Options listing and cost structure rigidity

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Abstract

We find that firms increase cost structure rigidity (i.e., the proportion of fixed costs relative to variable costs in the cost structure) when having options listed. The effect of options listing is stronger on selling, general, and administrative expenses, which are relatively committed and stable, than on the cost of goods sold, which closely tracks sales revenue. The results are robust to modifications in fixed effects included in the tests, the sample constructions, and the empirical methods used to control for the endogeneity concern in the association between cost structure and options listing. Cross-sectional tests suggest that options listing increases firms' cost structure rigidity by relaxing financial constraints and/or reducing managers' incentive to hoard bad news. In sum, we document how options listing affects the underlying firm's cost management, which reflects its financial flexibility and risk-taking capacity.

Keywords: cost structure rigidity; options listing; financial constraints; bad-news hoarding

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1. Introduction

Understanding cost behavior is important for investors to assess a firm's prospects because costs, particularly operating costs, have significant impacts on firm profitability and firm valuation (Garrison, Noreen, and Brewer, 2012; Horngren, Datar, and Rajan, 2012). An important line of the literature in cost behavior is about the rigidity of cost structure—that is, the sensitivity of a change in operating costs to a change in sales. In general, a firm's cost structure is considered rigid if the firm commits more fixed costs than variable costs to its operations.[1]

In theory, fixed costs substitute for variable costs because firms that rely more on fixed costs are more profitable (Chen, Harford, and Karama, 2019) by avoiding the congestion costs due to unexpectedly high realization of demand for products or services (Banker, Byzalov, and Plehn-Dujowich, 2014). Besides, committing more fixed costs—mainly selling, general, and administrative (SG&A) expenses—allows firms to build up organizational capital to enhance long-term sustainability and therefore improve performance (Lev, Radhakrishnan, and Zhang, 2009; Eisfeldt and Papanikolaou, 2014; Peters and Taylor, 2017).[2] However, similarly to having too much debt in the capital structure, relying too much on fixed costs in operations would result in a sharp decline in earnings during bad times. As a result, managers choose a level of fixed costs, together with financial leverage, to maximize the firm's value by balancing the benefits and costs of fixed costs (Chen, Harford, and Karama, 2019).

Previous studies on cost rigidity have mainly focused on the impacts of product market attributes (e.g., demand uncertainty, contribution margin variance, industry competition, and product fluidity) on cost rigidity (e.g., Banker, Byzalov, and Plehn-Dujowich, 2014; Holzhacker, Krishnan, and Mahlendorf, 2015; Chang, Hall, and Paz, 2021; Chang, Kwok, and Wong, 2024; Chen, Kama, and Lehavy, 2024; Chen, Liang, Yang, and Zhu, 2024). Few studies have examined

[1] Cost structure is also used as a proxy for operating leverage to capture the proportion of fixed costs in total costs.
[2] Chen, Harford, and Kamara (2019), among others, find that the cost of goods sold (COGS) is significantly more sensitive to sales revenue than SG&A expenses, suggesting that SG&A expenses have a significant fixed component.

the impact of financial market activities on cost rigidity. Our study aims to fill this gap in the literature by examining how firms adjust their cost structure and the resources committed to operations in response to the listing of options of their underlying stocks. In particular, we hypothesize that options listing increases cost rigidity of the underlying firm via two channels, as discussed below.

First, low cost and high leverage make options trading ideal for informed investors to profit from trading with uninformed investors (Black, 1975; Diamond and Verrecchia, 1987; Mayhew, Sarin, and Shastri, 1995), which creates incentives for informed investors to collect information (Cao, 1999; Cao, Goyal, Ke, and Zhan, 2024). As options prices and the underlying stock prices are theoretically linked, information revealed in the options markets will immediately influence the underlying stock prices. Therefore, options trading should deliver a sizeable enhancement effect on information flows on top of stock trading (Chakravarty, Gulen, and Mayhew, 2004). By enhancing information flows, options trading mitigates the information asymmetry between managers and external investors, thereby reducing the underlying firm's cost of capital (Naiker, Navissi, and Truong, 2013). A lower cost of capital makes it easier for firms to access external financing, thereby encouraging firms to commit resources and make fixed investments, such as investments in large in-house production facilities or the purchase of equipment.

