk may be related to declining market power or increased competition. For this analysis, I only consider firms from years $t = -1$ to $t = +3$ for two reasons.[13] The first is that if there are effects on market share because of reductions to a marketing channel, it is reasonable to assume those effects could be detected within three years of the reduction. Including years beyond $t = +3$ introduces the possibility of capturing unrelated effects on the longer timeline. The second reason is practical: Specific unemployment data before

[13] Regression results that include the entire sample from years -6 to +6 using interpolated estimates of area employment are indistinguishable from the -1 to +3 results.

2001 is unavailable, so allowing the years t = -6 to t = -2 creates several observations with missing independent variables.

Table 5 describes the changes to market share and industry HHI for firms in the treated and control sample. The three columns sort sample firm-years into those which increased their market share by more than 1%, stayed within +1% or -1% of last year's market share, or had decreased market share of more than 1%. The table also includes mean and median industry HHI for the sample.

Table 6 shows regression results for a firm's market share on the independent variables described previously. A reduction in media employment, levels of advertising expense, and the interaction between those two variables do not appear to be significantly related to a firm's market share. In fact, only the variables associated with firm size appear to correlate significantly with a firm's market power, as measured by share of industry sales. Therefore, I fail to reject the null hypothesis that a reduction in media employment does not have a significant effect on the product market share or competition levels.

Although the regression results indicate no significant relationship between a firm's market share and a reduction in media employment, the distribution of firms in Table 5 may indicate disparate treatment effects that are lost when concentration is examined in pooled OLS. In Table 5, more than 40% of the matched-sample control firms have similar market share to the year before, while only 7% of treated firms do. The treated firms with changes appear to be as equally likely to gain or lose market share relative to the rest of their industries, and the constraint of +1% or -1% are intentionally strict to approximate no change in market share. I examine subsamples using the Justice Department categories for competitiveness to determine whether the competitive environment plays a role in the increased idiosyncratic risks of the treated firms. None of the independent variables of interest are significant in any subsample. The adjusted R^2 of the regressions on the moderately competitive and concentrated subsamples are much higher (near 70%), implying the control variables are less effective in explaining market share in competitive environments.

Estimation risk: Bid-ask spread

Having failed to reject the null hypotheses that declines in attention or increases in competition are responsible for increased idiosyncratic risk of firms headquartered in areas with reductions in local media employment, I turn to the third possible channel: estimation risk. Market-makers are believed to optimize their positions by setting bid-ask prices that maximize profits from liquidity-motivated traders and protect against losses from information-motivated traders (Copeland and Galai, 1983; Kyle, 1985). And HFTs largely function as market-makers (Menkveld, 2013). Information traders would include those with inside information, such as managers. As the threat of adverse selection from informed traders increases, so does the bid-ask spread set by market-makers. Increases in bid-ask spread can be the result of more informed traders moving into the market for a security or the result of fewer liquidity, or noise, traders participating. Both situations result in a higher proportion of informed traders and fewer opportunities for market-makers to recover their information-driven losses (Glosten and Milgrom, 1985).

The dynamic difference-in-differences model suggests that in years $t = -1$ to $t = +6$ after a reduction in media employment is recorded in the area of a firm's headquarters, the bid-ask spread increases significantly (at $p < 0.1$ to $p < 0.01$) when compared with matching sample firms. Both elements, awareness and agreement on fair value estimates, of the Merton (1987) and Klein Bawa (1977) models imply this result. However, having found little evidence that awareness has declined, the increase in information asymmetry reflected in the bid-ask spread would appear to be most likely the result of increased estimation risk. The control variables, particularly those measuring the information environment, provide support for this conclusion. Both *Analyst Coverage* and the S&P 1500 *Index dummy* have significantly negative correlations ($p < 0.01$) with bid-ask spread, consistent with more uninformed traders participating in the market for the security and implying that estimation risk does not prevent them from participating. An increased share of noise traders allows market-makers to hedge their adverse-selection risk and keep bid-ask spreads relatively lower. *Price* has a significantly positive ($p < 0.01$) correlation with bid-ask spread, consistent with higher share prices limiting the role of noise traders in a security's market. *Trade,* which measures average daily volume traded and is used as a rough proxy for depth, also has a significantly positive ($p < 0.01$) correlation with the quoted spread.

Estimation risk: Number of shareholders

Our final test examines the interaction term between a reduction in local media employment, *Treat,* and the number of shareholders of a security, *Holders.* Both of the variables individually are insignificant; however, the interaction term is significantly positive ($p < 0.01$) and suggests that information asymmetry is increased for those firms with a relatively higher number of shareholders and headquartered in an area that has experienced a reduction in media employment. A higher number of shareholders suggests relatively fewer institutional investors and relatively more uninformed or retail investors. Therefore, the results are consistent with increased estimation risk from a local reduction in media employment most affecting retail investors.

Conclusion

As the number of people employed as members of the media continues to decline across the United States, the information environment for publicly traded firms may continue to be affected. My findings suggest that effects of local media reductions on firm risk are significant and persistent. After additional tests, the source of that idiosyncratic risk appears to be increased estimation risk among investors, rather than decreased awareness about the investment opportunities or real effects on firms' product market competition.

TABLES

Table 1.1, Panel A: Descriptive statistics of firms that experience shock to media employment

Summary statistics of firms identified with headquarters in Census Bureau statistical areas that recorded a 25% or more reduction in employment of Reporters and Correspondents or Brooadcast News Analysts over a two-year period. Values are those from the time when the shock is recorded (at time t = 0). Beta is calculated using the Fama-French Three Factor model.

N = 1,373	Mean	25th percentile	Median	75th percentile	St. Dev.
Market value*	$1,421.77	$136.50	$438.52	$1,242.82	$3,673.63
Total assets*	$3,247.46	$188.21	$644.5	$1,837.40	$16,737.21
3-Factor Beta	0.8377	0.5054	0.8987	1.1547	0.4674
St. dev. returns	2.6394%	1.8206%	2.4382%	3.2303%	0.0109
Mean spread	$0.3507	$0.1606	$0.2751	$0.4406	$0.3116
3-Factor R^2	24.39%	6.65%	22.15%	37.28%	18.83%
No. of analysts	6.30	0	4	10	7.05
Area employees	1,480,867	798,120	1,127,100	1,839,170	1,197,639

* in millions

Table 1.2, Panel B: Descriptive statistics of control firms

Summary statistics of firms matched with those identified with headquarters in areas that recorded a reduction in media employment over a two-year period. Firms are matched by year, market value, and beta from the Fama-French 3 factor. Values are from the time when the shock is recorded for the matched, treated firm (at time t = 0).

N = 1,373	Mean	25th percentile	Median	75th percentile	St. Dev.
Market value*	$1,434.40	$139.23	$423.50	$1,214.81	$4,599.65
Total assets*	$3,557.43	$198.42	$624.36	$1,705.63	$19,296.63
3-Factor Beta	0.8377	0.5027	0.8924	1.1625	0.4689
St. dev. returns	2.6182%	1.8301%	2.4337%	3.2507%	0.0108
Mean spread	$0.3370	$0.1567	$0.2608	$0.4135	$0.4274
3-Factor R^2	24.68%	6.59%	21.94%	38.00%	19.43%
No. of analysts	6.44	1	5	10	6.78
Area employees	975,782	143,925	450,890	1,433,880	1,186,725

* in millions

Table 1.2: Measures of risk

The results from the dynamic difference-in-differences with state and year fixed effects. Two measures of risk, overall standard deviation of log returns and the root mean squared error from Fama-French 3 Factor regressions, are the dependent variables. Control variables include the log of market value, a dummy variable indicating whether a firm is covered by analysts, a dummy variable indicating whether a firm is included in the S&P 1500, the log of the book-to-market ratio and lagged beta. Variables year t = -1 to + 6 show the correlation with a reduction in media employment where a firm is headquartered. Coefficients are significant at *** = p < 0.01; ** = p < 0.05; * = p < 0.10

Independent variables	Standard deviation of log returns	Mean-squared residuals from 3 Factor regression
Market value	-0.0029***	-0.0033***
(t stat)	(-24.74)	(-26.92)
Analyst Cover	-0.0007**	-0.0008***
	(-2.49)	(-3.15)
Index dummy	-0.0013***	-0.0014***
	(-4.20)	(-4.80)
Book-to-market	-0.0015***	-0.0014***
	(-6.76)	(-7.02)
Lag beta	0.0072***	0.0043***
	(39.62)	(22.73)
Year t = -1	0.0007**	0.0005*
	(2.20)	(1.84)
Year t = 0	0.0005*	0.0005*
	(1.90)	(1.76)
Year t = 1	0.0008***	0.0005*
	(3.36)	(1.95)
Year t = 2	0.0004	0.0003
	(1.52)	(1.24)
Year t = 3	0.0008**	0.0006***
	(2.53)	(2.71)
Year t = 4	0.0006*	0.0007**
	(1.65)	(2.30)
Year t = 5	0.0004	0.0006*
	(1.08)	(1.85)
Year t = 6	0.0006	0.0007*
	(1.23)	(1.61)
State and year fixed effects	Yes	Yes
Observations	24,621	24,621
R^2	93.98%	63.88%
Adjusted R^2	93.96%	63.75%

Table 1.3: SVI and ASVI

The results from OLS regression of the dependent variables natural log(1 + SVI) and natural log(1+ ASVI) on a dummy variable indicating whether the area where a firm is headquartered has experienced a 25% reduction in media employment, advertising expenditures scaled by total sales, and control variables that include natural log(1 + the number of Analysts covering a firm), a dummy variable variable indicating whether a firm is included in the S&P 1500, the log of the book-to-market ratio and prior year returns. Asterisks indicate coefficients are significant at *** = p < 0.01; ** = p < 0.05; * = p < 0.10

Independent variables	ln(1+SVI)	ln(1+ASVI)
Treat	0.0066	1.8109***
(t stat)	(0.13)	(3.06)
Advertising expenditures	-0.1173	3.0530*
	(-0.37)	(1.84)
Treat*advertising expenditures	0.2039	-2.8213
	(0.60)	(-1.64)
ln(1 + Number of analysts)	0.1753***	-0.5573***
	(5.50)	(-2.87)
ln(Market value)	0.0701***	-0.1267
	(3.12)	(-0.81)
S&P 1500 index dummy	0.0976**	3.2230***
	(1.98)	(6.35)
B2M	0.0095	0.2275
	(0.26)	(0.86)
Prior year's returns	0.1143***	0.5990**
	(2.64)	(1.99)
State, year, industry fixed effects	Yes	Yes
Observations	18,490	12,930
R^2	86.52%	13.26%
Adjusted R^2	86.42%	12.42%

Table 1.4: Trade volume

Dynamic difference-in-differences with state and year fixed effects. The dependent variable is the natural log of volume of shares traded. Control variables include reciprocal of share price at the end of the year, advertising expenses scaled by total sales, the log of market value, a dummy variable indicating whether a firm is covered by analysts, a dummy variable indicating whether a firm is included in the S&P 1500 index, the log of the book-to-market ratio, lagged beta, and prior year returns. Variables year t = -1 to +6 show the correlation of a reduction in media employment where a firm is headquartered with the outcome variable. Asterisks indicate coefficients are significant at *** = $p < 0.01$; ** = $p < 0.05$; * = $p < 0.10$sss

Independent variables	Ln(Annual volume of total shares traded)		
1/share price	0.8101***	Observations	24,636
(t stat)	(3.98)	R^2	72.77%
Ad expense/sales	0.6991***	Adjusted R^2	72.68%
	(2.90)		
Market value	0.7337***		
	(39.09)		
Analyst Coverage	0.1507***		
	(2.65)		
S&P Index dummy	0.3201***		
	(4.91)		
Book-to-market	-0.0934***		
	(-3.41)		
Lag beta	1.0323***		
	(28.88)		
Prior year returns	-0.5154***		
	(-10.38)		
Year t = -1	0.0536		
	(1.13)		
Year t = 0	0.0805*		
	(1.75)		
Year t = 1	0.1271**		
	(2.28)		
Year t = 2	0.1097*		
	(1.88)		
Year t = 3	0.0924		
	(1.50)		
Year t = 4	0.0483		
	(0.72)		
Year t = 5	0.0539		
	(0.76)		
Year t = 6	-0.0333		
	(-0.45)		
State and year fixed effects	Yes		

Table 1.5: Market share & HHI

Table 5 shows concentration increases and decreases for firms that experienced a decrease in media employment near their headquarters and a matching set of control firms. The columns indicate whether market share (percent of industry sales) increased by more than 1%, stayed relatively the same (between 1% increase and 1% decrease), or decreased by more than 1% percent.

	> 1% concentration increase	Same concentration (between +1% and -1%)	>1% concentration decrease
	Treated firm-years from time t = -1 to +3		
Number of firms	3,027	444	2,453
Percentage of firms in group	51.10%	7.49%	41.41%
Mean firm % of market	0.61%	0.76%	0.57%
Median firm % of market	0.13%	0.13%	0.09%
Mean Industry HHI	856	828	814
Median industry HHI	698	726	615
	Control firm-years from time t = -1 to +3		
Number of firms	1,863	2,372	1,427
Percentage of firms in group	32.90%	40.04%	25.20%
Mean firm % of market	0.75%	0.47%	0.65%
Median firm % of market	0.11%	0.09%	0.08%
Mean Industry HHI	881	795	856
Median industry HHI	731	586	669

Table 1.6: Market share regression results

The results from OLS regression of firm market share (% of total firm sales of total industry saless) on a dummy variable indicated whether media employment has declined in the area where a firm is headquartered, advertising expenses scaled by total sales, the HHI of the firm's industry, the total number of people employed in the area, and firm market-to-book, operating profit, and log of total assets. Asterisks indicate coefficients are significant at *** = p < 0.01; ** = p < 0.05; * = p < 0.10s

Independent variables	Market share % (total sample)	Market share % (competitive industries)	Market share % (moderate competition)	Market share % (concentrated)
Treat	0.0002	0.0002	0.0018	-0.0003
(t stat)	(0.44)	(0.29)	(0.75)	(-0.09)
Advertising expenditures	-0.0012	-0.0008	-0.0144	-0.0131
	(-0.24)	(-0.14)	(-0.84)	(-0.49)
Treat*advertising expenditures	0.0011	0.0010	-0.0453	-0.0127
	(0.22)	(0.18)	(-0.87)	(-0.40)
HHI of industry	0.0002	0.0038	-0.0683	-0.0103
	(0.02)	(0.23)	(-1.59)	(-0.17)
Number of firms in industry	0.0000	0.0000	-0.0002	-0.0001
	(-1.44)	(-1.29)	(-1.31)	(-0.78
Log (No. employed in area)	-0.0001	-0.0001	0.0003	-0.0020
	(-0.26)	(-0.17)	(0.35)	(-0.94)
Market-to-book	0.0002***	0.0002***	-0.0003	0.0006***
	(6.80)	(6.92)	(-0.52)	(1.14)
Operating profit margin	0.0000	0.0000	-0.0003	0.0015
	(1.21)	(1.20)	(-0.14)	(0.39)
Log of total assets	0.0049***	0.0048***	0.0072***	0.0057***
	(8.78)	(8.78)	(3.48)	(4.85)
State, year, industry FE	Yes	Yes	Yes	Yes
Observations	11,669	10,541	818	310
R^2	41.79%	37.61%	70.75%	76.71%
Adjusted R^2	41.13%	36.88%	67.88%	70.26%

Table 1.7: Bid-ask spread

Panel A	Panel B
The results from the dynamic differences-in-difference with state and year fixed effects. Bid-ask spread (annual mean of daily high bid minus low ask divided by 2) is the dependent variable. Control variables include the log of share price, a dummy variable equal to 1 if a firm is covered by analysts, a dummy variable equal to 1 if a firm is included in the S&P 1500, the natural log of total number of shares available, and prior year monthly returns. Variables year t = -1 to + 6 show the effect of a reduction in media employment where a firm is headquartered.	OLS regression for years t = -1 to +3 with state, year and industry fixed effects. Bid-ask spread (annual mean of daily high bid minus low ask divided by 2) is the dependent variable. Independent variables of interest are a dummy variable, *Treat* equal to 1 if the area of a firm's headquarters experienced a reduction in media employment and the log of number of shareholders. Control variables include log of share price, a dummy variable equal to 1 if a firm is covered by analysts, a dummy variable equal to 1 if a firm is included in the S&P 1500, and prior returns.

Independent variables	Bid-ask spread	Independent variables	Bid-ask spread
Price per share (log)	0.2724***	Treat	0.0045
(t stat)	(15.76)	(t stat)	(0.29)
Analyst Cover	-0.0560***	Holders	0.3528
	(-6.25)		(0.93)
Index dummy	-0.0760***	Treat * Holders	0.6626***
	(-6.21)		4.89
Mean daily volume	0.0265***	Price per share (log)	0.3215***
	(7.02)		(10.59)
Year t = -1	0.0141*	Analyst Cover	-0.0653***
	(1.83)		-5.52
Year t = 0	0.0114	Index dummy	-0.0597***
	(1.32)		-4.97
Year t = 1	0.0226*	Trades	0.0037
	(1.94)		0.76
Year t = 2	0.0259*		
	(1.83)		
Year t = 3	0.0362**		
	(1.99)		
Year t = 4	0.0453**		
	(2.06)		
Year t = 5	0.0344*		
	(1.67)		
Year t = 6	0.0541***		
	(2.76)		
Fixed effects	State, year	Fixed effects	State, year, industry
Observations	26,406	Observations	10,665
R^2	66.96%	R^2	68.73%
Adjusted R^2	66.85%	Adjusted R^2	68.31%

Asterisks indicate coefficients are significant at *** = p < 0.01; ** = p < 0.05; * = p < 0.10

FIGURES

Figure 1.1: Number of firms with negative shocks to media employment by year

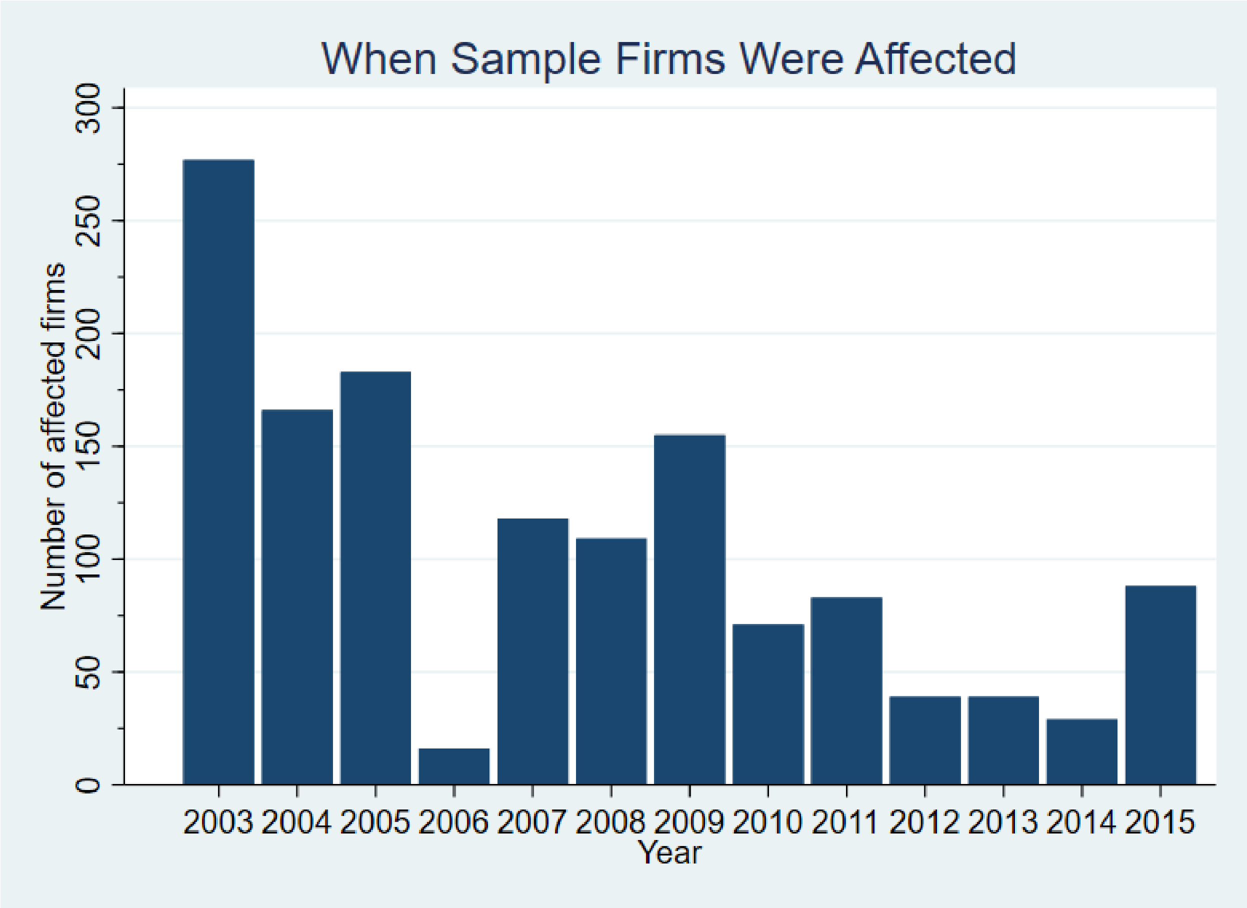

Figure 1. The frequency of the years in which the negative shock to employment of Reporters and Correspondents and Broadcast News Analysts occurred. From 2003 to 2015, 1,373 firms were identified as containing a shock to media employment in the geographic region in which they are headquartered.

Figure 1.2: Number of firms with negative shocks to media employment by state

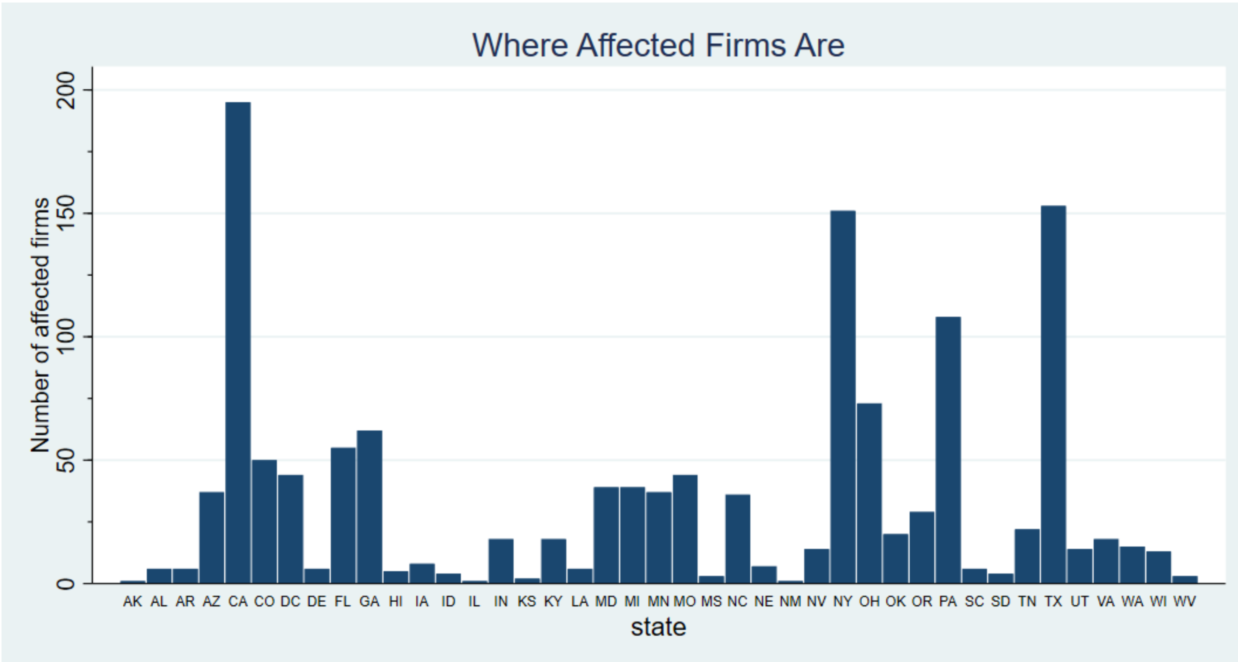

Figure 2. The frequency by state/territory in which the negative shock to employment of Reporters and Correspondents and Broadcast News Analysts occurred. From 2003 to 2015, 1,373 firms were identified as containing a shock to media employment in the Occupational Employment Statistical area in which they are headquartered. The OES areas spanned 41 states and the District of Columbia.

Figure 1.3: Mean Standard Deviation of Returns

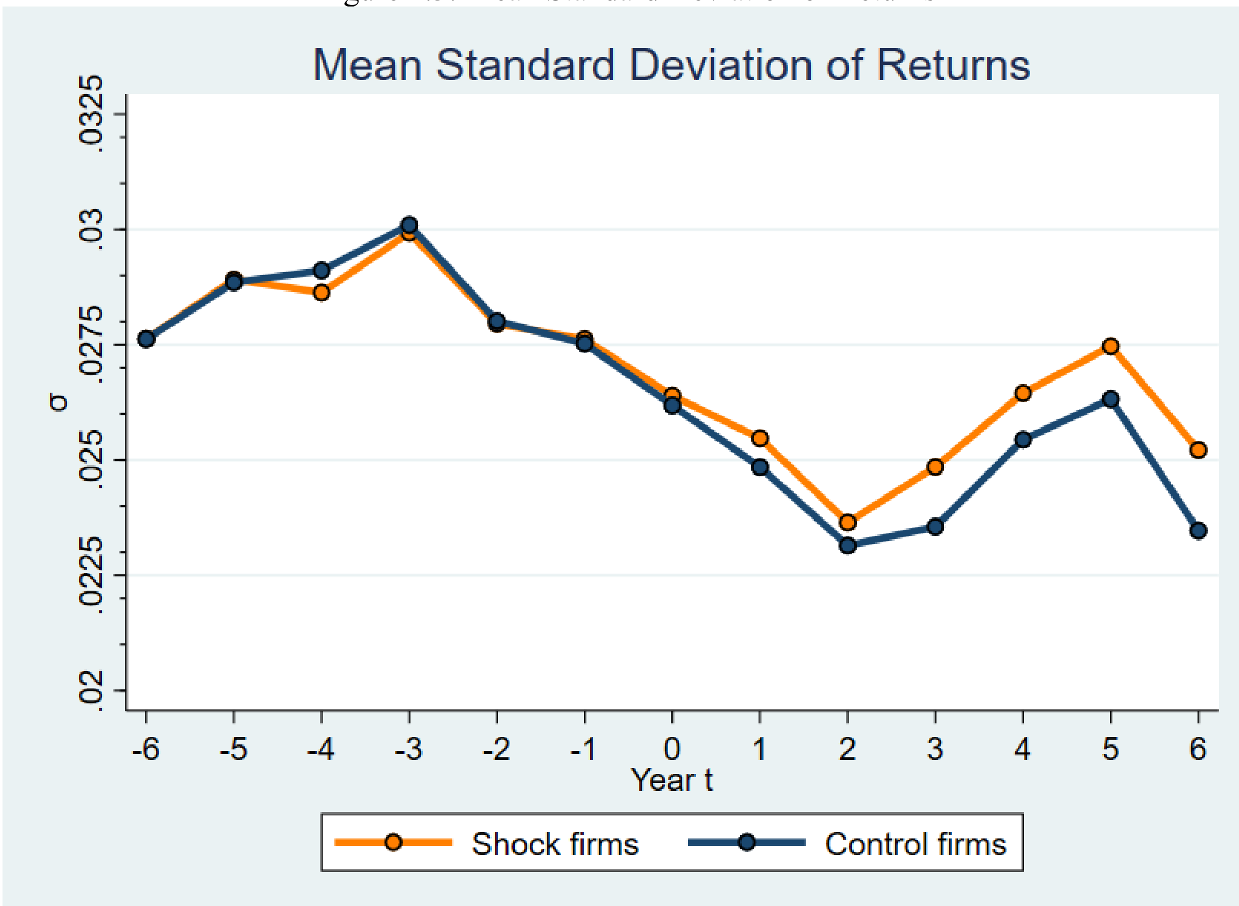

Figure 3: Mean standard deviation of log returns of firms that experienced a reduction in media employment in the area where the firm is headquartered, recorded at year t = 0, and a matching control sample of firms.

Figure 1.4: Median Standard Deviation of Returns

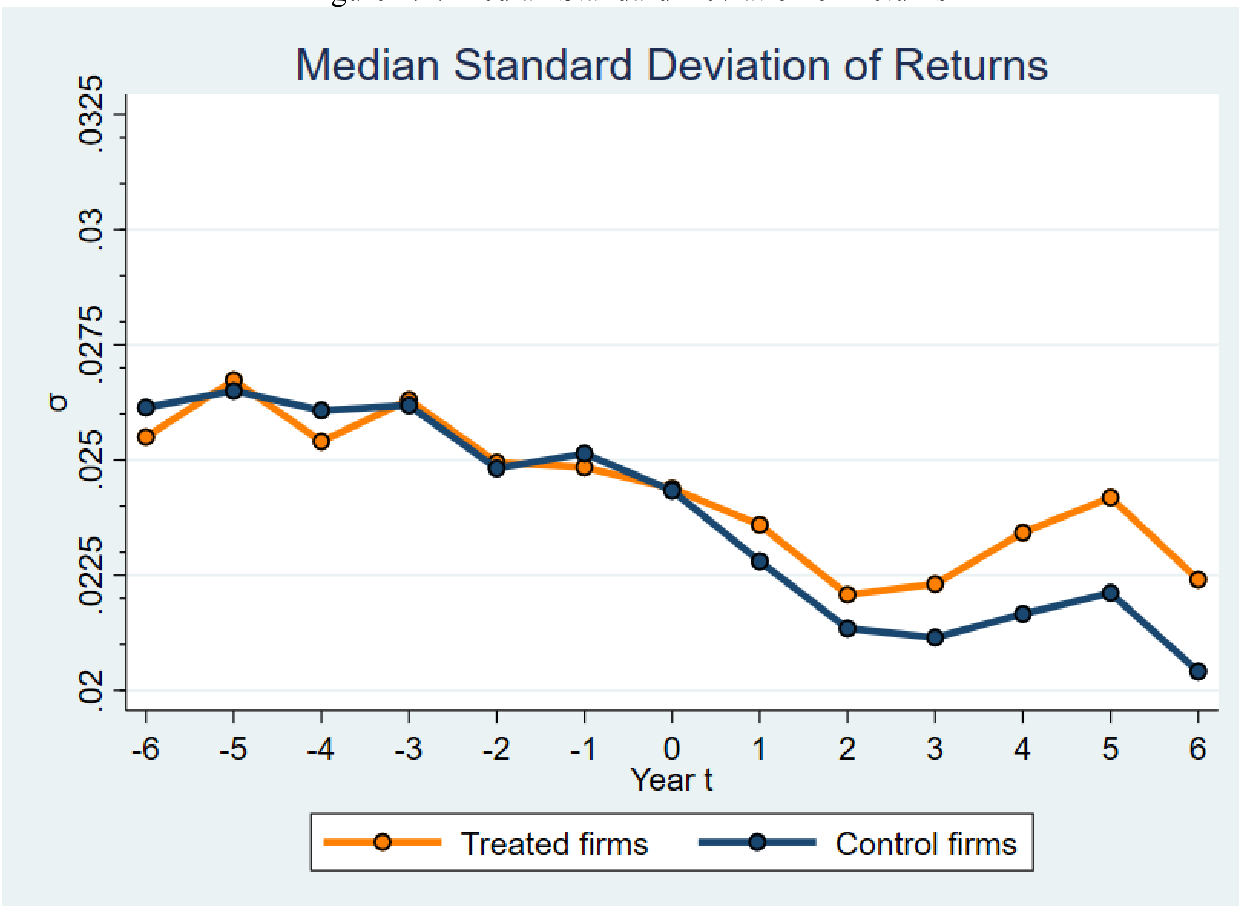

Figure 4: Median standard deviation of log returns of firms that experienced a reduction in media employment in the area where the firm is headquartered, recorded at year t = 0, and a matching control sample of firms.

Figure 1.5: Residuals from Fama French 3 Factor model

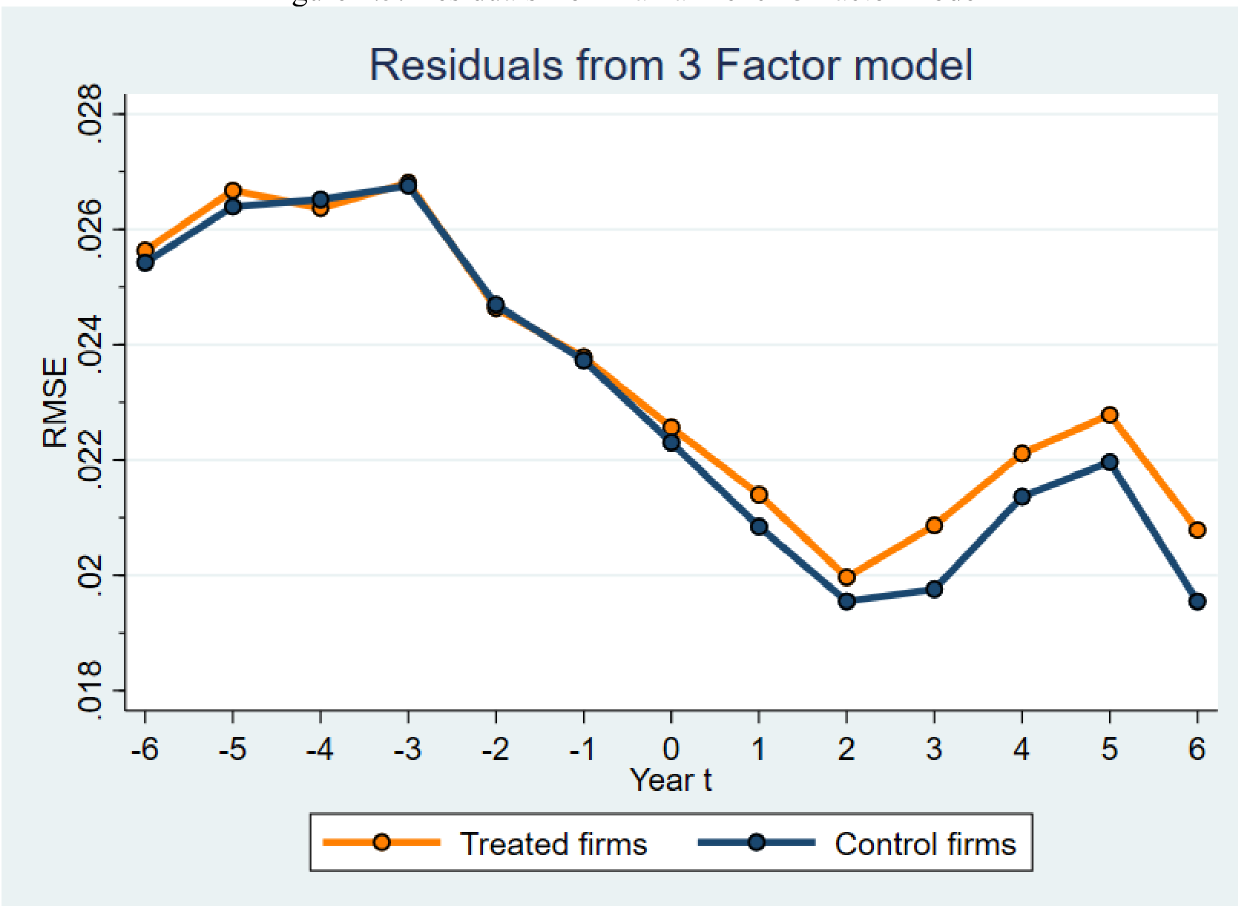

Figure 5: Mean residuals from Fama-French 3 Factor regressions for firms that experienced a reduction in media employment in the area where the firm is headquartered, recorded at year t = 0, and a matching control sample of firms.

Figure 1.6: Average Median Annual Search Volume Index

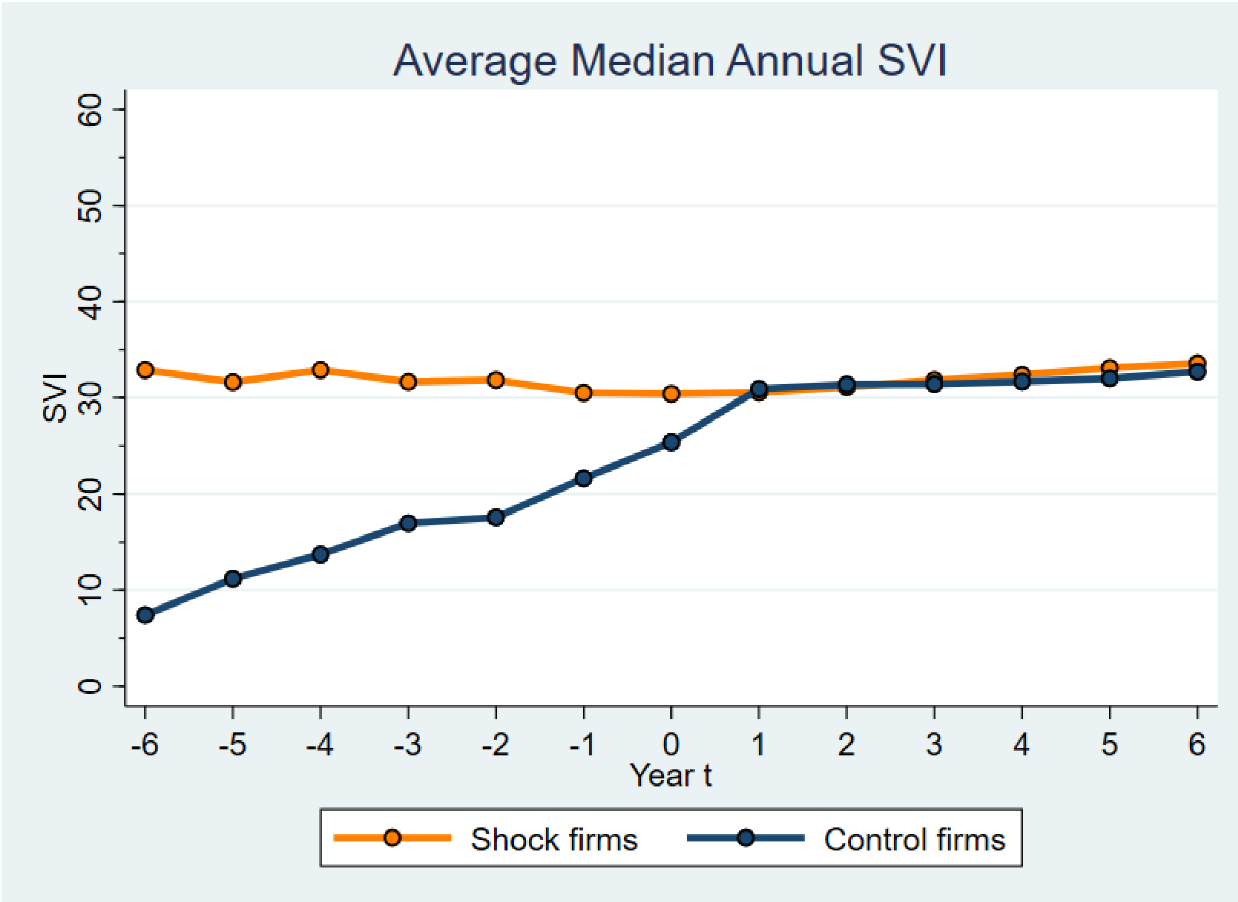

Figure 6: Overall levels of Google search traffic (SVI) for the ticker symbol of firms that experienced a reduction in media employment in the area where the firm is headquartered, recorded at year t = 0, and the ticker symbols for a matching control sample of firms. SVI is measured on a scale of 0 to 100 and compares a search term's popularity to total Google search traffic for the period and region examined and to itself over time. SVI is measured by year in the United States.

Figure 1.7: Abnormal Search Volume Index

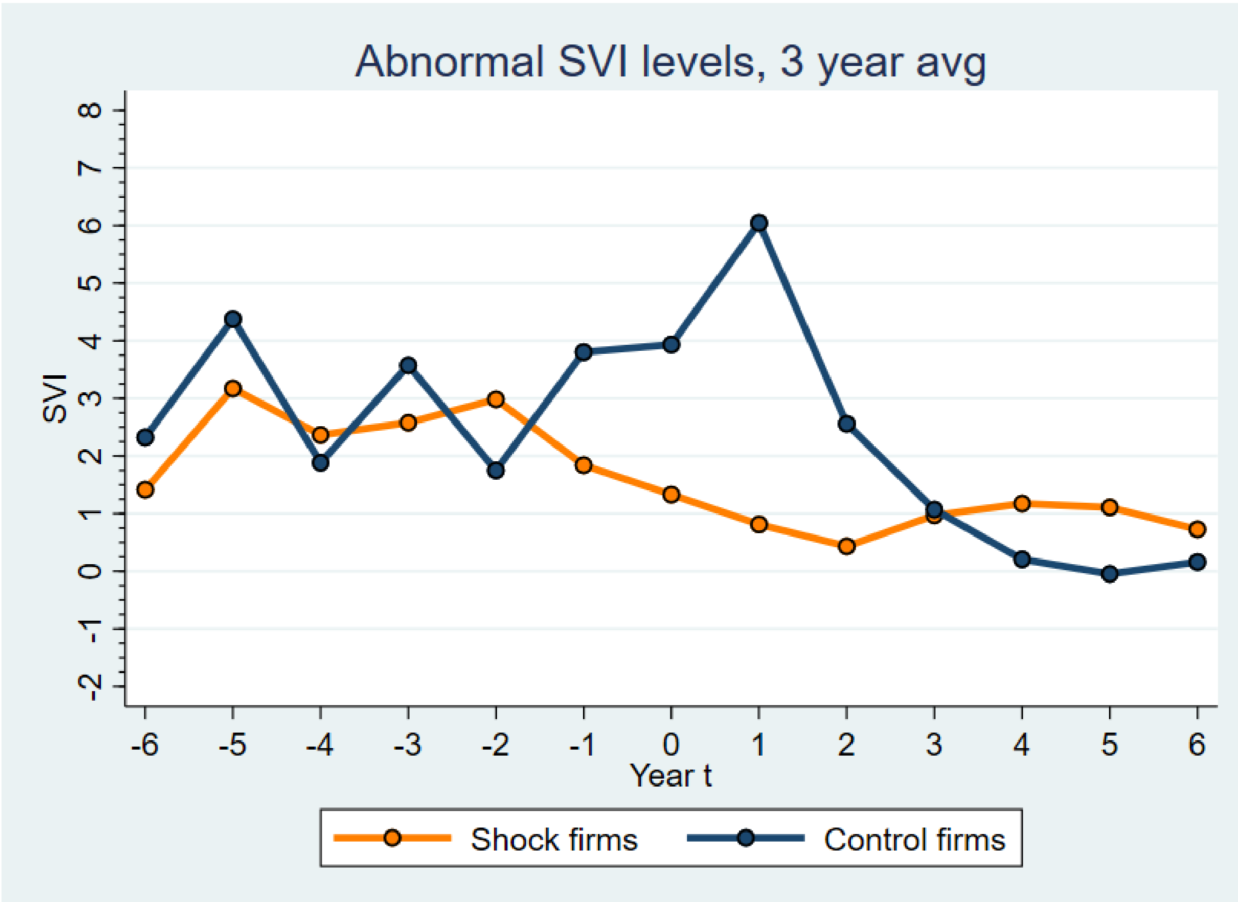

Figure 7: Abnormal levels of Google search traffic (ASVI) for the ticker symbols of firms that experienced a reduction in media employment in the area where the firm is headquartered, recorded at year t = 0, and the ticker symbols of a matched control sample of firms. ASVI compares the SVI of the ticker symbol at year *t* to its median search traffic for the period t = -3 to -1. SVI is measured by year in the United States.