Second, a transparent information environment motivates managers to make cost decisions that maximize long-term firm value rather than short-term performance. Managers are likely to be myopic when the firm's underlying stock price fails to reflect the long-term value implications of corporate decisions. To avoid being punished by poor short-term operating performance, managers tend to reduce fixed costs. However, cutting fixed costs could sacrifice the firm's long-term prospects by lowering the accumulation of intangible capital. A more transparent information environment, on the other hand, causes investors to evaluate managers based on the long-term implications of corporate decisions (Wurgler, 2000; Durnev, Morck, and Yeung, 2004) rather than the fluctuations of short-term profits. As a result, managers can make cost decisions that maximize long-term firm value rather than short-term profits. Besides promoting information flow, options markets allow investors to trade on negative information at a lower cost than the short selling of underlying stocks (Easley, O'Hara, and Srinivas, 1998; Johnson and So, 2012; Ramachandran and Tayal, 2021). Therefore, the existence of listed options should mitigate managers' incentives to meet short-term profit targets at the expense of long-term development (Blanco and Wehrheim, 2017).

The above argument assumes that options trading is informative. Some studies, however, show that individual investors, who tend to be noise traders, account for a significant portion of options trading volume (Lemmon and Ni, 2014), and options trading provides no additional information in financial markets (Choy and Wei, 2023; Muravyev, Pearson, and Broussard, 2013). If options trading is largely speculative and noisy, then stock prices will be clouded by noisy options prices (Hu, 2018; Roll, Schwartz, and Subrahmanyam, 2010) and become less informative (Stein, 1987). Therefore, after options listing, investors would rely less on the stock price and more on the short-term profit of the underlying firm as a signal for trading. Such short-termism by investors results in over-punishing managers for short-term losses and over-rewarding managers for short-term gains. Moreover, a cloudier information environment heightens financial constraints, increasing the need to reserve financial slack. In response, managers rely more on variable costs than fixed costs. For instance, the use of committed resources is reduced through business outsourcing, equipment leasing, recruitment of temporary workers, and so on.

To test the two competing hypotheses, we collect a sample of U.S.-listed firms in 1996-2022 from the CRSP/Compustat Merged database. After removing firms in financial and utility industries, we obtain a final sample of 43,445 firm-years in the sample period. Cost rigidity is measured as the sensitivity of the annual change in log operating costs to the annual change in log sales. A higher sensitivity indicates a lower cost rigidity. To examine whether the listing of a firm's options affects the firm's cost rigidity, we interact the firm's change in log sales with an indicator for the period after the initial listing of its options in our regression analysis. Our baseline result indicates that options listing is associated with a lower sensitivity of a change in operating costs to a change in sales—that is, a higher cost rigidity. The result is robust to the inclusion of other firm characteristics. The economic significance is that options listing causes the sensitivity of a change in operating costs to a change in sales to reduce by 0.082, or 11.9%, from a base value of 0.689. The finding is consistent with the efficient options trading hypothesis that options listing causes the cost structure to be more rigid.

We then perform several additional analyses to examine our baseline result. First, we show that relatively speaking, the rigidity of SG&A expenses increases more than that of COGS after options listing. Therefore, the change in cost rigidity after options listing mainly comes from substituting fixed SG&A for variable SG&A.

Second, we show that firms are less likely to rely on outsourcing and more likely to internalize operations after the initial listing of their options. Outsourcing provides operational flexibilities by offering firms an opportunity to renegotiate contracts with their suppliers. It also allows firms to free up resources for more productive investments (Choi, Ju, Trigeorgis, and Zhang, 2021) and to respond to demand uncertainty and financial risk (Holzhacker, Krishnan, and Mahlendorf, 2015; Moon and Phillips, 2021) when they are financially constrained. Using the data provided by Moon and Phillips (2021), we show that firms outsource their activities to a lesser extent and that they are more likely to purchase than to lease assets after the initial listing of their options.

Finally, we repeat our tests with the initial listing of options replaced by various measures of options trading volume because previous studies show that options trading volume affects corporate finance and investment decisions (e.g., Blanco and Wehrheim, 2017; Anagnostopoulou, Trigeorgis, and Tsekrekos, 2023; Hsu, Ke, Ma, and Ruan, 2024), cost of capital (Naiker, Navissi, and Truong, 2013), and disclosure (Chen, Ng, and Yang, 2021). Motivated by those studies, we repeat our tests with the initial listing of options replaced by variables of options trading volume, including the raw options trading volume, options trading volume relative to stock trading volume, and an exogenous shock to options trading volume—the Penny Pilot Program. The results with each of these options trading volume variables indicate that cost structure is more rigid when firms have a higher options trading volume.