Figure 1.8: Median of Total Trade Volume of firms' stocks

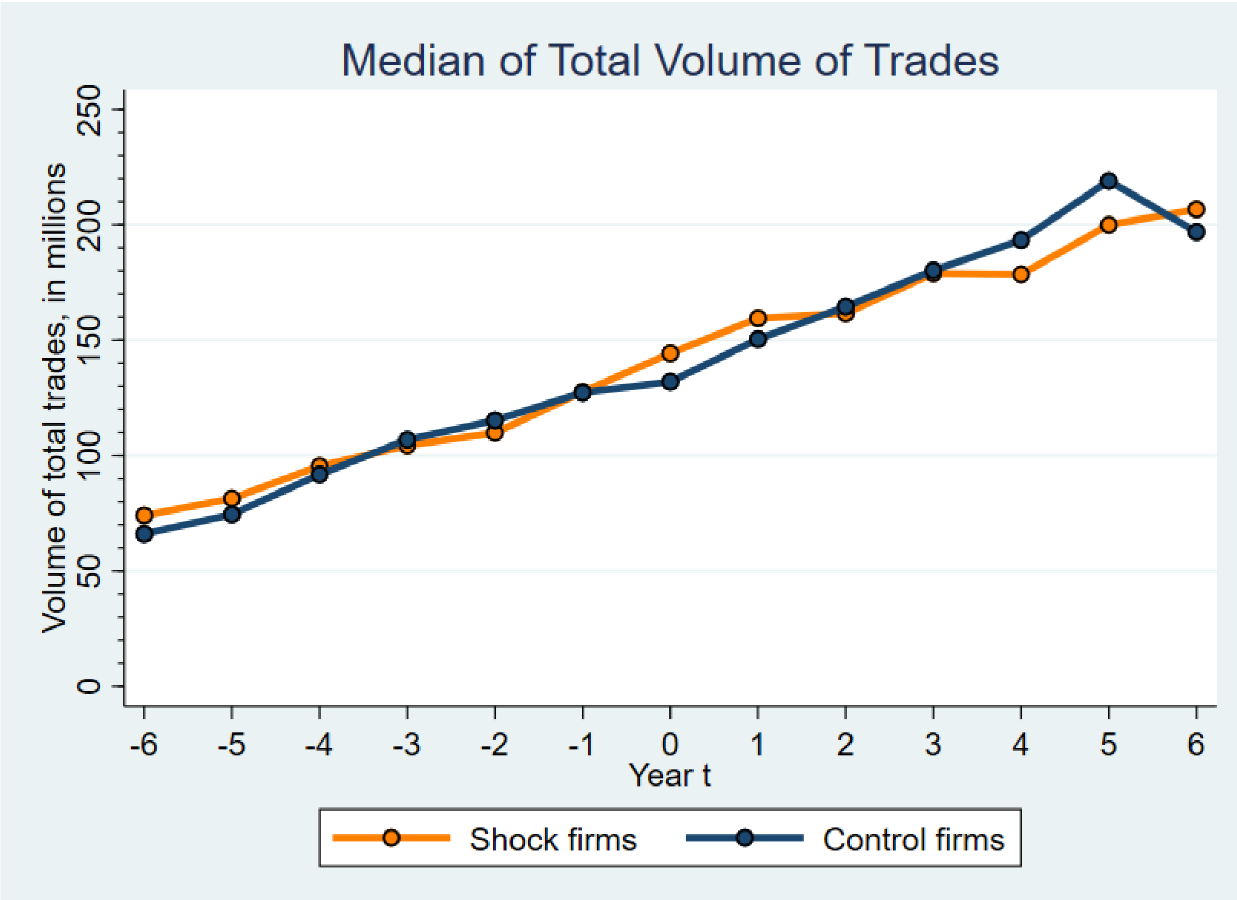

Figure 8: Trading volume of shares among firms that experienced a reduction in local media employment nearby and control firms.

Figure 1.9: Mean Bid-Ask Spread for Firms' Stock

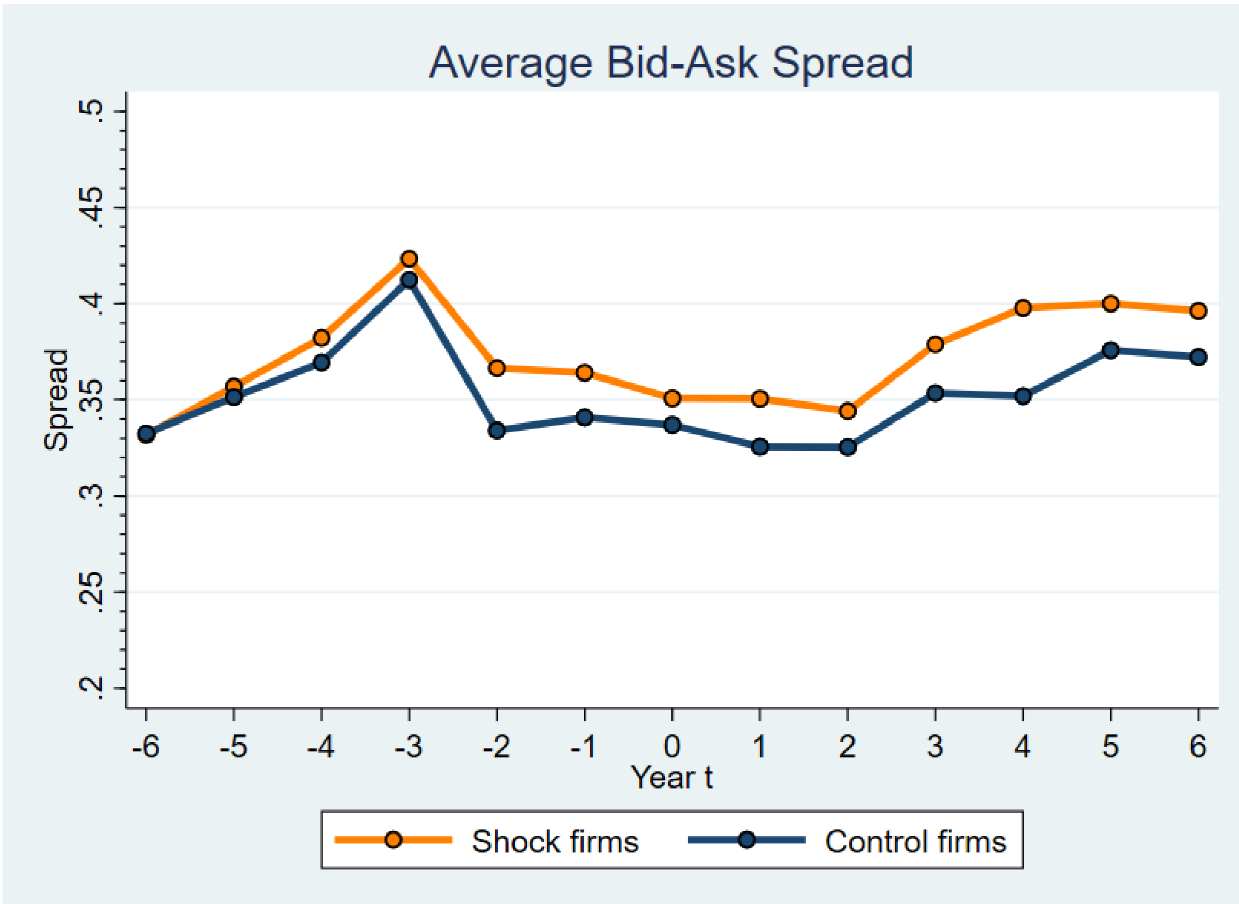

Figure 9: Average bid-ask spread among firms that experienced a reduction in local media nearby and control firms.

Chapter 2: Effects of local media reductions on returns, valuation, and debt

INTRODUCTION

Theory predicts that as investor attention declines or investor estimation risk increases (Klein and Bawa, 1977; Merton, 1987), securities will deliver higher levels of returns. Consistent with that, prior research has found that a lack of media coverage is associated with higher returns (Huberman and Regev, 2001; Chan, 2003; Fang and Peress, 2009; Antweiler and Frank, 2011). However, in the years following a reduction in nearby local media employment, I find that firms appear to not only fail to deliver higher levels of returns but to actually have lower returns than a set of peer firms that have not been affected by a reduction in media employment. However, affected firms do appear to be more likely to issue debt and to pay a debt premium from lenders. The effect on loan spreads is similar to the effect found for municipal governments in Gao, Lee and Murphy (2019) in the municipal bond market after a newspaper in the county closes.

This paper begins by identifying whether the stock return effects of local levels of media employment are similar to the effects of levels of "news coverage" as defined by Fang and Peress (2009). I find similar return premiums for firms headquartered in areas with relatively lower number of media employees when compared with those with a higher number of media employees. The return premium is similar in magnitude to that found in Fang and Peress (2009) for securities that received no coverage in four major national newspapers each month compared with those identified as "high coverage," i.e. being the subject of the highest number of articles in those publications. As in Fang and Peress (2009), I construct long-short portolios of low "coverage" firms and high "coverage" firms, where "coverage" in my setting is defined as the number of media employees in the Bureau of Labor Statistics Occupational Employment Statistical area where the firm is headquartered. I assume that those journalists employed nearby the firm's headquarters are most able and available to cover a firm (Engelberg and Parsons, 2011; Waldman, 2011; Gurun and Butler, 2012; Call *et al.*, 2018; Cage, Herve and Viaud, 2019). I find statistically and economically significant differences in overall returns and alpha from the market model and factor index models, consistent with the findings in Fang and Peress (2009) where "coverage" is defined by number of articles.

Next, I construct a sample of firms that have experienced a 25% reduction in media employment in the areas where they are headquartered to determine whether they experience the implied required increase in returns. I find evidence of the opposite in the six years after the reduction is recorded. In tests of valuation, market-to-book value is significantly positively correlated with a decrease in local media employment. Both findings are contrary to economic theory about investor attention and firm information environment and dissemination, which suggest that investors demand a return premium in less "covered" firms and that valuation will decrease as fewer investors participate in the market for a security (Klein and Bawa, 1977; Grossman and Stiglitz, 1980; Merton, 1987). In *Essay 1,* I find that the idiosyncratic risk of the sample firms increases after a reduction in local media employment and an increased estimation risk among investors is the likely reason. Theoretically, increased estimation risk has the same consequences as decreased attention, and implies lower valuations and higher required returns (Klein and Bawa, 1977).

Although I do not find a relationship between reduced media employment and returns, reductions in local media employment do appear to be correlated with a change in the propensity to issue debt and the spreads of those loans. Firms headquartered in areas that experience such a reduction appear more likely to issue debt, and private lenders require higher spreads from them, even after controlling for prior leverage levels. These findings are robust to controlling for firm characteristics, as well as other information environment factors and performance of the firm.

MOTIVATION

Most prior literature on media coverage and capital markets measures the numbers of articles appearing about a firm[14] or the number of television segments[15]. In representative examples, Fang and Peress (2009) examine stories that appeared in four major national newspapers; Engelberg, Sasseville and Williams (2012) count segments of *Mad Money* on CNBC; Twedt (2016) counts the number of articles transmitted by the Dow Jones Newswires; and Fedyk (2018) examines headlines published on the front page of Bloomberg News Terminals. However, my research attempts to explore the effect of news that is not produced or published. Similar in spirit to Gao, Lee and Murphy (2019) who examine newspaper closures, I explore whether a reduction in the number of media employees in the area where a firm is headquartered affects companies' financing costs. Gao, Lee, and Murphy (2019) explore municipal financing and find that the threat of government inefficiencies and informational frictions after a newspaper closure leads lenders to demand higher yields. In the corporate finance realm, Antweiler and Frank (2011) find that returns around firm-generated news are affected by whether *the Wall Street Journal* reports on it, and Bushee *et al.* (2010) and Bonsall, Green, and Muller (2019) suggest that less media coverage can increase information asymmetry and delay price discovery among both retail and institutional investors. Bradshaw, Wang, and Zhou (2019) find that the number of news articles

[14] Examples of research that examines print and online article publications include Pritamani and Singal (2001); Dyck and Zingales (2002); Chan (2003); Fang and Peress (2009); Engelberg and Parsons (2011); Ahern and Sosyura (2014); Fedyk and Hodson (2014); Peress (2016); Blankespoor, DeHaan and Zhu (2018); and Fedyk (2018)

[15] Research focusing on television segments include Engelberg, Sasseville and Williams (2012); Aman, Kasuga and Moriyasu (2018); Baloria and Heese (2018); and Peress and Schmidt, 2020).

about a firm are positively associated with analyst revisions, and that the tone of those articles predicts the direction of revisions. Guest and Kim (2019) also find that analysts and media play a complementary role, with each reducing the costs of generating information about covered firms for the other. Kothari, Li, and Short (2009) find that disclosures by business press have significant effects on firm cost of capital and return volatility.

Media can also play a role in firm decisions, as they attempt to manage their reputation capital (Dyck, Volchkova and Zingales, 2008; Dyck, Morse and Zingales, 2010; Liu and McConnell, 2013; Baloria and Heese, 2018), suggesting that firms with less potential for media coverage may act differently than those subject to higher levels of media scrutiny. The lack of media coverage may also influence investor behavior. Blankespoor, DeHaan, and Zhu (2018) find that the appearance of articles produced by an algorithm and containing no information beyond firm press releases increase trading activity and improve liquidity. Barber and Odean (2008) and Engelberg and Parsons (2011) find evidence of increased local trading when local media cover nearby firms.

Collectively, those findings suggest that differences in levels of media coverage, irrespective of actual articles, may be have similar effects to those of firm-level commitments to disclosure. Healy and Palepu (2001) count the press, along with analysts, auditors, and others, among the information intermediaries who can enhance the credibility of firm-produced financial disclosures and help reduce the agency problem between investors and managers seeking funding (Akerlof, 1970; Jensen and Meckling, 1976). Consistent with the predictions of the estimation risk hypothesis of Klein and Bawa (1977) and the implications of the Merton (1987) attention hypothesis, firms with higher levels of disclosure have been found to have lower costs of capital than those firms with lower levels of disclosure and thus higher information risk (Healy, Hutton and Palepu, 1999; Verrecchia, 2001; Botosan and Plumlee, 2002; Gelb and Zarowin, 2002). Leuz and Wysocki (2016) model a link between firm information quality and cost of capital and find that it affects investment decisions and expected cash flows. And Miller and Skinner (2015) suggest that media coverage influences management's disclosure decisions. Significantly, Miller (2006) shows that the media identifies almost one-third of fraud cases before they are announced by the firm, suggesting investors and lenders can, to an extent, rely on media to encourage disclosure. In the absence of outside information intermediaries, as Kothari, Shu and Wysocki (2009) find, managers on average delay disclosure of bad news relative to good news, which has implications for shareholders and lenders.

HYPOTHESIS DEVELOPMENT

Klein and Bawa (1977) and Merton (1987) propose models in which investor attention has implications for returns. Theoretically, investors are able to diversify away all idiosyncratic risk by using the essentially infinite number of assets available to their portfolios; however, actual investor portfolios are functionally limited to the assets with which investors are aware and for which they can arrive at fair value assessments. Merton models "awareness" as both knowing about a firm, i.e., that it exists an investment opportunity, and being able to properly assess parameters that affect its operation, i.e. estimate its value. If investors are "unaware" of an asset, either because they do not know that an investment opportunity for a firm exists or because they cannot properly estimate its value, then it cannot be included in their portfolios. In Merton's model, firms have an optimal number of investors and limits to attention can prevent firms from reaching that optimal number. For firms with a suboptimal number of shareholders, investors will require higher returns because of the additional risk carried. The Merton model also predicts that as awareness increases and more investors buy securities in the firm, the value of the company will increase.

Merton (1987) includes estimated valuations as an element of attention, but only includes one level of information in his model. Klein and Bawa (1977) explicitly discuss the implications of estimation risk among investors and suggest the effects would be similar to declining attention, with risk-averse investors moving their investments away from assets with difficult-to-assess fair values. The effect, as with declining attention, is reduced valuations and increased required returns to the remaining investors, which is also consistent with increased information risk in the disclosure literature (Healy, Hutton and Palepu, 1999; Verrecchia, 2001). Lehavy and Sloan (2008) find that decreased investor recognition, as determined by 13F filings, is linked to higher returns, and dissemination and recognition appear to be even more important than earnings news in explaining stock returns. Li (2015) finds that the most price-informative articles are produced by journalists who rely most heavily on first-hand access to management, institutional investors, and other experts. This suggests that local media employees may be uniquely positioned, by geography, to reveal news about a nearby firm. Gurun and Butler (2012) consider this role of journalists, but empirically find that local media are subject to "hype" when covering firms headquartered nearby and tend to cover local firms with a positive slant because of the influence of advertising relationships. In a survey of journalists, which includes several local and regional news outlets in addition to national and financial press, Call *et al.* (2018) find that journalists acknowledge pressure from management to avoid unfavorable stories, but that they consider monitoring companies one of journalism's most important functions.

My first hypothesis considers whether local media employees have a similar impact on returns as the well-established link between disclosure generally and the national and business press specifically.

Hypothesis 1: Firms headquartered in areas with relatively fewer local media employees will deliver higher returns than those firms headquartered in areas with relatively more local media employees.

A reduction in media does not appear to produce a reduction in measures of attention; however, it does appear to be associated with increased information asymmetry between investors and the firm, and investors face higher levels of idiosyncratic risk (*Essay 1*). If those are priced elements of valuation and returns, then returns should increase as valuation decreases.

Hypothesis 2a: Firms that experience a reduction in local media employment will deliver higher returns and suffer valuation declines relative to a sample of control firms.

Idiosyncratic risk; however, is diversifiable and therefore may not be reflected in lower valuations or higher returns. Hughes, Liu and Liu (2007) find no information asymmetry effect on risk premiums, suggesting rather that the cost of equity is driven by betas and factor risk premiums. Niessner and So (2018) find that financial news is more likely to focus on negative news, and suggest that findings about abnormal returns around high and low coverage are the results of high coverage firms experiencing negative returns rather than low coverage firms experiencing abnormally positive returns[16]. And Maskara and Mullineaux (2011) find that firms can publicize debt issuances to maintain valuations, essentially using lenders' presumed access to proprietary information to certify that a borrowing company is a safe investment. Therefore, I also state the null hypothesis:

Hypothesis 2_0: Firms that experience a reduction in local media employment maintain consistent or lower returns and consistent or higher valuations than a sample of control firms.

Jensen and Meckling (1976) outline agency problems between both managers and equity holders, as well as agents and debt holders. Leland and Pyle (1977) also detail some of the information risks faced by lenders. Smith and Warner (1979); Bushman and Smith (2001); Dichev and Skinner (2002); Bharath, Sunder and Sunder (2008); Zhang (2008); Nikolaev (2010); and Kim, Song and Zhang (2011), among others, demonstrate that lenders use the spread of loans and covenants to mitigate the risks associated with the information asymmetry between lenders and borrowers. Although little research has examined the role of the financial press in debt contracting, several papers discuss the role of media and the auditing function, another information intermediary that can certify a firm's disclosure (Healy and Palepu, 2001). Mutchler, Hopwood and McKeown (1997) and Joe (2003) find that media coverage can influence auditors' going-concern decision, and Gong, Gul and Shan (2018) find auditors charge higher fees for firms that are the subject of higher levels of media coverage. Collectively, the findings suggest that auditors may recognize an increased likelihood of disclosure of bad news that threatens the validity of their audit opinions and increases their litigation risk. Importantly, increased media coverage is unlikely

[16] This is in explicit contrast to Fang and Peress (2009), who find that increased returns to a portfolio long in No coverage firms and short in High coverage firms are driven by the higher returns to the No coverage firms.

to be associated with the existence of bad news; the risk, instead, is the result of media coverage increasing the likelihood that extant bad news is revealed. Debt holders have incentives to discover bad news about borrowers, but no agent can monitor all the events that are potentially relevant to their decisions (Nimark and Pitschner, 2019). If lenders face decreased channels for such discovery, they may charge a premium for their increased information risk.

Hypothesis 3: Firms that experience a reduction in local media employment face higher interest rates (spreads) from lenders than a sample of control firms that also issue debt.

DATA AND DESCRIPTIVE STATISTICS

BLS data

Information on the number of reporters and correspondents in an area is taken from the Bureau of Labor Statistics' Occupational Employment Statistics program. The survey produces annual estimates of employment for 810 specific occupations in more than 580 areas, including metropolitan statistical areas (MSAs) and nonmetropolitan areas throughout the U.S. states, the District of Columbia, and U.S. territories. The statistics are compiled annually after each year's May survey and are released in March of the following year. The Metropolitan Statistical Areas used by the BLS are designated by the U.S. Census Bureau. I link ZIP codes to MSAs using Census Bureau designations for 2010, and identify the ZIP code of a firm's headquarters using Compustat listings for its mailing address, because many companies incorporate outside of the state where their operational headquarters may be located for tax and governance reasons.

I use the total numbers of employees in the Reporter and Correspondent categories, which include both the Publishing Industries, such as newspapers and periodicals, and the Radio and

Television Broadcasting Industries. The employment total used in my analysis includes Broadcast News Analysts, which are listed in one category with Reporters and Correspondents in some years.

Low-High coverage portfolios

I construct an equally weighted portfolio of firms, similar to the portfolios constructed in Fang and Peress (2009). In their sample, firms are identified as No, Low, and High coverage firms, based on articles that appeared in four major newspapers in the United States. Those newspapers are the *Wall Street Journal,* the *New York Times,* the *Washington Post,* and *USA Today.* The majority of their firms fell into the No coverage category, with no stories appearing. Those firms delivered significantly higher returns than the High Coverage firms, and the authors attribute the differences to the higher returns required by investors to invest in lower-attention and less-liquid securities. To assess whether local media employment produces a similar effect, I construct portfolios based on terciles of the overall number of media employees in the geographic area where a firm is headquartered. Firms in the lowest tercile are identified as Low coverage firms, and those in the highest tercile are identified as High coverage. Because of the BLS data reporting schedule, my portfolios are rebalanced annually. To generate comparable annual returns, I take the monthly return and multiply it by the 12 months of the year[17]. I then repeat a subsample analysis included in Fang and Peress (2009). To demonstrate that effects are not driven by other firm characteristics, I separate firms into portfolios based on terciles by firm size, book-to-market, and share price, and analyze returns to each of those portfolios. Following Fang and Peress (2009), I eliminate firms with a closing share price for the year under $5.

[17] If I assume returns are compounded monthly, the returns to each portfolio are implausibly high, although directionally consistent.

Following Fang and Peress (2009), I regress daily returns for an equally weighted long-short portfolio of Low Coverage-High Coverage firms on the market return using the Capital Asset Pricing Model, the Fama-French (1993) Three-Factor model, and the Carhart (1997) Four-Factor model. The full four-factor model is:

$$R_{p,t} - R_{f,t} = \alpha_p + \beta_p^M\left(R_t^M - R_{f,t}\right) + \beta_p^{SMB}R_t^{SMB} + \beta_p^{HML}R_t^{HML} + \beta_p^{UMD}R_t^{UMD} + \varepsilon_{p,t}$$

Where $R_{p,t}$ is the return to the portfolio p at time t and $R_{f,t}$ is the risk-free rate. The Fama-French factors are R_t^M, the return to a value-weighted market portfolio; R_t^{SMB}, the difference of the return to small market value stocks minus larger firms; and R_t^{HML}, the difference between the return to high book-to-market stocks and low book-to-market firms. The Carhart (1997) Four-Factor model includes the Fama-French factors, as well as R_t^{UMD} — also known as the momentum factor (Jegadeesh and Titman, 1993, 2002, 2011) — the difference in returns between the lowest past-performing firms from the highest past-performing firms. The intercept, α_p, is interpreted as the abnormal return to the portfolio. Regressions use daily returns for firms and the daily returns from the value-weighted market portfolio from the CRSP database, and daily returns to the small-minus-big, high-minus-low and positive momentum-minus-negative momentum portfolios, R_t^{SMB}, R_t^{HML}, and R_t^{UMD} are from the factor file on Kenneth French's website.

Media reductions

Next, I create a sample of firms that have experienced a reduction in local media employment. I calculate the percentage change to employees in a metropolitan area as the sum of the two journalist categories across two-year windows to account for the rolling nature of the survey. Years in which the number of employees in the Reporters and Correspondents and Broadcast News Analyst categories falls 25 percent or more from the OES survey released two years prior are identified as negative shocks to coverage.[18] Because of changes to the survey, observations of shocks to employment are limited to 2003, as measured by the difference from 2001, and beyond.[19]

Bhojraj, Lee and Oler (2003) find that GICS classifications explain stock return comovements better than SIC and NAICS industry classification systems, and Levi and Welch (2017a, 2017b) find that a firm market value strongly correlates with a firm's beta and other market model factor similarities when compared with other firms. I find exact matches for each sample firm by year and by 6-digit G industry code, and then match, without replacement, by market value within 15% and within 15% of the beta of the treated firm's calculated beta in the Fama-French Three Factor model regressions.

I eliminate firms with relevant missing control variable data from Compustat or CRSP and firms with common stock share prices that closed the year trading under $2 to mitigate small changes in share price driving high variation in returns. My final sample consists of 1,373 firms headquartered in locations that experienced media employment reductions and their 1,373 controls. Available accounting and return information from Compustat and CRSP leave 24,532 firm-year observations across the 13 years from 2003 to 2018 considered.

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[18] The 25% cutoff is used because it closely approximates the employment reductions explored in Gao, Lee, and Murphy (2019) when one of a county's newspapers closes. Two of the examples mentioned in Gao, Lee, and Murphy (2019) are the closure of the *Denver Rocky Mountain News* in Colorado and the *Cincinnati Post* in Ohio. Those closures were reflected as 28% and 27%, respectively, losses in media category employment in the BLS data. Testing other cutoff levels of 20% and 30% yield similar results to those reported here.

[19] In 2003 and 2004, the OES was released twice, in May and November. I use the data released for the May survey results to ensure the survey windows are consistent across years. In 2004, the overall number of Reporters and Correspondents and Broadcast News Analysts were the same in both the May and November survey results.

Other variables

Annual returns are calculated using daily price changes from the CRSP database. Lagged monthly returns are calculated as the sum of daily log returns for each of the 12 individual months prior to the year being evaluated. Market value, total assets and other accounting information are taken from the Compustat database. Analyst coverage information is from the IBES database. Information on debt issuance and terms, including spread and covenant information, is taken from the Thomson Reuters DealScan database and linked by ticker symbol. The procedure provides 3,903 firm-year observations in which treated or control firms are identified as having issued debt.

METHODOLOGY

Returns, portfolios of Low-High media coverage firms

I use an unpaired t test to determine whether returns are significantly different from Low coverage firms and High coverage firms. I use the same test for each of the subsample portfolios based on size, book-to-market, and share price. I also conduct two nonparametric tests for robustness, a Wilcoxon rank sum test between the Low and High coverage firms and a test of median equivalence between the two sets of firms. I then calculate abnormal returns (the intercept, or alpha, α_p) to the long-short portfolio of Low-High coverage firms, employing the market model, the Fama-French Three Factor Model and the Carhart (1997) Four Factor model, and determine whether the intercepts are significantly different from zero.

Effects of reduction in media employment

I repeat the firm-level unpaired t test of overall returns to determine whether the difference between returns for firms that experience a reduction in nearby media employees is significantly different than the control sample at each year t from -6 to +6, where t = 0 is the year in which the reduction in media employment is recorded. I also conduct the two nonparametric tests for robustness.

Next I employ a dynamic difference-in-differences test to determine whether a reduction in the number of nearby media employees is associated with a significant difference in returns, abnormal returns as measured by alphas calculated from the Fama-French Three Factor Model, and valuation, as measured by market-to-book value. Figures 1 and 2 show median and mean overall returns for the sample of treated and control firms at year *t* = -6 to +6. Although returns vary considerably each year for both sets of firms, they appear to follow a consistent pattern, with treated firms' recording higher returns until year *t* = 0, when treated firm returns fall below control firm returns for several years before recovering their $t < 0$ trend. The model (1) calculates differences between firms that have experienced the media employment shock and those that have not in each of the periods $t \geq -1$, as well as controlling for state- and year-specific fixed effects.

(1)

$$Outcome_{\,i,t} = \sum_{k \geq 1} \beta_k \, Reduction\ in\ coverage\ has\ been\ in\ effect\ for\ k\ periods_{s,t} + \sum_{s} State\ fixed\ effects_s + \sum_{t} Time\ fixed\ effects_t + \beta_{i,t} Controls + \varepsilon_{i,s,t}$$

Where $Outcome_{\,i,t}$ is annual returns, abnormal returns (alpha from the Fama-French Three Factor model), and market-to-book value. Because the dependent variables are associated with market trading, I follow Levi and Welch (2017a, 2017b), who suggest that market value of equity and lagged returns are the most important determinants in market model specifications. I use the natural log of the firm's market value to normalize the values and calculate the lagged returns of each of the 12 months in the prior year and use each month's total as a control. I include the standard deviation of lagged returns as a control for risk. I also include controls that may affect the information environment of a firm, as compiled in Drake, Guest and Twedt (2014). *Analyst Coverage* is a dummy variable equal to 1 if the firm has at least one estimate issued during the year (LaFond, 2005). The *S&P 1500 Index* dummy variable is equal to 1 if a firm is listed on the S&P 1500 because being included on the high-profile index could both raise attention and scrutiny. I also include a firm's book-to-market ratio as a relevant element of an investors' value and estimation risk assessment (Desai, Rajgopal and Venkatachalam, 2004), and lagged beta to further control for the systematic elements of risk realized in the prior year. I include accounting measures of firm performance, as those may affect returns and valuation. ROA is net income scaled by book value of assets, and ROE is net income scaled by book value of equity. The first variable indicates efficient use of assets, and the second is a measure of delivered return to shareholders. All standard errors, in model (1) and all other regression models employed in this paper, are robust to heteroscedasticity and clustered by OES area number (the level of assignment of treatment) to address possible serial correlation in the error terms (Bertrand, Duflo and Mullainathan, 2004).

Debt spread terms

To explore the cost of debt, I employ a dynamic difference-in-differences design with the minimum spread in basis points of each loan facility recorded in DealScan for my matched sample:

(3)

$$Spread_{i,t} = \sum_{k\geq 1} \beta_k \, Reduction\ in\ coverage\ has\ been\ in\ effect\ for\ k\ periods_{s,t} + \sum_{s} State\ fixed\ effects_s + \sum_{t} Time\ fixed\ effects_t + \beta_{i,t} Controls + \varepsilon_{i,s,t}$$

I use four models, each including different controls. The first model uses controls typical in prior literature for firm and loan characteristics (Bharath, Sunder and Sunder, 2008). The firm characteristics included are market-to-book value, tangibility, leverage, and the current ratio. Market-to-book value, calculated as the market value of the firm at the end of the year over the book value of its assets, is an indicator of growth prospects. Tangibility is the accounting value of property, plant, and equipment over total assets; leverage is long-term total debt over total assets; and the current ratio is current assets divided by current liabilities. Each of the three provides different measures of liquidity that lenders use to assess a borrower's ability to repay the loan. Loan characteristics are the amount of the loan facility and the maturity of the loan, both log-transformed, and a dummy variable indicating whether the facility includes a performance pricing provision. In addition to spread, lenders can use loan amounts and maturities to mitigate the risk of default. Performance pricing provisions allow lenders to increase spreads when financial performance deteriorates to the point in which a covenant violation occurs and provide another mechanism for lenders to mitigate risk (Smith and Warner, 1979; Dichev and Skinner, 2002).

My second model includes atypical debt controls for the information environment of the firm. Although the informativeness of accounting quality in lending decisions is well-documented (Francis *et al.*, 2005; Core, Guay and Verdi, 2008; Dechow, Ge and Schrand, 2010; McNichols and Stubben, 2015), other external monitoring of the firm has been less typical in the literature. Because the role of external information intermediaries are central to the research question here, I include controls for two sources of external monitoring that are well-established in the literature in regards to returns. Analyst coverage is a dummy variable equal to 1 if at least one sell-side analyst has issued an earnings report for the firm, and the S&P 1500 Index dummy variable is equal to 1 if the firm is included in the highly watched index.

My third model includes controls for performance of the firm during the year the loan facility is issued. Earnings is net income scaled by market value of the firm, and return on assets is net income scaled by total assets. The fourth model includes all of the previous explained controls, including the controls for firm and loan characteristics, information environment, and contemporary firm performance.

RESULTS

Abnormal returns related to coverage levels

The number of reporters and broadcast analysts in a metropolitan statistical area offers no evidence whether a firm was written about in a major newspaper. However, portfolios constructed of firms with the lowest number of overall media employees and the highest deliver similarly high returns (both economically and statistically) as portfolios constructed of No Coverage and High Coverage firms reported in Fang and Peress (2009). In their paper, No Coverage firms are defined as those with no articles about them during the month, while Low and High Coverage firms are determined by whether the number of articles about a firm are below or above the median number of total news articles about firms.

Univariate results, returns

I use an unpaired t test to determine whether returns for Low coverage firms are significantly different from High coverage firms. The Low coverage firms are significantly higher, as shown in Table 1. In panels A-C, I use the same test for each of the subsample portfolios based on size, book-to-market, and share price. I also conduct two nonparametric tests for robustness, a Wilcoxon rank sum test between the Low and High coverage portfolios and a test of median equivalence between the two portfolios, shown in Panel D and E of Table 1. In the case of the full sample, results are consistent with the findings in Fang and Peress (2009) between Low (No) coverage firms and High (High) coverage firms. The Low coverage firms (the middle tercile) in Fang and Peress (2009), however, often delivered anomalously high (when compared with the No coverage firms) or low (when compared with High coverage firms) returns. In my sample, the effect of coverage appears to be more linear, with Middle coverage firm returns falling between the other two terciles of firms. The results suggest that sorting firms by local media employment captures a similar effect to the more traditional definition of coverage based on article counts documented in Fang and Peress (2009).

Regression analysis

Next, I explore whether the media effect holds in constructing long-short portfolios based on number of media employees located in the same area as a firm's headquarters. Again following Fang and Peress (2009), I compute the monthly return on a zero-investment portfolio that longs the stocks with the lowest number of media employees and shorts the stocks with the highest number. Fang and Peress constructed their portfolios based on longing the No coverage firms and shorting the High coverage firms. The returns to the portfolio for each month yield a time series of returns for the zero-investment portfolio. The time-series returns are then regressed on factors known to affect the cross-section of returns. I examine three of the factor models in Fang and Peress (2009): the capital asset pricing model, the Fama-French (1993) three-factor model, and the Carhart (1997) four-factor model.

Using my full sample of firms from 2003 to 2018, the results, shown in Table 2, are directionally consistent with the results in Fang and Peress (2009). The intercept (interpreted as the abnormal return to the portfolio) is not significantly different from zero in my regressions, whereas it is significantly positive in Fang and Peress (2009). A subsample of the portfolio returns from 2003 to 2009 finds significant positive returns ($p < 0.05$) to the long-short media strategy and coefficients are similar to those reported in Fang and Peress (2009). It is possible the gains to the No coverage portfolio have been largely arbitraged away since the 2009 publication; however, the values of the coefficients on the media employment portfolios are actually larger than the annually adjusted coefficients in Fang and Peress (2009) while the R-squared for each model is considerably lower. Blitz (2019) finds that the widely used Fama-French factors delivered consistently negative returns from 2010 to 2019, while other, less common factors, delivered positive premiums. Therefore, comparing the results of the models may be inappropriate across these particular time frames. Fang and Peress' (2009) sample was from 1993 to 2002. In Fang and Peress' sample, HML had a significantly positive relationship ($p < 0.01$) with returns in the three- and four-factor models, and UMD was significantly positive ($p < 0.01$) in the four-factor specification. Neither of those factors were significant in my full sample tests; however, the SMB factor was positive in both my tests and Fang and Peress (2009), and HML was significant ($p < 0.05$) in my sample that considered only 2003-2009.

The Fang and Peress (2009) results are somewhat suspect in light of the results in Gurun and Butler (2012), who find that media coverage is essentially captured by advertisers and is subject to short-term reversal. Any negative return to High Coverage firms in the Fang and Peress (2009) sample, which is determined monthly, could be driven by those reversals. My results, in an annual setting, support the Fang and Peress (2009) conclusions that investors require higher returns from firms with less coverage, however "coverage" is defined.

After a reduction in media employment

The univariate t tests and nonparametric tests, shown in Table 3, find little evidence that returns are significantly different between the firms that experienced a reduction in nearby media employment and their controls after the reduction in media employment is recorded. Consistent with Figures 1 and 2, treated firms have consistently higher mean and median returns in the years prior to the reduction in media employment, significantly higher at years t = -3 and -4 ($p < 0.1$ for means and $p < 0.05$ for medians). No significant differences are found after the media reduction is recorded, although in many years the control firms' mean and median returns are higher. Nonparametric tests, a Wilcoxon rank sum test and a test of the equality of medians, support these findings.

In the dynamic difference-in-differences results only year $t = 2$ shows a significant relationship ($p < 0.05$) between a reduction in media employment and returns and abnormal returns, in columns 1 and 2 of Table 4. The negative correlation is the opposite of that predicted by attention and estimation risk models, which suggest that valuations will fall as current shareholders at the margins will move their investments away from the security and new shareholders will not know to invest or will decline to bear the added estimation risk. The remaining shareholders are predict to demand higher returns as additional risk premia. The market-to-book value has a significant positive relationship at years $t = 1$ ($p < 0.05$), t = 2 ($p < 0.01$), and $t = 3$ and 4 ($p < 0.1$). Again, this is the opposite direction predicted by attention and estimation risk models. Together, the results of the univariate tests and the dynamic difference-in-differences test provide support for the null, Hypothesis 2_0 rather than the affirmative, Hypothesis 2a.

Cost of debt

Univariate results

A reduction in media employment may not show the predicted decrease in valuation and increase in returns because the firm-level return effects are limited to idiosyncratic risk, which can be diversified away, or because firms take other measures to maintain valuation and return levels such as issuing debt (Maskara and Mullineaux, 2011). Lenders also face adverse selection risks and use loan term spread to mitigate their risk from their asymmetrical information with managers (Leland and Pyle, 1977; Smith and Warner, 1979; Bushman and Smith, 2001; Dichev and Skinner, 2002; Zhang, 2008; Kim, Song and Zhang, 2011). Auditors appear to recognize that media coverage can uncover bad news (Mutchler, Hopwood and McKeown, 1997; Joe, 2003; Gong, Gul and Shan, 2018), and lenders may charge additional risk premia, through increased spreads, when that external monitoring function is reduced.

I identify 3,903 firms in my treated and control samples with a debt issuance recorded in DealScan. Unpaired t tests between control firms and treated firms, reported in Panel A of Table 5, find no significant differences in the means of the spreads in the years t = -6 to +1 around a reduction in local media employment at year t = 0. However, in years t = +2 and +3, treated firms have significantly higher average spreads ($p < 0.1$ and $p < 0.05$, respectively). Nonparametric analysis, in Table 5, Panel B, finds significantly higher median spreads for treated firms at years t = +1, +3 and +5 ($p < 0.1$, $p < 0.05$, and $p < 0.01$, respectively). Figures 3 and 4 illustrate the mean and median spreads for treated and control firms from year t = -6 to +6.

Dynamic difference-in-differences

I employ four models, shown in Table 6, to examine the relationship between a reduction in media employment and spreads. Model (1) includes controls for firm and loan characteristics typically related to spreads, which are divided by 1,000 to allow coefficients to be interpreted as percentages. Market-to-book value and loan size are negatively correlated with spreads, as is the dummy variable equal to 1 if the loan includes a price performance provision that allows lenders to increase the spread if performance falls below a certain threshold. Firm leverage and loan maturity are positively correlated with spreads. The variables of interest, year dummies for years $t = -1$ to $+6$ around the reduction in media employment, are significant at years $t = +1$ through +3 ($p < 0.05$, $p < 0.1$, and $p < 0.01$, respectively). Year t = +2 and +3 are significantly positively correlated in all specifications and year t = +1 in all but the model (3) that includes performance measurements. Models 2 and 4 include other controls related to the information environment of the firm. Being included in the S&P 1500 index ($p < 0.01$) and analyst coverage ($p < 0.05$) are significantly correlated with lower spreads in model 2, which does not include performance measurements. Inclusion in the S&P 1500 is significantly negatively correlated with spreads in model 4, as well. This is consistent with lenders using information or certification from outside sources, as well as their private information from borrowers, to mitigate their risk.

Additional tests

Capital structure theory predicts that firms should prioritize equity financing to issuing debt (Myers and Majluf, 1984) to avoid adverse selection problems or that firms have a target capital structure that balances various costs (Graham and Leary, 2011). Despite evidence that firms that experience a reduction in nearby media employment do not face an increased cost of equity, I found no instances of seasoned equity offerings in my sample using the CRSP distribution dataset. Even if overall valuations had increased, Welch (2004) finds that U.S. corporations do not issue and repurchase debt and equity to counteract the mechanistic effects of stock returns on their debt-equity ratios.

Debt ratios and debt issuance

I employ one of the measurements, debt-to-equity ratio, common in the finance and economics capital structure literature. A firm's debt ratio is calculated as its book value of debt over the sum of its book value of debt and its market value of equity:

$$Debt\ ratio_{i,t} = \frac{D_{i,t}}{D_{i,t} + S_{i,t}P_{i,t}}$$

Where $Debt\ ratio_{i,t}$ is equal to the sum of short-term and long-term debt for firm i at time t, $S_{i,t}$ is the number of shares outstanding at the end of time t and $P_{i,t}$ is the share price. Figure 5 shows the mean debt ratios for treated and control firms; however, no pattern is evident, and the pre-treatment trends indicate a difference-in-differences is an inappropriate statistical tool.