We also use various ways to address the endogeneity concern of our baseline results. First, we conduct a parallel-trend test to show that there is no significant change in cost rigidity before options listing. Second, following Blanco and Wehrheim (2017) and Hsu, Ke, Ma, and Ruan (2024), we use propensity score matching and an entropy balancing model to control for differences in characteristics between firms with listed options and those without listed options. Our main result is robust to these two empirical strategies. Third, we add high-dimensional fixed effects to the model and show that unknown time-varying effects are unlikely to drive our main findings.

After confirming the robustness of our baseline result, we perform two sets of cross-sectional tests to examine the channels through which options listing affects cost structure. First, to show that options trading alleviates financial constraints and thereby allows firms to incur more fixed costs, we classify firms into groups according to two proxies for information asymmetry, including analyst forecast errors and the bid-ask spread of stock prices. As information asymmetry is a root cause of financial constraints (Kaplan and Zingales, 1997), the impact of options pricing on cost rigidity should be stronger for firms that exhibit higher information asymmetry. The empirical result is consistent with our expectations. We also classify firms into groups according to common proxies for financial constraints, including the SA index (Hadlock and Pierce, 2010), the WW index (Whited and Wu, 2006), the existence of long-term credit rating, and the dividend payout ratio. Again, the impact of options pricing on cost rigidity is stronger for more financially constrained firms.

Second, we test whether options trading alleviates managers' concerns of being excessively punished by short-term poor performance. If options trading helps stock prices to reflect information for product demand of a firm more promptly and accurately, then the firm managers will have less incentive to stabilize earnings by incurring relatively more variable costs than fixed costs. To test this prediction, we classify firms into groups according to a proxy of managers' bad-news hoarding incentive (Kama and Weiss, 2013). The result indicates that the impact of options trading on cost rigidity is larger when managers have stronger incentives to avoid reporting losses. We also classify firms into groups according to proxies for the short-sale constraints of underlying stocks. The idea is that short-sale constraints inhibit stock prices from incorporating negative news, resulting in higher crash risks. To avoid stock prices going bust, managers have incentives to stabilize earnings by adopting a flexible cost structure. A salient feature of options trading is that options trading allows investors to profit from negative information at a lower cost by circumventing the short-sale constraints of underlying stocks. If promoting information flow increases cost rigidity, the impact of options listing on cost rigidity should be stronger for stocks with larger short-sale constraints. We follow Asquith, Pathak, and Ritter (2005) in classifying firms as firms with high short-sale constraints if their stocks have low institutional ownership and high short interest. Consistent with our expectations, we show that the impact of options listing on cost rigidity is larger when the underlying stock's short-sale constraints are larger.

Previous studies on cost structure have linked cost rigidity to demand uncertainty. In particular, when determining their cost structures, firms trade off the downside risk of negative demand shock against reduced congestion cost to meet positive demand surprise (Banker, Byzalov, and Plehn-Dujowich, 2014). To reduce congestion cost to meet positive demand surprise, firms maintain a higher capacity by making more fixed investments—that is, adopting a more rigid cost structure when facing higher demand uncertainty, which tends to be high when firms serve few customers and/or operate in competitive industries (Chang, Hall, and Paz, 2021). Consistent with previous studies (Banker, Byzalov, and Plehn-Dujowich, 2014; Chang, Hall, and Paz, 2021), we find that cost structure is more rigid in firms with higher demand uncertainty, as indicated by higher firm-level sales uncertainty and higher product similarity. More importantly, the result is more pronounced for firms with options listed. Therefore, firms facing greater demand shock commit more fixed investments to maintain a higher capacity, and options listing further strengthens this tendency, probably by improving financial flexibilities for fixed asset investments.

Our study contributes to the literature in several ways. First, although many previous studies have explained how product market factors, such as demand uncertainty, affect firms' cost structure, few have examined how derivatives trading on financial markets affects cost structure.[3] One noticeable exception is Fang, Pu, and Wang (2023), who find that after the inception of CDS trading, reference firms exhibit an increase in the elasticity of cost structure due to increased creditors' liquidation incentives. Our results, on the other hand, show that firms' cost rigidity increases after the initial listing of their options, especially when firms' financial constraints are large in the first place. Besides, the impact of options listing is larger when negative information is more likely to be suppressed. This notable feature of options trading is not commonly available in other proxies for the information environment (Agarwal, Khizer, and Sethuraman, 2023).