To estimate a firm's financing decisions, instead I employ the full sample of firms that experienced a reduction in local media employment and their controls to estimate a probit model on a firm's decision to issue debt, where the outcome variable is set to 1 if a facility loan appears in the DealScan database and 0 otherwise.

$$Debt\ issue_i = \alpha_p + \beta_i Treat + \beta_i Cash + \beta_i OIBD + \beta_i CAPEX + \beta_i MTB + \beta_i MTB + \beta_i Sales + \beta_i RD + \beta_i RDD + \beta_i Adj\_returns + \beta_i Age + \varepsilon_{p,t}$$

Where *Treat* is a dummy variable equal to 1 if a firm is in year $t > -1$ and is headquartered in area that has experienced a media reduction. I include controls for the firm's level of cash, scaled by total book value of assets (CASH); operating income before depreciation, scaled by assets (OIBD); capital expenditures, scaled by assets (CAPEX); market to book value (MTB); the log-transformed value of sales (Sales); book leverage, defined as book debt scaled by assets; research and development expenditures, scaled by assets (RD); a dummy variable for missing values of R&D (RDD)[20]; the difference between the raw returns of the firm one year prior and the year-ago returns to the value-weighted market portfolio recorded in the CRSP database (Adj. returns); and firm age (Age). Firms with more cash and higher profitability are less likely to seek external financing. Growth firms, as measured by capital expenditures, higher market-to-book values, R&D expenditures, and firm returns, are more likely to seek external financing. Age is included because older firms are less likely to require external financing. Financial statement information is taken from the Compustat database; adjusted returns are calculated using information on firm returns and returns to the value-weighted market portfolio from the CRSP database; and Age is calculated as the year of the firm-year observation minus the year a market value first appeared for the firm in Compustat. Missing values in Compustat leave 24,982 firm-year observations in my analysis.

Table 7, Panel A reports the results of the probit regression. *Treat* is significantly positive ($p < 0.01$), suggesting firms are more likely to issue debt after a reduction in media employment. The marginal effect of a reduction in media employment on a debt issuance decision is 1.9% at the mean. I also estimate a logit regression to include fixed effects for state, year, and industry, as each of these may play a role in the borrowing environment and financing decisions of firms. The

[20] Following Huang and Ritter (2009), I set R&D equal to 0 if the value is missing in Compustat, both to preserve the sample and because generally no value recorded indicates a firm has not invested in R&D. However, to control for any variation explained by firms that may not record R&D expenses, Huang and Ritter include a dummy variable, RDD, equal to 1 if the value of R&D expense is missing, which I also employ here.

results, reported in Table 7, Panel B, are similar to the probit model, with *Treat* again showing significantly positive correlation on the debt issuance decision ($p < 0.01$). The marginal effect of a reduction in media employment is 2% at the mean.

CONCLUSION

Overall, lower levels of media coverage appear to be consistent with models that predict investors will require higher returns. The relationship appears robust to a variety of definitions of "coverage." In Chan (2003), "coverage" is the number of corporate press releases picked up by the Dow Jones News Wires. In Fang and Peress (2009), "coverage" is the number of articles that appear in four national newspapers. In Engelberg and Parsons (2011) "coverage" is articles in local newspapers. Here, I define "coverage" as the number of media employees nearby and able to cover a firm. The effects on firm valuations, returns, and risk appear similar across definitions.

However, a reduction in the number of media employees (i.e., a reduction in coverage) does not appear to affect returns in the predicted way. Local investors tend to overweight their portfolios with local stocks and are more likely to buy than sell (Huberman, 2001), which may hold prices and returns steady. Even though as shown in Essay 1, firm-specific risk increases after a reduction in local media employment, more sophisticated investors may be able to diversify that risk away. Maskara and Mullineaux (2011) suggest firms can use debt to "certify" their disclosures and improve share prices by announcing the loans.

Although firms that experience a decrease in media coverage, as defined here, do not appear to face increased costs of equity capital, the costs of debt capital appear to rise, even after controlling for leverage. Despite this increased cost of debt financing and steady or reduced costs of equity capital, firms that experience a reduction in media coverage appear more likely to opt to issue debt. The puzzling finding is not explained by current attention or investor estimation risk models, nor is it obviously explained by capital structure theory.

Tables

Table 2.1: Media Coverage and Stock Returns: Univariate Comparisons

This table presents average monthly returns for stocks with low and high numbers of media employees in the geographic area where they are headquartered and repeats the results of No Coverage and High Coverage firms in Fang and Peress (2009), FP. The table also shows the difference between Low (No) and High (High) returns and the results of a t-test of the differences between the means. Panel A-C presents the same analysis for select subsamples based on terciles of firm size, book-to-market, and share price, with 1 being the largest market value firms and 3 being the smallest, etc. Panels D and E present nonparametric tests of the differences in returns for the full sample.

	Average Annual Return				Average No. of stocks	
	Low (No)	High	Low-High	t-statistic for Low-High	Low (No)	High
All stocks by number of media employees	10.37%	6.56%	3.82%	7.21	795.56	775.00
FP annual return (monthly return X 12)	16.20%	11.52%	4.68%	2.13	1,430.08	245.40
Panel A: By Size						
1 By Number of Media Employees	9.91%	4.01%	5.90%	5.88	236.00	311.63
2 By Number of Media Employees	10.16%	6.80%	3.37%	3.58	271.25	249.63
3 By Number of Media Employees	10.94%	9.99%	0.95%	1.19	288.00	213.44
1 FP annual return (monthly return x 12)	16.92%	6.36%	10.56%	1.74	578.55	17.98
2 FP annual return (monthly return x 12)	16.08%	8.28%	7.80%	2.68	514.23	46.71
3 FP annual return (monthly return x 12)	15.24%	13.20%	2.04%	1.03	337.42	149.19
Panel B: By Book-to-market						
1 By Number of Media Employees	21.44%	18.32%	3.13%	2.56	254.13	174.94
2 By Number of Media Employees	10.65%	9.12%	1.53%	1.67	242.13	182.44
3 By Number of Media Employees	0.88%	0.97%	-0.10%	-0.09	252.13	184.00
1 FP annual return (monthly return x 12)	14.28%	10.44%	3.84%	1.25	441.79	81.50
2 FP annual return (monthly return x 12)	14.76%	6.48%	8.28%	3.13	450.03	74.64
3 FP annual return (monthly return x 12)	17.04%	14.28%	2.76%	0.85	460.78	70.57
Panel C: By Price						
1 By Number of Media Employees	21.44%	18.32%	3.13%	2.56	254.13	174.94
2 By Number of Media Employees	10.65%	9.12%	1.53%	1.67	242.13	182.44
3 By Number of Media Employees	0.88%	0.97%	-0.10%	-0.09	252.13	184.00
1 FP annual return (monthly return x 12)	12.12%	-1.32%	13.44%	3.14	545.18	35.76
2 FP annual return (monthly return x 12)	16.68%	6.48%	10.20%	3.77	500.77	60.14
3 FP annual return (monthly return x 12)	21.24%	16.20%	5.04%	2.62	384.13	128.21
Panel D: Wilcoxon rank sum test of differences						
	Low median	High median			Rank sum Z score	
All stocks by number of media employees	8.90%	5.00%			8.29	
Panel E: Test of median equivalence						
	Low	High			Pearson chi^2	
All stocks below median	5,973	6,592			97.73	
All stocks above median	6,756	5,808			P(medians equal) < 0.001	

Table 2.2: Media-Related Trading Profits

Panel A of this table shows the profitability of a trading strategy that longs stocks with the lowest number of media employees in the firm's area by tercile and shorts stocks with the largest number of media employees. Both the long and short positions are equally weighted, and portfolios are rebalanced annually. The resulting time-series returns on the long–short portfolio are regressed on widely accepted risk factors. Panel B highlights results from Fang and Peress (2009) shows the profitability of a trading strategy that longs stocks with no media coverage (no articles in four publications) and shorts stocks with high media coverage (above the median of firms that had any articles written about it). Both the long and short positions are equally weighted, and portfolios are rebalanced monthly. The resulting time-series returns on the long–short portfolio are regressed on widely accepted risk factors. Coefficients in the Fang and Peress (2009) sample have been adjusted to reflect compounded annual returns from the monthly portfolios.

Panel A: Low-High coverage portfolios based on media employment

	Model (1) 2003-2018 CAPM	Model (2) 2003-2018 3 Factor	Model (3) 2003-2018 4 Factor	Model (1) 2003-2009 CAPM	Model (2) 2003-2009 3 Factor	Model (3) 2003-2009 4 Factor
Mkt-rf	5.362**	1.977	2.684	-15.315***	-18.442***	-18.836***
	(2.636)	(2.861)	(2.964)	(2.847)	(3.257)	(3.555)
SMB		14.431***	14.129***		1.740	1.894
		(4.790)	(4.803)		(5.579)	(5.638)
HML		1.749	3.200		11.438**	11.029**
		(4.361)	(4.640)		(4.741)	(4.980)
UMD			2.448			-0.729
			(2.666)			(2.556)
Alpha (intercept)	0.138	0.138	0.133	0.250**	0.236*	0.235*
	(0.107)	(0.105)	(0.105)	(0.123)	(0.121)	(0.122)
Obs.	192	192	192	84	84	84
R-squared	0.021	0.067	0.072	0.261	0.311	0.312

Panel B: No-High coverage portfolios in Fang and Peress (2009)

	Model (1) FP CAPM	Model (2) FP 3 Factor	Model (3) FP 4 Factor
Mkt-rf	-0.844***	-0.779***	-0.682***
	-	-	-
SMB		43.489***	14.129***
		-	-
HML		4.814***	39.108***
		-	-
UMD			1.427***
			-
Alpha (intercept)	0.055**	0.043**	0.029**
	-	-	-
Obs.	119	119	119
R-squared	0.11	0.58	0.62

Standard errors are in parenthesis in Panel A; Standard errors not reported in Fang and Peress (2009)

Table 2.3: Test of differences in returns for treated, control firms at year t

Table 3 shows results of parametric (Panel A) and nonparametric tests (Panel B) of differences between average returns for a sample of firms that experienced a 25% media reduction in the area in which they were headquartered and a control sample of firms matched on 6-digit GICS industry category, year, market value and market beta from Fama-French 3 Factor model regressions. Year t = 0 is the year in which the reduction to media employment was recorded.

Panel A: Unpaired t tests

	Control	Treated	Control Avg. ret.	Treated Avg. ret.	Diff	St_Err	t value	p value
Year t = -6	859	887	-.009	.013	-.022	.022	-1	.313
Year t = -5	925	952	-.053	-.048	-.005	.024	-.25	.815
Year t = -4	992	1,024	-.006	.034	-.04	.022	-1.75	.079
Year t = -3	1,078	1,082	-.105	-.061	-.043	.023	-1.9	.058
Year t = -2	1,173	1,158	.025	.030	-.005	.021	-.25	.788
Year t = -1	1,264	1,256	-.044	-.043	-.001	.020	-.05	.966
Year t = 0	1,373	1,373	.055	.051	.003	.018	.2	.861
Year t = 1	1,279	1,324	-.036	-.032	-.004	.019	-.25	.82
Year t = 2	1,135	1,245	-.025	-.047	.022	.018	1.25	.209
Year t = 3	991	1,170	-.097	-.106	.009	.021	.4	.686
Year t = 4	800	1,008	-.102	-.099	-.003	.026	-.1	.922
Year t = 5	685	908	-.085	-.083	-.002	.029	-.05	.941
Year t = 6	599	824	.046	.047	-.001	.024	0	.985

Panel B: Two-sample Median Equivalence, Wilcoxon rank-sum (Mann-Whitney) test

	Control	Treated	Control Med. ret.	Treated Med. ret.	Pearson Chi-square	P(equal)	Rank sum Z	p value
Year t = -6	859	887	.022	.032	.1123	.738	-0.724	.469
Year t = -5	925	952	-.009	.004	.1543	.694	-0.540	.589
Year t = -4	992	1,024	.017	.058	3.1754	.075	-2.125	.033
Year t = -3	1,078	1,082	-.028	.002	2.9630	.085	-2.049	.041
Year t = -2	1,173	1,158	.066	.074	.3606	.548	-.0421	.674
Year t = -1	1,264	1,256	.034	.042	.1016	.750	-0.224	.823
Year t = 0	1,373	1,373	.060	.072	.1763	.675	-0.251	.802
Year t = 1	1,279	1,324	.041	.040	.0464	.830	-0.302	.763
Year t = 2	1,135	1,245	.017	-.009	1.4165	.234	1.406	.160
Year t = 3	991	1,170	-.001	-.008	.1037	.747	0.448	.654
Year t = 4	800	1,008	-.008	-.012	.0807	.776	0.421	.674
Year t = 5	685	908	.030	.024	.0755	.783	0.397	.692
Year t = 6	599	824	.100	.093	.3761	.540	0.175	.861

Table 2.4: Regression results on returns, alpha, and market-to-book

The results from a dynamic difference-in-differences model with state and year fixed effects. Two measures of returns, overall (1) and abnormal returns as measured by the intercept from Fama-French 3 Factor regression (2), and market-to-book value (3) are the dependent variables. Variables Years t = -1 to + 6 show the correlation between a reduction in media employment where a firm is headquartered and the outcome variable.

	(1) Returns	(2) Alpha	(3) MTB
Market value	0.056***	0.021***	-0.356***
	(0.006)	(0.002)	(0.077)
Lag market beta	-0.131***	-0.069***	0.411***
	(0.010)	(0.004)	(0.092)
Book-to-market	-0.117***	-0.043***	-1.841***
	(0.011)	(0.004)	(0.201)
ROA	-0.426**	-0.155**	12.289***
	(0.192)	(0.075)	(3.216)
ROE	0.491***	0.181***	-4.551***
	(0.120)	(0.047)	(1.338)
Analyst Coverage	0.006	0.001	-0.257***
	(0.009)	(0.004)	(0.095)
S&P 1500 Index	-0.052***	-0.018***	-0.187**
	(0.008)	(0.003)	(0.078)
Lag S.D. of returns	0.177	0.026	
	(0.590)	(0.223)	
Year t = -1	-0.003	0.005	0.131
	(0.017)	(0.005)	(0.096)
Year t = 0	0.011	0.005	0.148
	(0.013)	(0.005)	(0.122)
Year t = 1	-0.024	-0.010	0.232**
	(0.017)	(0.006)	(0.115)
Year t = 2	-0.027**	-0.010**	0.212***
	(0.013)	(0.005)	(0.075)
Year t = 3	0.004	0.001	0.249*
	(0.014)	(0.006)	(0.131)
Year t = 4	-0.013	-0.007	0.170*
	(0.016)	(0.007)	(0.096)
Year t = 5	-0.011	-0.007	0.149
	(0.017)	(0.007)	(0.115)
Year t = 6	0.021	0.005	0.076
	(0.015)	(0.006)	(0.115)
Obs.	24,532	24,532	24,533
R-squared	0.356	0.204	0.701
Adj. R-squared	0.354	0.202	0.700
Year fixed effects	Yes	Yes	Yes
State fixed effects	Yes	Yes	Yes

Standard errors are in parenthesis
**** p<0.01, ** p<0.05, * p<0.1*

Table 2.5: Tests of differences between loan interest rates for treated, control firms

Table 5 shows results of parametric (Panel A) and nonparametric tests (Panel B) of differences between loan interest spreads for firms that took loans, taken from a sample of firms that experienced a 25% media reduction in the area in which they were headquartered and a control sample of firms matched on 6-digit GICS industry category, year, market value and market beta from Fama-French 3 Factor model regressions. Year *t* = 0 is the year in which the reduction to media employment was recorded.

Panel A: Unpaired t tests

	Control N	Treated N	Control avg. spread	Treated avg. spread	Diff	St. Err.	t value	p value
Year t = -6	150	172	158.81	171.97	-13.15	19.55	-.65	.5
Year t = -5	159	208	172.84	161.72	11.12	17.73	.65	.53
Year t = -4	157	241	159.08	162.5	-3.42	15.47	-.2	.83
Year t = -3	155	234	172	173.08	-1.08	13.57	-.1	.94
Year t = -2	143	244	189.72	178.73	10.98	15.95	.7	.49
Year t = -1	170	241	187.74	182.41	5.33	14.07	.4	.7
Year t = 0	157	234	179.64	195.69	-16.05	14.46	-1.1	.27
Year t = 1	176	289	186.24	203.81	-17.57	12.36	-1.4	.16
Year t = 2	128	276	181.4	212.13	-30.73	17.1	-1.8	.07
Year t = 3	167	248	184.75	212	-27.25	12.96	-2.1	.04
Year t = 4	137	228	176.4	192.52	-16.12	13.09	-1.25	.22
Year t = 5	122	204	187.68	210.14	-22.46	14.56	-1.55	.12
Year t = 6	101	209	187.55	190.66	-3.11	10.24	-.3	.76

Panel B: Two-sample Wilcoxon rank-sum (Mann-Whitney) test, Median Equivalence

	Control	Treated	Control median spread	Treated median spread	Pearson Chi-square	P(equal)	Rank sum Z score	p value
Year t = -6	150	172	115.825	150	.4915	.483	-0.546	.5852
Year t = -5	159	208	125	162.5	.0321	.858	0.193	.8471
Year t = -4	157	241	125	132.5	1.5685	.210	-0.826	.4087
Year t = -3	155	234	150	145	.4091	.522	0.168	.8668
Year t = -2	143	244	150	138.75	.0279	.867	0.682	.4950
Year t = -1	170	241	175	150	.4127	.521	0.701	.4833
Year t = 0	157	234	165	150	.3125	.576	-0.106	.9159
Year t = 1	176	289	165	175	.7818	.377	-1.696	.0899
Year t = 2	128	276	150	160	.07503	.386	-1.033	.3018
Year t = 3	167	248	150	200	11.1861	.001	-2.030	.0424
Year t = 4	137	228	150	165	2.5028	.114	-0.978	.3279
Year t = 5	122	204	150	175	8.6984	.003	-2.676	.0075
Year t = 6	101	209	162.5	175	.1413	.707	-0.630	.5289

Table 2.6: Loan spread of firms that experience media-related shock, controls

Results of dynamic differences-in-differences with state and year fixed effects. Loan facility spreads, the dependent variable, are taken from DealScan and matched with a sample of firms that experienced a reduction in media employment in their geographic areas and a sample of control firms. Variables Years t = -1 to + 6 show the effect of a reduction in media employment. Model 1 includes standard firm and loan characteristic controls. Model 2 includes previous controls and dummy variables representing the information environment of the firm. Model 3 controls for firm performance. And Model 4 includes both information environment and firm performance controls.

	(1) Standard	(2) Information Environment	(3) Performance Measures	(4) Info & performance
Analyst Coverage		-0.024**		-0.015
		(0.010)		(0.010)
S&P 1500 Index		-0.049***		-0.043***
		(0.009)		(0.008)
ROA			-0.222***	-0.202***
			(0.045)	(0.045)
Earnings			-0.004***	-0.003***
			(0.001)	(0.001)
MTB	-0.009**	-0.009**	-0.009**	-0.009**
	(0.004)	(0.004)	(0.004)	(0.004)
Tangibility	-0.000	-0.003	-0.001	-0.003
	(0.012)	(0.012)	(0.011)	(0.011)
Leverage	0.187***	0.169***	0.175***	0.160***
	(0.025)	(0.026)	(0.020)	(0.021)
Current ratio	0.003	0.002	0.004	0.003
	(0.004)	(0.003)	(0.004)	(0.004)
Log loan size	-0.023***	-0.017***	-0.021***	-0.016***
	(0.006)	(0.005)	(0.005)	(0.005)
Log loan maturity	0.036***	0.033***	0.038***	0.035***
	(0.007)	(0.007)	(0.007)	(0.006)
Price changes	-0.067**	-0.047	-0.085***	-0.067***
	(0.031)	(0.029)	(0.020)	(0.020)
Year t = -1	0.006	0.010	0.007	0.010
	(0.011)	(0.011)	(0.011)	(0.011)
Year t = 0	0.016	0.017	0.010	0.011
	(0.014)	(0.014)	(0.012)	(0.012)
Year t = 1	0.023**	0.023**	0.016	0.017*
	(0.010)	(0.010)	(0.010)	(0.010)
Year t = 2	0.019*	0.020**	0.016*	0.018**
	(0.010)	(0.010)	(0.009)	(0.009)
Year t = 3	0.032***	0.032***	0.026**	0.027***
	(0.011)	(0.011)	(0.010)	(0.010)
Year t = 4	0.001	0.004	0.004	0.006
	(0.010)	(0.011)	(0.010)	(0.011)
Year t = 5	0.005	0.012	0.002	0.009
	(0.013)	(0.013)	(0.013)	(0.013)
Obs.	3903	3903	3903	3903
R-squared	0.729	0.738	0.741	0.747
Adjusted R-squared	0.724	0.733	0.736	0.742
Year fixed effects	Yes	Yes	Yes	Yes
State fixed effects	Yes	Yes	Yes	Yes

Standard errors are in parenthesis
**** p<0.01, ** p<0.05, * p<0.1*

Table 2.7: Likelihood of Debt Issue After Media-Related Event

The results from an estimation of the probability a firm decides to issue debt, where the outcome variable is set to 1 if a facility loan appears in the DealScan database and 0 otherwise. The full sample of firms that experienced a media employment reduction in their geographic area and a set of control firms are included, and *Treat* is a dummy variable equal to 1 if a firm experienced the media reduction and the firm year is after the reduction in media employment. Panel A reports the results of a probit model that includes controls for characteristics associated with debt issues, and Panel B reports the results of a logit model that includes state, year, and 6-digit GICS industry fixed effects.

Panel A: Probit model

P(Debt issue)	Coef.	St.Err.	t-value	p-value	[95% Conf	Interval]
Treat	0.070***	0.020	3.44	0.001	0.030	0.110
OIBD	0.221**	0.103	2.14	0.032	0.019	0.423
CAPEX	1.411***	0.140	10.05	0.000	1.136	1.687
MTB	-0.022**	0.010	-2.19	0.028	-0.042	-0.002
Sales	0.261***	0.007	37.33	0.000	0.248	0.275
RDD	-0.055**	0.024	-2.25	0.024	-0.102	-0.007
RD	-0.188	0.297	-0.63	0.528	-0.771	0.395
Adjust returns	0.122***	0.026	4.77	0.000	0.072	0.172
Age	0.002**	0.001	2.47	0.013	0.000	0.003
Constant	-2.631***	0.054	-49.15	0.000	-2.736	-2.526

Mean dependent var	0.180	SD dependent var	0.384
Pseudo r-squared	0.121	Number of obs	25,291
Chi-square	2119.934	Prob > chi2	0.000

Panel B: Logit model with year, state, and industry fixed effects

P(Debt issue)	Coef.	St.Err.	t-value	p-value	[95% Conf	Interval]
Treat	0.140***	0.036	3.92	0.000	0.070	0.210
OIBD	0.743***	0.216	3.43	0.001	0.319	1.167
CAPEX	2.392***	0.248	9.63	0.000	1.905	2.878
MTB	-0.064***	0.020	-3.25	0.001	-0.103	-0.025
Sales	0.458***	0.012	36.83	0.000	0.434	0.483
RDD	-0.128***	0.045	-2.88	0.004	-0.216	-0.041
RD	-1.221*	0.704	-1.74	0.083	-2.601	0.159
Adjust returns	0.226***	0.047	4.82	0.000	0.134	0.318
Age	0.003***	0.001	2.63	0.009	0.001	0.005
Constant	-4.544***	0.100	-45.66	0.000	-4.739	-4.349

Mean dependent var	0.180	SD dependent var	0.384
Pseudo r-squared	0.123	Number of obs	25,291
Chi-square	2199.794	Prob > chi2	0.000

**** p<0.01, ** p<0.05, * p<0.1*

FIGURES

Figure 2.1: Annual median returns

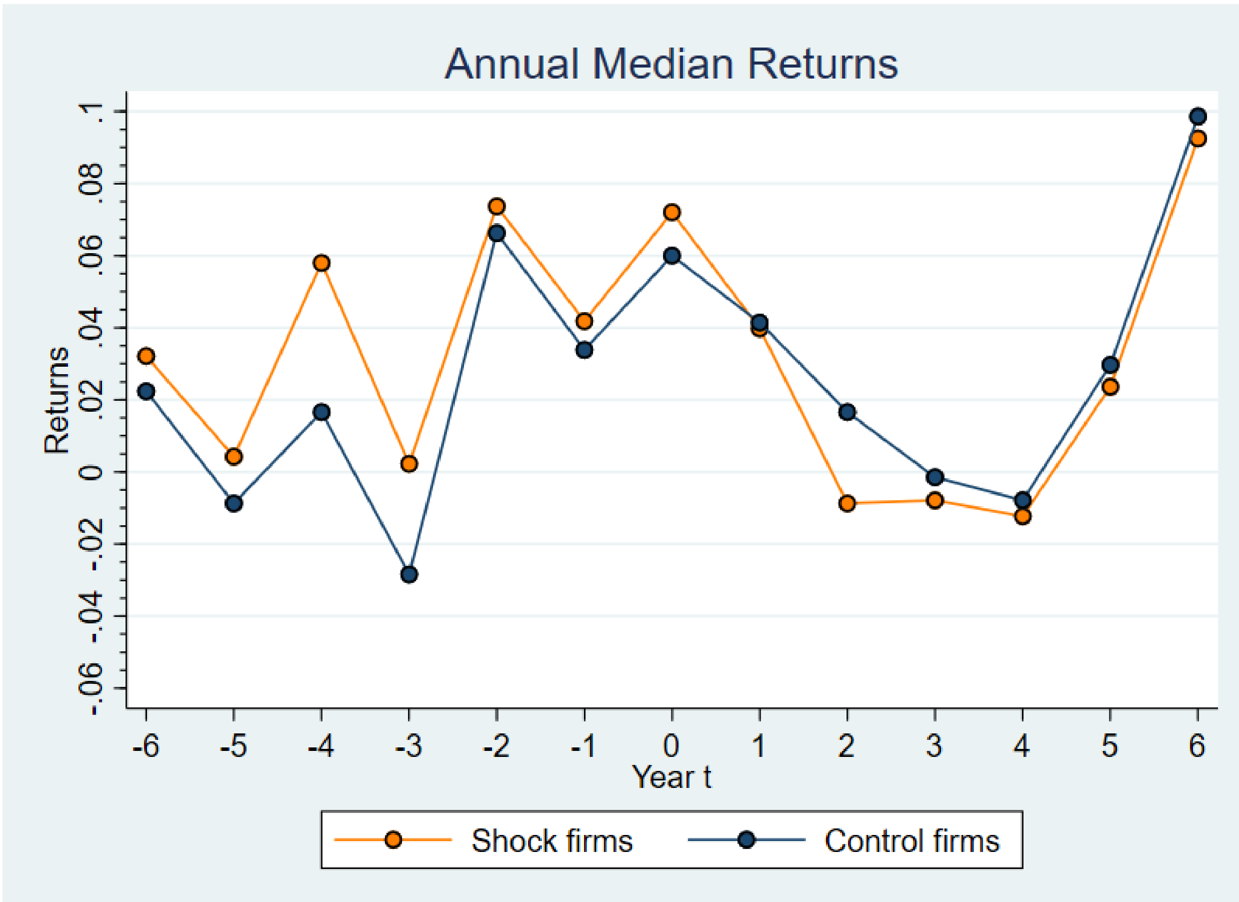

Figure 1: The annual median returns of firms that have recorded a 25% or more reduction in local media employment in the geographic area where the companies are headquartered at year t = 0 and a sample of control firms matched on market value and systematic risk (market beta from the Fama-French 3 Factor model).

Figure 2.2: Annual average returns

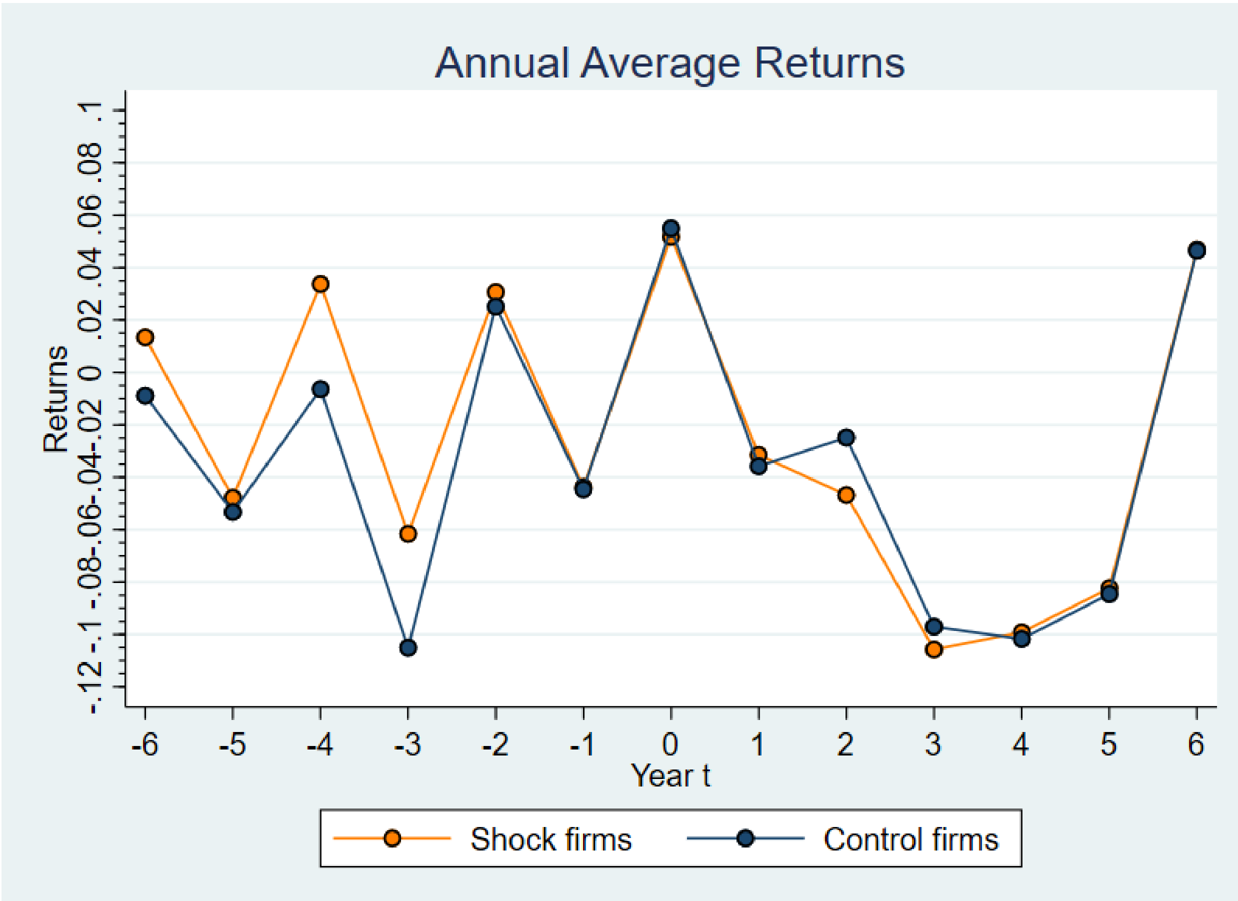

Figure 2: The annual average returns of firms that have recorded a 25% or more reduction in local media employment in the geographic area where the companies are headquartered at year t = 0 and a sample of control firms matched on market value and systematic risk (market beta from the Fama-French 3 Factor model).

Figure 2.3: Mean spread

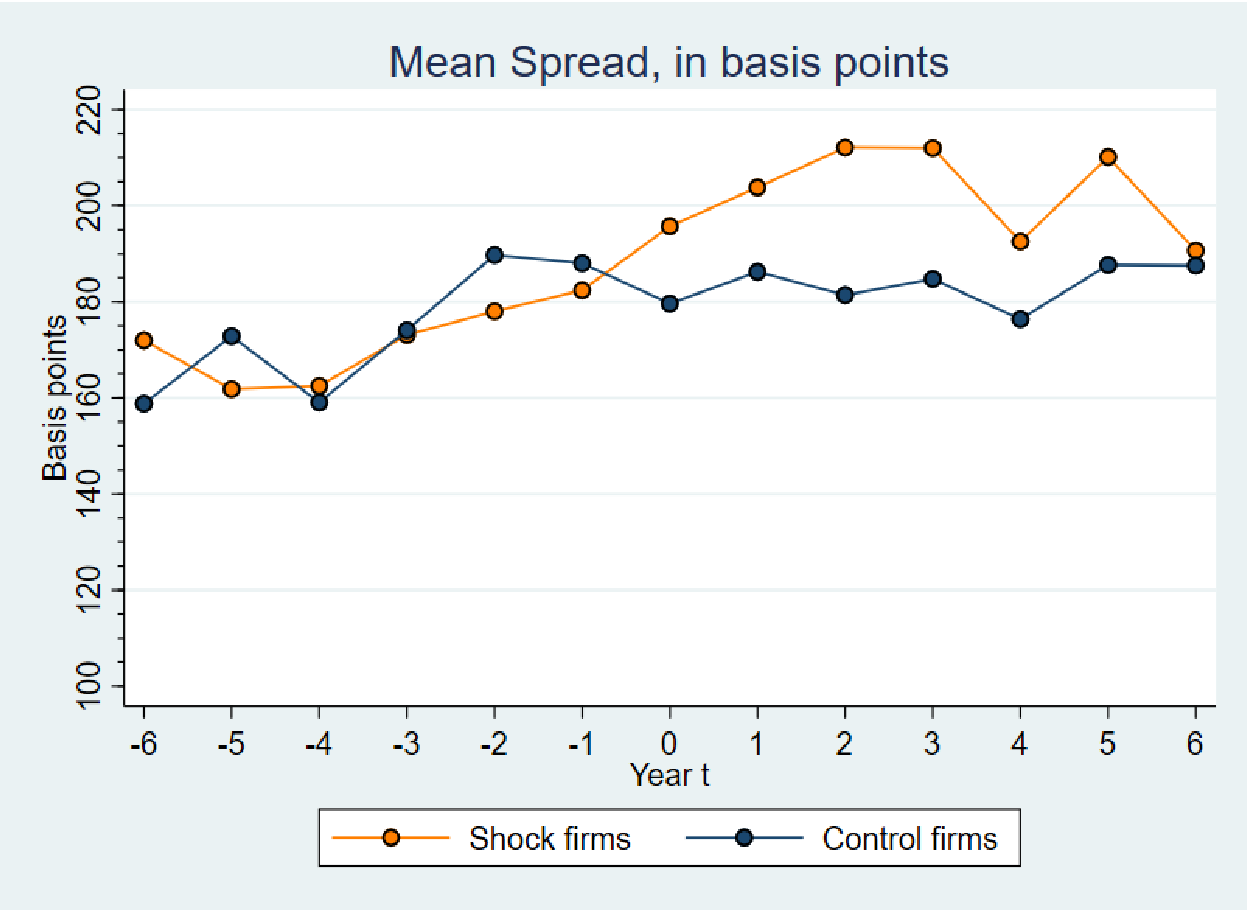

Figure 3: Mean spread from facility loans reported in DealScan for firms that have recorded a 25% or more reduction in local media employment in the geographic area where the companies are headquartered at year t = 0 and a sample of control firms matched on market value and systematic risk (market beta from the Fama-French 3 Factor model).

Figure 2.4: Median spread

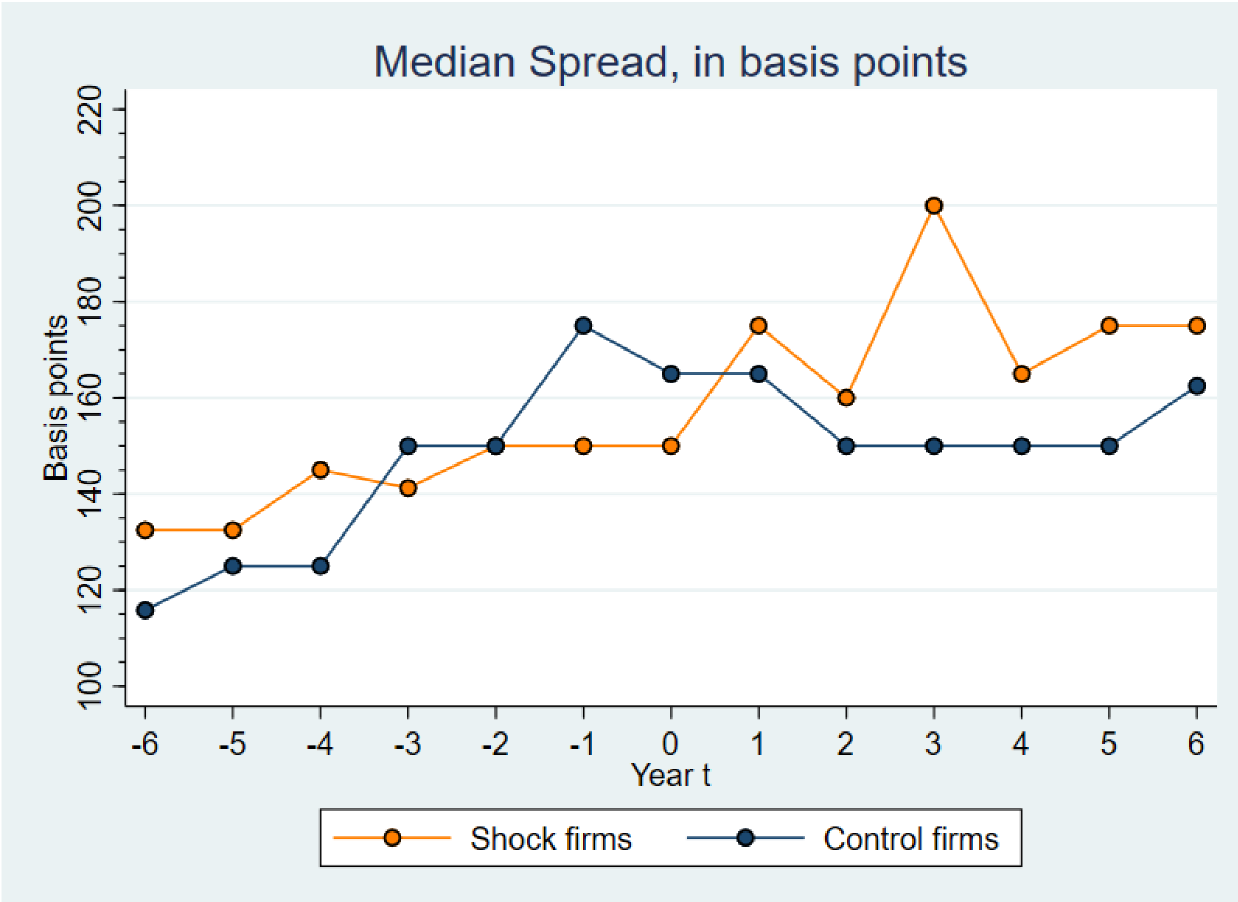

Figure 4: Median spread from facility loans reported in DealScan for firms that have recorded a 25% or more reduction in local media employment in the geographic area where the companies are headquartered at year t = 0 and a sample of control firms matched on market value and systematic risk (market beta from the Fama-French 3 Factor model).

Figure 2.5: Average debt ratios

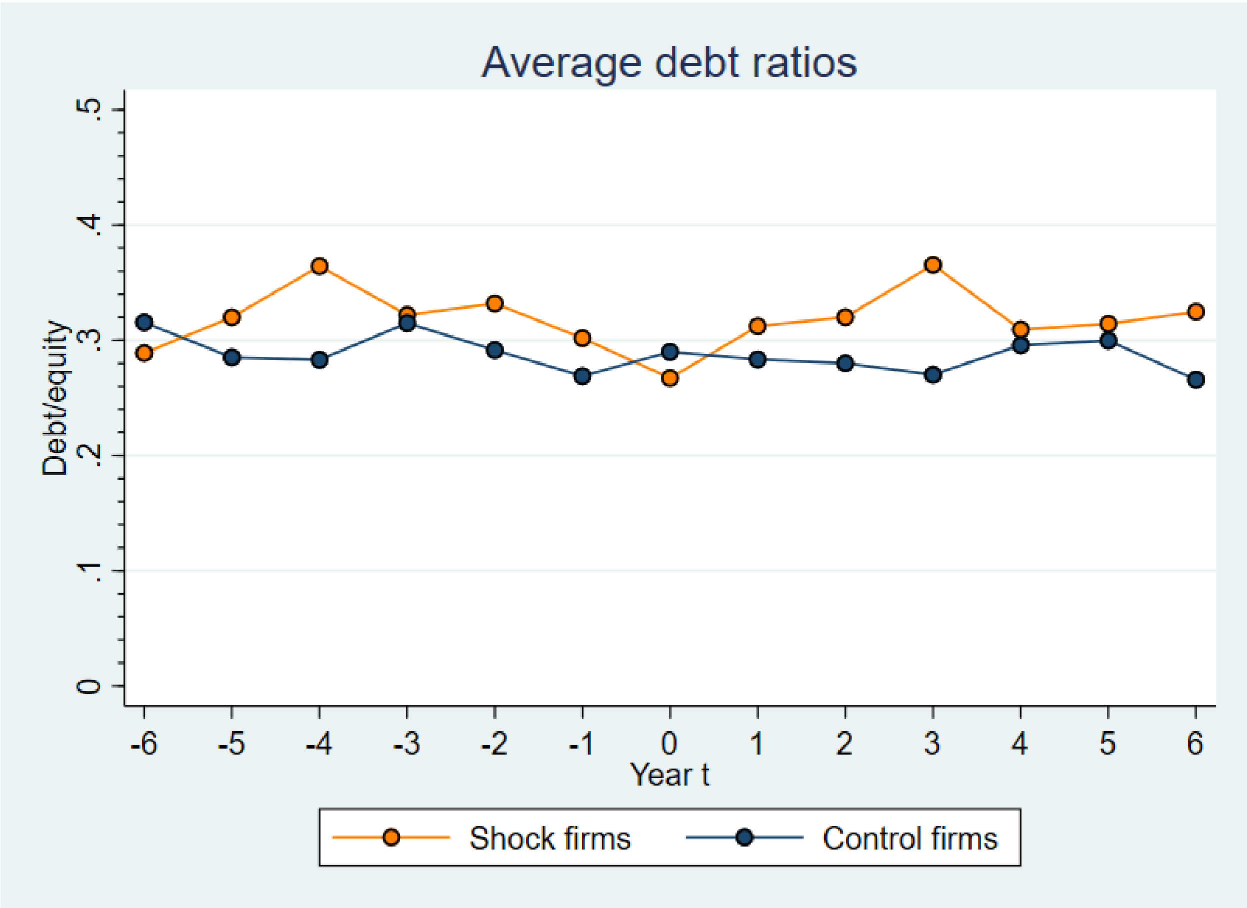

Figure 5: Mean debt ratios for firms that have recorded a 25% or more reduction in local media employment in the geographic area where the companies are headquartered at year t = 0 and a sample of control firms matched on market value and systematic risk (market beta from the Fama-French 3 Factor model).