Second, our study follows a recent stream of studies on the role of options trading in corporate decision-making.[4] In particular, Blanco and Wehrheim (2017) find that options trading causes the underlying stock price to be more informative. As a result, investors are less likely to overreact to short-term operating performance and make more long-run investments, such as research and development investments and corporate innovation. We provide consistent evidence showing that after options listing, firms undertake a more rigid cost structure, which highlights the information role of options trading in facilitating risk-taking and encouraging managers to be more long-term oriented.

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[3] Prior studies have documented various determinants of cost structure. The most important determinants are product market factors, including demand uncertainty (Kallapur and Eldenburg, 2005; Banker, Byzalov, and Plehn-Dujowich, 2014; Holzhacker, Krishnan, and Mahlendorf, 2015; Chang, Kwok, and Wong, 2024; Chen, Kama, and Lehavy, 2024), firm-level bullwhip effect (Chen, Di, Jiang, and Li, 2024), and concentration degree in customers or suppliers (Chang, Hall, and Paz, 2021; Pizzini and Vansant 2024; Chen, Liang, Yang, and Zhu, 2024). Corporate governance and institutions are also relevant in determining cost structure (Aboody, Levi, and Weiss, 2018; Chang, Xin, Lohwasser, and Chiu, 2022; Siciliano and Weiss, 2023).

[4] Recent empirical studies indicate that options trading can promote information flows in financial markets, thereby reducing the equity cost of capital (Naiker, Navissi, and Truong, 2013), deterring earnings management (Dai, Qiao, and Xia, 2024; Delshadi, Hosseinniakani, and Rezaee, 2023), and enhancing the efficiency of corporate decisions (Bernile, Hu, Li, and Michaely, 2023), including capital investments (Anagnostopoulou, Trigeorgis, and Tsekrekos, 2023; Hsu, Ke, Ma, and Ruan, 2024) and innovation activities (Blanco and Wehrheim, 2017).

Finally, our findings are consistent with a broad line of studies that propose a substitution relationship between financial risk and operating risk (e.g., Lev, 1974; Kumar and Yerramilli, 2018; Chen, Harford, and Kamara, 2019). That is, firms choose a lower operating leverage in response to a higher financial risk. In this regard, Fang, Pu, and Wang (2023) find that firms respond to an increase in financial risk (caused by the empty creditor problem derived from the CDS trading inception) by reducing operation leverage. Our results show that firms increase operating leverage after the initial listing of their options that are expected to reduce their financial risk by improving their information transparency.

2. Literature and Hypotheses

From the theoretical perspective, substituting fixed costs for variable costs makes earnings more sensitive to changes in sales revenue (Garrison, Noreen, and Brewer, 2012; Horngren, Datar, and Rajan, 2012). The increased sensitivity implies an increase in systematic risk and therefore an increase in the expected return of the underlying business.[5] From the operational perspective, benefits from using fixed costs versus variable costs come from reduced adjustment costs to scale up the operations when demand for products/services increases unexpectedly. Consistent with the above predictions, Banker, Byzalov, and Plehn-Dujowich (2014) empirically show that firms facing more uncertainty in demand are more likely to commit to a more rigid cost structure—that is, the one that relies more on fixed costs than variable costs. By doing so, firms can avoid the congestion costs due to an unexpectedly high realization of demand for products or services. More recently, Chen, Harford, and Kamara (2019) document a positive relation between firm profitability and the use of fixed costs.

Recent studies suggest that part of the operating expenses, specifically the SG&A expenses, aim to build up intangible capital and have a significant impact on a firm's long-term value (Peters

[5] The impact of operating leverage on systematic risk has been discussed in classic studies (e.g., Lev, 1974; Gahlon, 1981; Mandelker and Rhee, 1984). In general, according to classic asset pricing theories, higher operating leverage leads to higher systematic risk and therefore higher expected return.

and Taylor, 2017; Banker, Huang, Natarajan, and Zhao, 2019; Iqbal, Rajgopal, Srivastava, and Zhao, 2023). Therefore, while spending on SG&A would reduce corporate profit in the short run, it would improve long-term performance by strengthening the firm's organizational capital and knowledge development. As investments in intangible capital aim to deliver superior long-term performance by combining human skills and physical capital into systems for producing and delivering want-satisfying products (Lev, Radhakrishnan, and Zhang, 2009), these investments are unlikely to be linked to the current level of operations but should be relatively fixed and committed.