Chapter 3: Effect of local media reductions on firm meet-beat behavior, CEO turnover, and CEO compensation

INTRODUCTION

Theory and research suggest that media may play a monitoring role to constrain corporate activity[21], and journalists believe that one of their most important functions is uncovering mismanagement and financial malfeasance (Call *et al.*, 2018). However, few have examined the role of local media specifically, and what research does exist suggests that local media may be more important for raising attention than conveying price relevant information about firms (Barber and Odean, 2008; Engelberg and Parsons, 2011) or may even serve as marketers more than monitors because of advertising relationships (Gurun and Butler, 2012).

I examine metropolitan areas in the United States that have experienced a reduction in local media employment, according to an annual survey conducted by the Bureau of Labor Statistics. Firms headquartered in those geographic areas show an increased propensity to report earnings per share that demonstrate discontinuity, consistent with earnings management to meet or beat consensus analyst estimates, year-ago EPS, and reporting zero or just-positive earnings. I also find abnormal discretionary accruals, commonly employed as a measurement for managed earnings (Jones, 1991; Dechow, Sloan and Sweeney, 1995; Ball, 2013), increase for firms headquartered in areas that experience a reduction in media employment when compared with a matching set of control firms.

In addition to behavior consistent with earnings management, I find some evidence that in the years after a reduction in local media employment, CEOs are less likely to experience turnover

[21] Examples include Dyck and Zingales (2002); Miller (2006); Core, Guay and Larcker (2008); Dyck, Volchkova and Zingales (2008); Joe, Louis and Robinson (2009); Dyck, Morse and Zingales (2010); Kuhnen and Niessen (2012); and Liu and McConnell (2013).

and more likely to have higher levels of pay and percentage change increases. The change in pay appears to be driven by profiting on stock and options, consistent with managers using their inside information to profit after a reduction in monitoring and reputational risk that could accompany a reduction in media employment.

MOTIVATION

Healy and Palepu (2001) count media among the information intermediaries — which also include financial analysts, rating agencies and auditors — that engage in private information production to uncover managerial misuse of firm resources and thereby mitigate agency problems that arise from the separation of ownership and management (Jensen and Meckling, 1976). Dyck, Morse and Zingales (2010) examine where corporate fraud allegations originate and find that the media is one of the most important actors in detecting fraud, accounting for more eventual investigations than the SEC, auditors, private litigation, or debt and equity holders.

Miller (2006) finds that the press fills a dual role in revealing accounting fraud, both rebroadcasting information from analysts, auditors, and lawsuits and by providing new information and analysis that uncover accounting irregularities. Dai, Parwada, and Zhang (2015) also find media play a role in corporate governance by disseminating news about insider trading profits, which effectively reduces both overall trading and future profits by corporate insiders. Rogers, Skinner, and Zechman (2016) also study insider trading information and find media play an important role in market reaction to information, even when that information is publically available before media reporting. More broadly, Li, Ramesh and Shen (2011) find that newswires help investors identify news and trade on that information even though SEC reports were previously available to the public.

Nyhan and Reifler (2015) find the threat of fact-checking can constrain lawmakers' willingness to engage in potential falsehoods, and several researchers have found a willingness among firms to change their behavior, as well, to manage their reputational capital in the face of media scrutiny (Dyck, Volchkova and Zingales, 2008; Liu and McConnell, 2013; Baloria and Heese, 2018). Dyck, Volchkova and Zingales (2008) find that media coverage can encourage firms to reverse corporate governance violations. Baloria and Heese (2018) find that firms are willing to delay the release of negative information if they fear the loss of reputational capital from slanted news sources. And Liu and McConnell (2013) find that differences in media tone and attention can affect managers' decision to abandon value-reducing acquisition attempts. Liu and McConnell (2013) conclude that the risk to managers' reputational capital levied by media exposure can help align agent and shareholder interests. Niessner and So (2018) demonstrate that the media prioritizes publicizing negative news about firms, consistent with journalists' priorities as detailed in the Call *et al.'s* (2018) survey of members of the press.

Collectively, the above research demonstrates that media, considered broadly, can play an important role in constraining financial mismanagement through the threat of exposure. However, most prior research focuses on national or international news sources, such as the *Wall Street Journal* (Farrell and Whidbee, 2002); Dow Jones news releases (Li, Ramesh and Shen, 2011; Dai, Parwada and Zhang, 2015); or major broadcast networks such as Fox News (Baloria and Heese, 2018). It is unclear whether local media exert similar reputational risks to managers and firms. Local and regional newspapers appear to drive retail investor trading activity (Barber and Odean, 2008; Engelberg and Parsons, 2011); however, little evidence exists that local media serve as effective corporate monitors, even if their geographic proximity to managers and employees provides added opportunities to detect and reveal financial mismanagement (Gurun and Butler, 2012).

Gao, Lee, and Murphy (2019) show that a reduced threat of local media exposure can mean increased borrowing costs for municipal governments. Consistent with Gao, Lee, and Murphy, Essay 2 demonstrates a similar effect among corporate borrowers. In Essay 1, I demonstrate that firms experience increased idiosyncratic risk after reductions in local media employment, and find evidence consistent with increased levels of information asymmetry driving the change rather than reductions in attention or changes to the competitive environment.

HYPOTHESIS DEVELOPMENT

Despite changes to the media landscape, in which a variety of online and other digital sources may serve as substitutes for traditional journalists, Waldman (2011) and Cage, Herve and Viaud (2019) find the most information still originates with local and regional newspapers and broadcast networks. If even the threat of exposure of evidence of agency conflicts constrains managers' ability or willingness to engage in such activity, then a reduction in local media employment may be associated with increased levels of behavior consistent with financial mismanagement. In the auditor fraud triangle (Creesey, 1973), the risk of fraud is associated with three conditions: perceived financial pressure, rationalizations for engaging in potentially fraudulent activity; and the perceived opportunity to avoid detection. A reduction in the number of local media members available to monitor a firm likely impairs their ability to effectively discover information about a firm and could potentially increase the "perceived opportunity" by managers to engage in behavior that is misaligned with shareholder interests.

Earnings management

Degeorge, Patel and Zeckhauser (1999) and Healy (1985) detail many reasons managers may have incentives to manage earnings, including an array of employment opportunities and compensation benefits. Graham, Harvey and Rajgopal (2005) survey executives and find that managers set smooth earnings reports as a high priority. Further, they find that a majority of managers are willing to destroy firm value to achieve favorable earnings reports. Leuz, Nanda and Wysocki (2003) suggest that managers engage in earnings management to protect their benefits of private control.

If a significant number of managers manipulate earnings to just avoid reporting a loss, then the earnings distribution will be discontinuous at zero, with unusually many few small losses and unusually many small profits. If some managers just avoid year-over-year earnings decreases, then a similar discontinuity arises for earnings changes (Burgstahler and Dichev, 1997). Managers may also avoid just missing a consensus analyst forecast (Degeorge, Patel and Zeckhauser, 1999).

In the context of the fraud triangle, managers may interpret a reduction in local media employment as an increased opportunity to avoid detection of earnings management.

Hypothesis 1: Discontinuity around standard earnings benchmarks will increase for firms headquartered in areas that experience a reduction in nearby media employment.

Discretionary accruals

I next explore one of the common mechanisms for managing earnings, the use of accruals to temporarily boost or reduce reported income. Accruals are components of earnings that are not reflected in current cash flows, and a great deal of managerial discretion goes into their construction. Beneish and Vargus (2002) demonstrate that abnormal accruals can predict insider trading activity by managers, and this evidence of earnings management at least partially explains the accrual anomaly documented in Sloan (1996) and Collins and Hribar (2000). Bergstresser and Philippon (2006) provide evidence that abnormal discretionary accruals are more pronounced for firms with CEOs who have greater compensation incentives to meet earnings benchmarks. As with the meet-or-just-beat earnings incentives, if a reduction in local media provides executives an increased opportunity to adjust accruals with reduced perceived risk of detection, then increased measurements of abnormal accruals would be positively correlated with being headquartered in an area that experienced a media employment reduction.

Hypothesis 2: Firms headquartered in areas that experience a reduction in media employment will demonstrate higher relative levels of discretionary accruals.

Executive turnover

Lowenstein (1996) argues that the presence of potential media coverage can encourage corporate boards to be more effective because of the threat that shareholders might respond to negative press coverage by selling their shares, thereby reducing market value. Negative media coverage of firm performance could also affect director reputations and create incentives for directors to remove the CEO in an effort to salvage their reputations (Farrell and Whidbee, 2002). Miller (2006); Dyck, Volchkova and Zingales (2008); and Dyck, Morse and Zingales (2010) find that media coverage can expose managerial and governance problems at firms. Considering the relationship between CEO turnover and media coverage, Farrell and Whidbee (2002) find that the volume of negative coverage in the Wall Street Journal is correlated with an increased probability of CEO turnover.

If the threat of media exposure of financial mismanagement and the reputational risks to managers and directors is lower after a reduction in local media employment, then CEOs may face less likelihood of being fired or forced to resign.

Hypothesis 3: The probability of CEO turnover is reduced after a reduction in local media employment in the area where the firm is headquartered.

Compensation

Kuhnen and Niessen (2012) find that media coverage of executive compensation can affect both compensation levels and structure. They find that negative press coverage focuses especially on stock options, and that reductions in option compensation are more severe when managers and directors face higher levels of reputational risks. Core, Guay and Larcker (2008) find that the press monitors excess compensation, and that it is more likely to focus on large stock and option elements of executive compensation. However, they find that firms do not change compensation in response to press coverage. Dai, Parwada and Zhang (2015) find that disseminating news about insiders' trading activity can effectively constrain both the volume and the profitability of future trading. If a reduction in local media decreases a potential channel for disseminating information about levels of compensation and trading activity, then managers of firms headquartered in areas that have experienced a reduction in the number of journalists able to disseminate that information may be able to more effectively increase their compensation levels and profitability of trading. This leads to hypotheses 4 and 5:

Hypothesis 4: CEOs of firms headquartered in areas that have experienced a reduction in media employment will see relatively higher levels of compensation than chief executives of a matching control sample.

Bhojraj *et al.* (2009) find that CEOs of firms that exhibit meet-or-just-beat behavior consistent with earnings management and those with poor-quality accruals are more likely to engage in insider selling of their company's stock and options because they understand the potentially firm-value-destroying nature of their actions. If managers face reduced monitoring and reputational risk after a reduction in nearby media employment, levels and changes in pay should be driven by sales of stock and options.

Hypothesis 5: Increased executive pay for firms that have experienced a reduction in media employment in their area will be reflected more in calculations of compensation that include actual profits from stock and option sales rather than estimates of value reported by the firms to the SEC.

DATA AND DESCRIPTIVE STATISTICS

Bureau of Labor Statistics data

Information on the number of reporters and correspondents in an area is taken from the Bureau of Labor Statistics' Occupational Employment Statistics program. The survey produces annual estimates of employment for 810 specific occupations in more than 580 areas, including metropolitan statistical areas (MSAs) and nonmetropolitan areas throughout the U.S. states, the

District of Columbia, and U.S. territories. The statistics are compiled annually after each year's May survey and are released in March of the following year. I use the total numbers of employees in the Reporter and Correspondent categories, which include both the Publishing Industries, such as newspapers and periodicals, and the Radio and Television Broadcasting Industries. I calculate the percentage change to employees in the sum of these two categories across two-year windows to account for the rolling nature of the survey. Years in which the number of employees in the Reporters and Correspondents and Broadcast News Analyst categories falls 25 percent or more from the OES survey released two years prior are identified as negative shocks to coverage.[22] Observations of shocks to employment are limited to 2003, as measured by the difference from 2001, and beyond.[23]

The Metropolitan Statistical Areas used by the BLS are designated by the U.S. Census Bureau. I link ZIP codes to MSAs using Census Bureau designations for 2010, and identify the ZIP code of a firm's headquarters using Compustat listings for its mailing address, because many companies incorporate outside of the state where their operational headquarters may be located for tax and governance reasons. To ensure that total changes in overall employment are not affecting the results, I drop observations in which a 25% reduction in the sum of Reporters and Correspondents and Broadcast News Analysts is accompanied by a reduction in overall employment as recorded by the OES. Bhojraj, Lee and Oler (2003) find that GICS classifications explain stock return comovements better than SIC and NAICS industry classification systems, and

[22] The 25% cutoff is used because it closely approximates the employment reductions explored in Gao, Lee, and Murphy (2019) when one of a county's newspapers closes. Two of the examples mentioned in Gao, Lee, and Murphy (2019) are the closure of the *Denver Rocky Mountain News* in Colorado and the *Cincinnati Post* in Ohio. Those closures were reflected as 28% and 27%, respectively, losses in media category employment in the BLS data. Testing other cutoff levels of 20% and 30% yield similar results to those reported here.

[23] In 2003 and 2004, the OES was released twice, in May and November. I use the data released for the May survey results in those years for consistency. In 2004, the number of Reporters and Correspondents and Broadcast News Analysts were the same in both surveys.

Levi and Welch (2017) find that a firm market value strongly correlates with a firm's beta and other market model factor similarities when compared with other firms. I find exact matches for each sample firm by year and by 6-digit G industry code, and then match, without replacement, by market value within 15% and within 15% of the beta of the treated firm's calculated beta in the Fama-French Three Factor model regressions. The Fama-French Three Factor model is:

$$R_{i,t} - R_{f,t} = \alpha_i + \beta_i^M\left(R_t^M - R_{f,t}\right) + \beta_t^{SMB} R_t^{SMB} + \beta_i^{HML} R_t^{HML} + \varepsilon_{i,t}$$

Where $R_{i,t}$ is the return to the firm *i* at time *t*, $R_{f,t}$ is the risk-free rate, R_t^M is the return to a value-weighted market portfolio, R_t^{SMB} is the difference of the return to small market value stocks minus larger firms, and R_t^{HML} is the difference between the return to high book-to-market stocks and low book-to-market firms. Regressions use daily returns for firms and the value-weighted market portfolio from the CRSP database, and daily returns to the small-minus-big and the high-minus-low portfolios, R_t^{SMB} and R_t^{HML}, are from the factor file on Kenneth French's website.

I eliminate firms with relevant missing control variable data from Compustat or CRSP and firms with common stock share prices that closed the year trading under $2 to mitigate the effects of the smallest value firms affecting results.

My final sample consists of 1,373 firms headquartered in locations that experienced media employment reductions and their 1,373 controls in areas that have not.

Earnings

I use three earnings benchmarks to evaluate meet-or-just-beat behavior. The first examines only those firms in the sample that are covered by at least one financial analyst and calculates analyst forecast error as:

$$FERR_{i,t} = EPS_{i,t} - Forecast_{i,t-1}$$

Where $EPS_{i,t}$ is the reported annual earnings per share (EPS) for firm *i* for the year *t,* and $Forecast_{i,t-1}$ is the mean of all analyst forecasts during the three-month period before the end of the fiscal year. When $FERR_{i,t} = \$0$, a firm's reported EPS matched the consensus forecast exactly. The statistical test uses bins set at a width of $0.0025 (a quarter of a cent) to determine the distribution discontinuity of reported earnings around the benchmark $FERR_{i,t} = \$0$. A large discontinuity around the benchmark is the probability that earnings were managed to meet or just beat the benchmark (Byzalov and Basu, 2019). The statistical test considers a subsample of bins on either side of the benchmark to determine the parameters of the "normal" distribution of earnings around the benchmark. Observations, as reported in Tables 1 and 2, are not firm-years but the number of firms that reported earnings within the subsample considered.

The second measure is employed for firms that are not covered by financial analysts. It considers the difference in EPS reported in year *t* from year *t-1.*

$$\Delta EPS_{i,t} = EPS_{i,t} - EPS_{i,t-1}$$

The final measurement, used for all treated and control firms, examines the discontinuity distribution around a reported net income of zero and is scaled by market value of the firm, following Burgstahler and Chuk's (2015) recommendations for discontinuity tests.

$$Earnings_{i,t} = \frac{NI_{i,t}}{MV_{i,t}}$$

Where $\boldsymbol{NI_{i,t}}$ is net income for firm *i* at year t and $\boldsymbol{MV_{i,t}}$ is market value (number of common shares outstanding multiplied by share price) of firm *i* at year *t.*

Mean analyst EPS forecasts are taken from the Summary file in the IBES database; actual EPS, net income, and market value are taken from the Compustat database.

Discretionary accruals

I employ two well-established measures of discretionary accruals, the Jones model (Jones, 1991) and the modified Jones model proposed in Dechow, Sloan and Sweeney (1995). TA is defined as the change in non-cash current assets minus the change in current liabilities minus depreciation and amortization, scaled by lagged total assets. The Jones model measures discretionary accruals using the following regression estimated cross-sectionally each year for all firm-year observations in the same two-digit GICS code:

$$TA_{i,t} = \beta_0 + \beta_1 \left(\frac{1}{Assets_{i,t-1}} \right) + \beta_2 \Delta Sales_{i,t} + \beta_3 PPE_{i,t} + \beta_4 ROA_{i,t} + \varepsilon_{i,t}$$

Where $\Delta Sales_{i,t}$ is change in sales and $PPE_{i,t}$ is net property, plant, and equipment. Both values are scaled by lagged total assets ($Assets_{i,t-1}$) to mitigate heteroscedasticity in the residuals. I follow Kothari, Leone, and Wasley (2015) and estimate the regression with the intercept, β_0, to provide an additional control for heteroscedasticity and to mitigate problems stemming from an omitted size variable. Kothari, Leone, and Wasley (2015) also recommend the addition of the performance control $ROA_{i,t}$ because firms experiencing extreme performance may exhibit higher levels of "normal" discretionary accruals. Abnormal discretionary accruals are measured as the absolute value of the residual, $\varepsilon_{i,t}$, from the equation.

The modified Jones model employs the same equation except a firm's change in accounts receivable ($\Delta AR_{i,t}$) is subtracted from $\Delta Sales_{i,t}$ before estimation.

All values are taken from the Compustat database, as are the additional control variables employed in the dynamic difference-in-differences regression, market-to-book and the log-transformed market value of the firm.

Executive turnover

The dependent variable in the probit model of executive turnover takes a value of 1 if the Execucomp database indicates a CEO left the firm for a reason other than retirement in the three years after the reduction in local media employment is recorded (year $t = 0$) for firms in the treated sample or a matched control firm. Data from Execucomp and Compustat is available for a total of 751 firms, with 362 of them having experienced a reduction in local media employment in the area where they are headquartered. In addition to firm and return characteristics, I use CEO age and tenure as controls, both of which are taken from the Execucomp database.

Executive compensation

I use two measures of executive compensation from the Execucomp database, Total SEC and Total compensation — Alternate Method 2. Total SEC is taken from firm filings with the SEC on the overall level of payment to executives and includes salary, bonus, stock awards, option awards, nonequity incentives, pension changes, and other compensation. Alternate Method 2 uses most of the elements of the Total SEC compensation except that stock and option awards are valued using the value realized from option exercise or stock vesting instead of the amount charged to the income statement in filings to the SEC.

METHODOLOGY

Earnings

Early research on earnings discontinuity focused on the empirical histogram of the bins around a theorized target (Burgstahler and Dichev, 1997; Degeorge, Patel and Zeckhauser, 1999; Burgstahler and Eames, 2003), which employs standardized difference tests and cannot easily incorporate multiple explanatory variables. To study the determinants of meet-or-just-beat behavior, researchers have generally employed a logit model that assigns a dummy variable of 1 or 0 based on whether an observation occurs at a particular bin (i.e., around zero or at round number such as 1 or 10 cents) of interest (Frankel, Johnson and Nelson, 2002; Matsumoto, 2002; Ashbaugh, LaFond and Mayhew, 2003; Cheng and Warfield, 2005; Jiang, Petroni and Wang, 2010). However, Byzalov and Basu (2019) argue that the logit model can yield erroneous inferences about the determinants of meet-or-just-beat behavior. They suggest that if a determinant affects the mean or variance of *pre-managed earnings,* then the probability of unmanaged small profits also varies with that determinant. The small- or zero-profit dummy variable employed in the logit model will include both managed and unmanaged earnings. Therefore, the probability that a reported earnings number will be assigned to the bin of interest varies with the determinant, even if the determinant does not affect meet-or-just-beat behavior. Byzalov and Basu (2019) develop a statistical test that allows the distribution shape to vary with multiple explanatory variables, by assuming a smooth distribution of pre-managed earnings and a discontinuous incremental effect at the benchmark of interest. They use local polynomial approximations to model the smooth pre-managed distribution and interact the polynomial terms with explanatory variables to implement the conditioning on determinants. The data outside the small-loss and small-profit intervals identify the pre-managed distribution conditional on the determinants, and the missing small losses or increased small profits identify meet-or-just beat behavior. Employing this method allows distribution discontinuity and its determinants to be identified with OLS regressions in each stage of the estimation. The first stage estimates parameters outside the bins of interest, and the second stage tests observations inside the bins of interest for distribution discontinuity.

Byzalov and Basu (2019) suggest using a third-order (cubic) polynomial[24] and their empirical tests demonstrate the bins from -1 cent to +1 cent are most suitable for examining discontinuity distributions The parameters estimated are an estimated intercept, α_0; a linear trend, α_1; a quadratic trend, α_2; and the cubic trend, α_3. The earnings management probability is calculated as π_0. I employ the Byzalov and Basu (2019) tests for distribution discontinuity using three settings. In the first test, I use the portion of the sample that is covered by analysts to assess meet-or-just-beat behavior of mean analyst EPS forecasts. I conduct separate tests on the treated sample of firms before (at year $t < -1$) and after (at year $t \geq -1$) a reduction in local media employment. I also examine the behavior of firms in the control sample. For firms that are not covered by analysts, I create the same subsamples and test the difference between current year reported EPS and prior-year EPS. These tests are unrestricted and do not control for possible determinants of meet-or-just-beat behavior. Following Byzalov and Basu (2019), I set the bin widths at 0.0025 and examine the 16 bins on either side of 0, [-0.04 to 0.04), to establish the first stage parameter estimates, and the 8 bins around 0 [-0.01 to 0.01) for the probability of discontinuity.