The above studies suggest that firms can improve expected performance by incurring more fixed costs. However, the benefits could be offset by potential distress costs that would occur when the product/service demand is significantly weaker than expected. The trade-off becomes more acute when firms face financial constraints. A firm is financially constrained if its external cost of capital is significantly higher than its internal cost of capital. The wedge between external and internal costs of financing is largely driven by information asymmetry between corporate insiders and outside investors. In their classic study, Myers and Majluf (1984) theoretically show that the information asymmetry problem results in adverse selection in financing choices and that as a result, investors require higher expected returns for securities that exhibit more information asymmetry. A key empirical implication is that firms cannot fund all positive-NPV projects; that is, they are financially constrained, because external funds are more costly than internal funds. The financial and real impacts of financial constraints have been widely documented and evaluated by previous studies on corporate investments (e.g., Fazzari, Hubbard, and Petersen, 1988; Hoshi, Koshyap, and Scharfstein, 1991; Kaplan and Zingales, 1997; Alti, 2003), cash holding (Almeida, Campello, and Weisbach, 2004, 2024; Almeida and Campello, 2007; Hadlock and Pierce, 2010), and expected stock returns (Whited and Wu, 2006), among others. To summarize, firms tend to behave more conservatively when they are more financially constrained.

Motivated by previous studies, we expect that financial constraints cause firms to rely less on fixed costs relative to variable costs. When financial constraints become more severe, managers adopt a more flexible cost structure by reducing fixed costs to avoid falling into financial difficulties. When financial constraints are relaxed, fixed costs will substitute for variable costs to achieve a higher expected profit and better long-term performance.

To test the above prediction, we use options listing as an exogenous reduction in information asymmetry and therefore a reduction in financial constraints for the underlying firm. The relatively low transaction costs and high leverage of options attract informed investors to collect information and profit from trades with uninformed investors (Black, 1975; Diamond and Verrecchia, 1987; Mayhew, Sarin, and Shastri, 1995; Skinner, 1997; Cao, 1999; Cao, Goyal, Ke, and Zhan, 2024), which makes that options trading potentially reflects information beyond stock trading. Numerous studies have found that options trading contains important private information about the underlying stock value and can enhance the information content of stock prices (Chakravarty, Gulen, and Mayhew, 2004; Cremers and Weinbaum, 2010; Johnson and So, 2012; An, Ang, Bali, and Cakici, 2014; Hayunga and Lung, 2014; Hu, 2014; Cao, Goyal, Ke, and Zhan, 2024). Therefore, the incremental value-relevant information embedded in options trading can alleviate the information asymmetry between the firm and the market, thereby reducing the firm's cost of capital (Naiker, Navissi, and Truong, 2013) and increasing the firm's access to debt and equity financing (Li, 2021). Lower financial constraints allow managers to take more risk in operating decisions for higher profits and to invest in organizational capital for long-term success.

Furthermore, the role of options trading in mitigating information asymmetry may alter managerial incentives, which can affect a firm's cost decisions. If a firm's cost structure is rigid such that the cost is not sufficiently adjusted downward when sales revenue declines, the firm will experience a significant profit decline or even incur losses. Lacking sufficient information, the market puts more weight on the most recent observable output when revising its beliefs about managers' abilities (Holmström, 1999). Managers' failure to hit short-term earnings targets, such as analyst forecasts, often hinders managers' internal promotions or intra-industry mobility (Graham, Harvey, and Rajgopal, 2005). More importantly, missing earning targets amid declining sales revenue may exacerbate market penalties (Graham, Harvey, and Rajgopal, 2005; Rees and Sivaramakrishnan, 2007), exposing managers to greater career risks. Hence, if financial markets cannot accurately evaluate the long-term implications of managers' SG&A investments, the rational choice of self-interested managers would be to adopt a flexible cost structure to stabilize earnings. With options trading, the increase in value-relevant information flows will enhance investors' assessment of managers' cost decisions. As a result, managers are likely to be rewarded based on informative stock prices and are less likely to be wrongly punished for short-term underperformance (Dow and Gorton, 1997; Kang and Liu, 2008). Shielding managers from excessive career concerns can reduce their excessive focus on current earnings and strengthen their incentives to maintain valuable investments in organizational capital. Therefore, overall cost rigidity increases. We summarize the empirical prediction in the following efficient options trading hypothesis:

H1a: Ceteris paribus, options listing increases firms' cost rigidity.

However, options trading can also be speculative. Previous studies find that options trading is associated with speculative trading by uninformed investors (Wei, Poon, and Zee, 1997; Lemmon and Ni, 2014). Uninformed investors, lacking information advantages, may focus on a firm's short-term profits without considering the long-term value of committed fixed resources. Thus, speculative options trading not only fails to enhance value-relevant information flows in the financial market but also leads investors to misprice managers' cost decisions, exacerbating the information asymmetry between the firm and the market. In an imperfect market, information asymmetry exists and results in risk premiums that increase the cost of capital (Hughes, Liu, and Liu, 2007; Armstrong, Core, Taylor, and Verrecchia, 2011; Lambert, Leuz, and Verrecchia, 2012) and reduce a firm's access to credit and equity market (Dierkens, 1991; Tang, 2009). The resulting financial constraints could lead to a reduction in cost rigidity.

Meanwhile, in response to speculative stock prices, managers tend to pursue short-term goals at the expense of long-term fundamental value (Bolton, Scheinkman, and Xiong, 2006).

Specifically, managers who adopt a more flexible cost structure to boost current earnings may be over-rewarded, while those who commit more fixed resources to the firm's organic growth but incur a decline in current earnings or losses may be over-punished by the market. To avoid unfair market penalties leading to reputation and compensation losses, managers with heightened career concerns would adopt a more conservative and flexible cost structure. Hence, we propose the following speculative options trading hypothesis as a competing hypothesis:

H1b: Ceteris paribus, options listing decreases firms' cost rigidity.

3. Research Design

3.1 Sample and data

We start with annual data from the CRSP-Compustat merged database and obtain options listing data from OptionMetrics. We first exclude samples from the financial industry (SIC code 6000-6999) and utility industry (SIC code 4900-4999). Next, we exclude samples with missing or non-positive total assets, sales revenue, and operating costs, as well as samples in which the stock closing price is less than $1 (Chen, Lu, and Sougiannis, 2012). Following Banker, Byzalov, and Chen (2013), we drop observations with extreme operating costs that are lower than 50% or higher than 200% of sales for the current or prior years.[6] Following Naiker, Navissi, and Truong (2013) and Hsu, Ke, Ma, and Ruan (2024), we define a firm's initial options listing date as the first date on which it appears in the OptionMetrics database. We drop firms that first appear in the OptionMetrics database in 1996 because we cannot distinguish between firms that were first listed in 1996 and firms that were listed before 1996. Further, we drop firms that were first listed at the end of the sample period (i.e., year 2022) and restrict the sample to a time window of 10 years before and after options listing to strengthen causal inference.[7] Finally, we exclude samples with missing values of variables in the baseline model, resulting in a final sample of 43,445 firm-year

[6] Our results remain robust even if we remove these data filtering processes.
[7] Alternatively, if we set this time window to [-3, 3] or [-5, 5] or do not impose any restrictions on the time window at all, our baseline results do not show any adverse changes.

observations covering 43 industries in the Fama-French 48 industries from 1996 to 2022. Table 1 details the year and industry distribution of the final sample. To control for inflation, we convert financial variables into their equivalent 1996 dollar values by applying the GDP deflator. To mitigate the influence of outliers, we winsorize all continuous variables at the bottom and top one percentiles.

[Insert Table 1 here]

3.2 Model specification

Following previous studies (Banker, Byzalov, and Plehn-Dujowich, 2014; Chang, Hall, and Paz, 2021; Chen, Liang, Yang, and Zhu, 2024), we use the following staggered difference-in-differences (DiD) model to detect the effect of options listing on firms' cost structure:

$$\Delta lnOPR_{i,t} = \alpha_0 + \alpha_1 \Delta lnSales_{i,t} + \alpha_2 OPList_{i,t} + \alpha_3 OPList_{i,t} * \Delta lnSales_{i,t} + \alpha_4 Controls_{i,t} + \alpha_5 Controls_{i,t} * \Delta lnSales_{i,t} + Year\ FEs + Firm\ FEs + \varepsilon_{i,t} \quad (1)$$