In my final test, I examine discontinuities around zero reported earnings[25]. This test uses the full sample of treated and matched control firms, and includes a number of controls that prior

[24] According to Byzalov and Basu (2019), cubic terms are often significant and improve approximation quality, while higher order terms are consistently insignificant in their explanatory power for the distribution continuity.
[25] Earnings is net income scaled by market value to control for size differences in the bins, as recommended by Burgstahler and Chuk (2015)

research finds are associated with meet-or-just-beat behavior. Following Burgstahler and Dichev (1997) I control for current asset (CA) intensity and current liability (CL) intensity as proxies for a firm's ability to manage earnings by manipulating working capital. Burgstahler and Chuk (2017) also suggest intensity of costs of goods sold (COGS) and research and development (RD) are implicit claims that could create contracting incentives for earnings management. CA intensity is the ratio of non-cash current assets to the market value of equity and CL intensity is the ratio of current liabilities to the market value of equity. COGS intensity is the ratio of cost of goods sold to total assets, and RD intensity is the ratio of R&D expense to total assets. I replace missing R&D expenses with zero.

Discretionary accruals

The change in abnormal discretionary accruals is evaluated using a dynamic difference-in-differences model with abnormal discretionary accruals calculated from the Jones (1991) and the modified Jones model (Dechow, Sloan, and Sweeney, 1995) as the dependent variable:

$$DA_{\,i,t} = \sum_{k\geq 1} \beta_k \, Reduction\ in\ coverage\ has\ been\ in\ effect\ for\ k\ periods_{s,t} + \sum_{s} State\ fixed\ effects_s + \sum_{t} Time\ fixed\ effects_t + \beta_{i,t} Controls + \varepsilon_{i,s,t} \quad (1)$$

All standard errors, in model (1) and all other regression models employed in this paper, are robust to heteroscedasticity and clustered by OES area number (the level of assignment of treatment) to address possible serial correlation in the error terms (Bertrand, Duflo and Mullainathan, 2004). The model includes state- and year-fixed effects. Additional controls include the market-to-book value of the firm as an additional control for growth opportunities and size as measured by the log-transformed market value of the firm. In separate tests (Table 3, Columns 3 and 4), I include other indicators of the firms information environment, a dummy variable equaling 1 if a firm has at least one analyst covering the firm (*Analyst Coverage*) and a dummy variable equaling 1 if the firm is listed on the S&P 1500 index (*S&P 1500*).

The dynamic difference-in-differences controls also include the original independent variables used to estimate the residuals in the Jones and modified Jones models. Chen, Hribar and Melessa (2018) find that the typical implementation of the Jones and modified Jones models that use residuals as a dependent variable generates biased coefficients and standard errors that can lead to incorrect inferences. Because the magnitude of the bias in coefficients and standard errors is a function of the correlations between model regressors, they find that including the independent variables from the original discretionary accrual estimation can correct the bias. Therefore, I include $\Delta Sales_{i,t}$ (or $\Delta Sales_{i,t}$ less $\Delta AR_{i,t}$, in the modified Jones tests) and $PPE_{i,t}$, scaled by lagged total assets, as well as the reciprocal of total assets as additional controls. I also include the Kothari, Leone, and Wasley (2015) performance control, $ROA_{i,t}$ used in the original abnormal discretionary accruals estimation.

Executive Turnover

I estimate a probit model for treated and control firms at year t = 0, where the dependent variable, *Turn*, takes a value of 1 if the CEO is replaced for any reason other than retirement in the following three years, according to the Execucomp database.

$$\begin{aligned} Turn_{i,t=1-3} = \beta_0 + \ & \beta_1 Treat_i + \beta_2 Analyst\ Coverage_i + \beta_3 S\&PIndex_i + \beta_4 BTM_i \\ & + \beta_5 Annual\ Returns_i + \beta_6 Leverage_i + \beta_7 ROA_i + \beta_8 TA_i + \beta_9 CEO\ age_i \\ & + \beta_{10} CEO\ tenure_i + \ \varepsilon_{i,t} \end{aligned}$$

Where the independent variable of interest *Treat* is a dummy variable equal to 1 if a firm recorded a reduction in local media employment at year t = 0. *Analyst Coverage* is a dummy variable equal to 1 if a firm has at least one analyst following it, and *S&P Index* is a dummy variable equal to 1 if the firm is listed in the S&P 1500 index. Both variables are intended to capture some of the other information environment around the firm and are likely related to board and managerial reputational risk. *Annual Returns* are the realized returns for the firm at year *t* = 0 minus the value-weighted return to the market portfolio recorded in CRSP (Weisbach, 2001). *ROA* is net income divided by total assets. Both variables are intended to control for past performance. Book-to-market (*BTM*) also controls for performance, as well as size, and total assets (*TA*) controls for size as measured by book value of the firm. *Leverage* is total debt over total assets and can be indicative of the bankruptcy risk of the firm; Strebulaev and Yang (2013) also find that firm debt levels are correlated with CEO characteristics. *CEO age* and *CEO tenure* control for CEO characteristics that may be associated with turnover and are calculated as described previously.

Executive compensation

My final test examines CEO compensation between treated and control firms in the period after a reduction in local media employment is recorded for the treated sample.

$$Comp_{i,t} = \beta_0 + \beta_1 Treat_{i,t} + \beta_2 Market\ Value_{i,t} + \beta_3 Annual\ Returns_{i,t-1} + \beta_4 Annual\ Returns_{i,t-2} + \beta_5 S.D.of\ Returns_{i,t-1} + \beta_6 Leverage_i + \beta_7 ROA_i + \beta_{10} CEO\ tenure_i + \varepsilon_{i,t}$$

Where *Treat* is a dummy variable equal to 1 if the firm has experienced a reduction in local media employment. Control variables follow Hwang and Kim (2009) and include market value of the firm, as a control for size; lagged annual returns for two years prior ($Annual\ Returns_{i,t-1}$ and $Annual\ Returns_{i,t-2}$) and the lagged standard deviation of returns ($S.D.of\ Returns_{i,t-1}$), which may affect evaluation of the CEO's performance. *Leverage, ROA,* and *CEO tenure* are as described previously.

RESULTS

Earnings

Table 1 shows the results for the unrestricted tests of discontinuity for actual EPS and consensus analyst forecast and actual EPS to year-ago EPS. The firms that experienced a reduction in media employment show significant probability ($p < 0.01$) of having managed earnings to meet analyst mean forecasts (Column 1)[26]. Significantly, the same firms do not show a significant probability of having managed earnings before the reduction in local media employment (Column 2). Control firms with analyst coverage (Column 3) also show significant probability of having managed earnings but with a lower probability ($p < 0.05$) than post-treatment firms. The tests of meet-or-just-beat behavior against year-ago EPS of firms without analyst coverage indicate significant probability of post-treatment firms' managing their earnings, while the discontinuity distribution is not significant for the same firms before the recorded reduction in media employment nor the sample of control firms. However, the relatively low number of observations in the meet-beat year-ago EPS test may indicate that year-ago EPS is not a particularly relevant benchmark for firms as so few observations are with the -4 cent to +4 cent range. Still, all six columns considered collectively are consistent with firms changing their behavior around reported earnings after a reduction in local media employment.

I use the zero net income benchmark in a full sample analysis with control variables, allowed by the Byzalov and Basu (2019) method. The dummy variable *Treat* is a dummy variable

[26] Results in Columns 1, 2, and 3 are nearly identical when using median analyst forecast as the benchmark rather than mean analyst forecast.

equaling 1 if a firm has experienced a reduction in local media and is in the period year t = -1 to +6 from the time the reduction was recorded at year t = 0. *Treat* is lightly significant ($p < 0.1$) and positively associated with the probability of earnings management. Control variables that have been shown to be significant influences on earnings management in prior research are insignificant, except for R&D Intensity ($p < 0.01$). *Analyst Coverage* is also insignificant, despite analysts theorized role in corporate governance and monitoring.

Discretionary accruals

The measures of abnormal discretionary accruals (Table 3) are highly significant ($p < 0.01$) in all specifications of the dynamic difference-in-differences model for years t = -1 to +1[27]. In the Jones model without the information environment controls (Column 1), year t = +6 is significantly positively correlated with a reduction in media employment ($p < 0.05$), and for the modified Jones model without information environment controls, the year t = + 3 is lightly significant ($p < 0.1$). With the information environment controls, *Analyst Coverage* and *S&P 1500*, the effect on discretionary accruals appear even more persistent, with years t = +3, +5, and +6 showing statistically significant variations in abnormal accruals in the years after a reduction in local media employment. Being included in the S&P 1500 index appears to have a negative correlation with abnormal discretionary accruals, consistent with increased attention to a firm dissuading managers from employing accrual management. However, *Analyst Coverage* is not significant for either measure of discretionary accruals.

[27] Although the media employment reduction is recorded in year t = 0, the actual reduction may have occurred as much as two years earlier, so the difference occurring in year t = -1 is consistent with the effect being driven by employment shock. In tests of parallel trends, I allow variables for each year t = -6 to +6 and no significant effects are detected before year t = -1.

CEO turnover

Being headquartered in an area where local media employment has fallen is lightly correlated ($p < 0.1$) with a decrease in the probability that a CEO will be replaced within the next three years (Table 4). The marginal effect of being headquartered in an area with fewer journalists is 3.3% reduced chance of turnover. The other information environment controls do not appear to play a significant role in the probability of turnover. Only the control for firm size, log of total assets, demonstrates a significant relationship to the probability of CEO turnover. The limited sample size may restrict the power of the test, but the test provides some evidence consistent with managers and members of the board of directors suffering less reputational risk when firms are headquartered in an area with fewer journalists.

CEO compensation

I test both overall levels of compensation (Table 5, Columns 1 and 2) and year-over-year change in compensation (Columns 3 and 4) for treated and control firms in the years after a reduction in media employment (i.e., years t = -1 to +6). The dummy variable is significantly positively correlated with overall levels of both measures of compensation ($p < 0.01$); however, in considering the changes in compensation, only the measure of overall compensation that includes actual value of stock and option sales is significantly correlated with a reduction in local media employment. Collectively, the findings are consistent both with firms being less constrained in their compensation levels and with managers being less constrained in their exercise and trading of shares and options after a reduction in media employment. This does not necessarily demonstrate agency problems, and in fact Jensen and Murphy (1990) argue that executive compensation is suboptimal *because* of sensational media coverage. However, these findings are consistent with media attention, even local media attention, putting constraints on executive compensation.

CONCLUSION

My findings provide evidence consistent with local media providing an element of monitoring and potential reputational risk documented among members of the national and financial press. Managers appear to exhibit behavior consistent with a perceived increased opportunity to engage in potentially value-destroying behavior and appear to face less employment risk despite that. Compensation levels, and specifically changes in profitable exercises of stock and options, are consistent with boards and managers facing less reputational risk for increases in executive compensation.

TABLES

Table 3.1: Unrestricted test for meet-beat analyst forecasts and year-ago EPS

Estimates for Byzalov and Basu (2019) distribution discontinuity test at the just-meet and just-beat mean analyst estimates and year-ago EPS bins for firms that have experienced a reduction in media employment in the areas where the firm is headquartered and matching sample of control firms. Bin widths are set at 0.0025, and discontinuity is tested for the 8 bins around a 0 difference between actual EPS and mean analyst forecast. The estimation interval is [-0.04, 0.04) difference between actual EPS and mean forecasts and actual EPS and year-ago EPS, following recommendations in Byzalov and Basu (2019). α is the polynomial coefficient in the probability function of pre-managed earnings at the intercept (α_0), a linear function (α_1), a quadratic function (α_2) and a cubic function (α_3). π_0 is the earnings management probability for just-meet, just-beat observations. Columns 1 and 4 are the results of the discontinuity test on firms after they have experienced a reduction in local media employment; Columns 2 and 5 are results of the test on the same firms before the reduction is recorded; Columns 3 and 6 show the results for a matched sample of control firms.

	Meet-beat analyst estimates			**Meet-beat year-ago EPS**		
	(1) Post-treatment firms	(2) Pre-treatment firms	(3) Control firms	(4) Post-treatment firms	(5) Pre-treatment firms	(6) Control firms
α_0	-0.007***	-0.003	-0.004**	0.043***	0.012	0.042***
	(0.002)	(0.002)	(0.002)	(0.011)	(0.013)	(0.010)
α_1	-0.125	0.290***	-0.064	-0.999*	-0.541	0.615
	(0.079)	(0.094)	(0.072)	(0.417)	(0.469)	(0.473)
α_2	3.018***	1.650***	2.522***	-1.141	2.176	-1.385
	(0.239)	(0.286)	(0.220)	(1.140)	(1.417)	(0.966)
α_3	2.593***	-1.102	2.135**	9.272*	7.841	-4.500
	(0.829)	(0.905)	(0.677)	(3.965)	(4.474)	(3.967)
π_0	**-3.119*****	**7.957**	**-4.974****	**0.163****	**0.232**	**-0.104**
	(1.001)	**(8.258)**	**(2.092)**	**(0.074)**	**(0.341)**	**(0.126)**
Obs.	2,736	1,629	2,661	143	107	160

Standard errors are in parenthesis
**** p<0.01, ** p<0.05, * p<0.1*

Table 3.2: Positive earnings with controls

Estimates for Byzalov and Basu (2019) discontinuity test for for zero or just-positive earnings. Earnings is net income scaled by market value to control for size differences in bins from [-0.06, 0.06). *Treat* is a dummy variable equal to 1 if a firm has experienced a reduction in media employment in the areas where the firm is headquartered. *Analyst Coverage* is a dummy variable equal to 1 if at least one analyst has issued an earnings forecast for the firm in the year before earnings is reported. Other control variables are intensity of Costs of Goods Sold, R&D, Current Assets, and Current Liabilities. Bin widths are set at 0.0025, and discontinuity is tested for the 16 bins around 0 earnings reported with parameters estimated from the 24 bins on either side of 0 earnings.. α is the polynomial coefficient in the probability function of pre-managed earnings at the intercept (α_0), a linear function (α_1), a quadratic function (α_2) and a cubic function (α_3). π_0 is the earnings management probability for the 16 bins on either side of 0 earnings. The coefficients for the interaction terms between the control variables and the parameters α_0, α_1, α_2, and α_3 are not shown.

	Earnings around 0
α_0	0.022***
	(0.003)
α_1	0.595***
	(0.092)
α_2	0.054
	(0.141)
α_3	-0.835**
	(0.344)
$\boldsymbol{\pi_0}$	-0.039
	(0.060)
π Treat	0.048*
	(0.028)
π Analyst Coverage	-0.066
	(0.053)
π COGS Intensity	-0.039
	(0.028)
π R&D Intensity	0.427***
	(0.120)
π CA Intensity	0.045
	(0.091)
π CL Intensity	-0.027
	(0.125)
Obs	10,086

Standard errors are in parenthesis
**** p<0.01, ** p<0.05, * p<0.1*

Table 3.3: Local media reductions and discretionary accruals

Table 3 shows the results of the correlation between the absolute value of discretionary accruals and years *t* = -1 to +6 when a reduction in local media employment is recorded at year *t* = 0. Columns 1 and 2 show results of discretionary accruals measured by the Jones (1991) model and the modified Jones (1995) model of Dechow, Sloan and Sweeney. Columns 3 and 4 include controls associated with the information environment of the firm.

	(1) Jones model discretionary accruals	(2) Modified Jones model discretionary accruals	(3) Jones model discretionary accruals	(4) Modified Jones model discretionary accruals
Market-to-book	0.010***	0.010***	0.008***	0.008***
	(0.002)	(0.002)	(0.002)	(0.002)
Market value	-0.002	-0.002	0.005	0.005
	(0.003)	(0.003)	(0.003)	(0.003)
ROA	-0.009	-0.007	-0.009*	-0.008
	(0.006)	(0.006)	(0.005)	(0.006)
1/Total assets	0.795***	0.732**	0.832***	0.768***
	(0.291)	(0.283)	(0.283)	(0.275)
Δ Sales	0.006***	0.006***	0.006***	0.006***
	(0.002)	(0.002)	(0.002)	(0.002)
Δ PPE	-0.013	-0.011	-0.018	-0.016
	(0.026)	(0.026)	(0.025)	(0.025)
Year t = -1	0.020***	0.020***	0.020***	0.020***
	(0.006)	(0.007)	(0.006)	(0.006)
Year t = 0	0.024***	0.022***	0.023***	0.021***
	(0.009)	(0.008)	(0.008)	(0.008)
Year t = 1	0.018***	0.017**	0.018***	0.017***
	(0.007)	(0.007)	(0.006)	(0.006)
Year t = 2	0.007	0.006	0.008	0.007
	(0.007)	(0.007)	(0.006)	(0.006)
Year t = 3	0.012	0.014*	0.014*	0.016**
	(0.008)	(0.008)	(0.007)	(0.007)
Year t = 4	0.002	0.002	0.005	0.005
	(0.006)	(0.006)	(0.006)	(0.006)
Year t = 5	0.011	0.007	0.015**	0.011*
	(0.007)	(0.006)	(0.007)	(0.006)
Year t = 6	0.019**	0.011	0.025***	0.016**
	(0.008)	(0.007)	(0.007)	(0.007)
Analyst Coverage			-0.001	-0.001
			(0.005)	(0.005)
S&P 1500			-0.040***	-0.039***
			(0.007)	(0.007)
Obs.	17,357	17,351	17,357	17,351
R-squared	0.213	0.208	0.217	0.213
Adjusted R-squared	0.209	0.205	0.214	0.210
Year fixed effects	Yes	Yes	Yes	Yes
State fixed effects	Yes	Yes	Yes	Yes

Standard errors are in parenthesis
**** p<0.01, ** p<0.05, * p<0.1*

Table 3.4: Probability of executive turnover

Table 4 shows the results of a probit model estimation in which the outcome variable is 1 if a CEO leaves the firm for any reason other than retirement in the three years after year t = 0 and 0 otherwise.

Probability of executive turnover in next 3 years	Coef.	St.Err.	t-value	p-value	[95% Conf	Interval]
Treat	-0.263*	0.149	-1.77	0.077	-0.556	0.029
Analyst Coverage	-0.096	0.248	-0.39	0.698	-0.582	0.390
S&P Index	-0.146	0.182	-0.81	0.421	-0.503	0.210
Book-to-market	0.061	0.129	0.47	0.634	-0.191	0.314
Annual returns	-0.152	0.178	-0.85	0.394	-0.501	0.197
Leverage ratio	-0.208	0.365	-0.57	0.568	-0.923	0.507
ROA	-0.727	0.479	-1.52	0.129	-1.666	0.212
Log of total assets	0.129***	0.049	2.65	0.008	0.034	0.225
CEO age	0.003	0.005	0.67	0.505	-0.006	0.013
CEO tenure	0.019**	0.010	2.02	0.043	0.001	0.038
Intercept	-2.277***	0.413	-5.51	0.000	-3.087	-1.467

Mean dependent variable	0.068	SD dependent var	0.252
Pseudo r-squared	0.056	Number of obs	751
Chi-square	20.724	Prob > chi2	0.023

**** p<0.01, ** p<0.05, * p<0.1*

Table 3.5: Executive compensation of firms with nearby media reduction vs. controls

Table 5 shows results from the OLS regression of CEO compensation and changes in CEO compensation for firms that have experienced a reduction in media employment in the area where the firm is headquartered and a matching set of control firms. Columns 1 and 3 use compensation and changed in compensation reported by the firm to the SEC, and Columns 2 and 4 use a calculation that includes the CEO's actual profits from stock and option sales and exercises.

	(1) Total compensation reported to SEC	(2) Total compensation adjusted for stock, option sales	(3) Δ Total compensation reported to SEC	(4) Δ Total compensation adjusted for stock, option sales
Treatment	2683.903***	9590.898***	-568.722	142.122***
	(724.653)	(2448.840)	(728.485)	(20.572)
One-year lag returns	70.612	79.659	16.353	19.507
	(63.997)	(76.146)	(79.765)	(26.799)
Two-year lag returns	78.417	180.968**	-23.121	-2.228
	(52.082)	(83.677)	(20.243)	(2.877)
Market value	240.299	-243.317	-68.344	57.810***
	(364.393)	(861.846)	(111.662)	(21.869)
Book-to-market	322.432	-159.024	-33.107	-9.719
	(377.626)	(499.411)	(47.630)	(12.946)
Leverage	1849.304*	1131.811	-1197.312	58.143
	(1023.778)	(2352.879)	(1624.598)	(97.806)
Lag st. dev. of returns	440.575	-920.422	-597.316	46.185
	(1141.403)	(2709.048)	(488.897)	(105.814)
ROA	1054.867	-197.409	-502.400	-20.985
	(659.895)	(2490.713)	(561.667)	(55.610)
Tenure	58.827**	496.051***	16.758	-6.882***
	(29.293)	(36.928)	(22.159)	(0.980)
Intercept	-3412.141	2943.716	856.578	-364.897**
	(2628.762)	(8675.958)	(1159.070)	(161.568)
Obs.	3,423	3,416	2,804	2,752
R-squared	0.772	0.562	0.173	0.456
Adjusted R-squared	0.714	0.450	-0.057	0.304

Standard errors are in parenthesis

**** p<0.01, ** p<0.05, * p<0.1*

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