$\Delta lnOPR_{i,t}$ refers to the log change of operating costs of firm *i* in year *t* relative to year *t*-1. Similarly, $\Delta lnSales$ refers to the log change of sales revenue. *OPList* is a variable that takes the value of 1 in the year of the initial listing of options and all subsequent years and takes the value of 0 otherwise. In model (1), the coefficient α_1 captures firms' cost structure, with a lower α_1 indicating that changes in firms' operating costs are less responsive to variations in sales; that is, the cost is more rigid. Furthermore, the coefficient α_3 captures the effect of options listing on firms' cost structure. If α_3 is significant and negative (positive), it indicates that options listing increases (decreases) firms' cost rigidity.

Controls include firms' basic characteristics that may determine their cost structure in the first place, such as firm size, firm age, and market-to-book value (Chang, Kwok, and Wong, 2024). Following Anderson, Banker, and Janakiraman (2003), we control for the adjustment costs in firms' cost decisions, proxied by asset intensity and employee intensity. In addition, Chen, Lu, and Sougiannis (2012) argue that managerial empire-building incentives may also influence firms' cost behaviors. Therefore, we control for free cash flows to account for possible managerial empire-building incentives. Detailed definitions and data sources of all the variables can be found in Appendix A. In the model, we also control for year fixed effects and firm fixed effects to mitigate endogeneity concerns related to omitted variables. Standard errors in all regressions are clustered at the firm-year level.[8]

3.3 Descriptive statistics

Table 2 presents the descriptive statistics of the main variables in the baseline model. On average, the variation in operating costs is roughly equivalent to the variation in sales (the mean of $\Delta lnOPR$ is slightly higher than the mean of $\Delta lnSales$). The mean value of *OPList* is 0.45, indicating that firm-years before and after options listing are relatively balanced in terms of sample size. In our sample, the average total assets of the firms are $240 million (before log transformation), and the average firm age is 15 years (177 months, before log transformation). On average, the market value of the sample firms is three times their book value. The distributions of asset intensity, employee intensity, and free cash flows are also similar to those in previous studies (e.g., Chang, Xin, Lohwasser, and Chiu, 2022).

[Insert Table 2 here]

4. Empirical Results

4.1 Options listing and cost rigidity

Table 3 reports the baseline regression results. In column 1, we do not include any control variables and control only for firm fixed effects and year fixed effects. The regression coefficient of $\Delta lnSales$ is significant and positive ($p < 0.01$), while the regression coefficient of $OPList*\Delta lnSales$ is significant and negative ($p < 0.05$), indicating that firms' operating costs become less sensitive to sales volatility after options listing. In column 2, we further include all control variables and their interactions with $\Delta lnSales$. The significant and negative ($p < 0.01$)

[8] Regardless of whether we cluster standard errors at the firm, industry, firm-year, or industry-year level, our results remain highly robust.

regression coefficient of *OPList**Δ*lnSales* lends strong support to our hypothesis; that is, options listing increases the underlying firms' cost rigidity. According to the result in column 2, firms' cost rigidity increases by 11.9% (= 0.082/0.689) after the initial listing of their options, which carries both statistical and economic significance.

[Insert Table 3 here]

4.2 Robustness Tests

4.2.1 Alternative dependent variables

In the baseline model, we define cost rigidity as the sensitivity of a change in operating costs to a change in sales. Aboody, Levi, and Weiss (2018) and Chen, Harford, and Kamara (2019) document that although both COGS and SG&A are the core cost components that are expected to respond to revenue shocks, SG&A expenses (i.e., resources consumed for marketing and advertising, distribution, and information technology) have more fixed elements. A significant portion of SG&A expenses aims to build up intangible capital (e.g., Lev, Radhakrishnan, and Zhang, 2009; Eisfeldt and Papanikolaou, 2014; Peters and Taylor, 2017), making SG&A expenses more committed than COGS. Therefore, we separate total operation cost into COGS and SG&A expenses and use Δ*lnCOGS* (the log change of the cost of goods sold) and Δ*lnSG&A* (the log change of SG&A) as separate dependent variables. The regression results in Panel A of Table 4 show that the rigidity of different cost components (either COGS or SG&A) increases after options listing, as evidenced by the significant regression coefficients of the interaction terms *OPList**Δ*lnSales* in both columns. More importantly, the results indicate that the increase in SG&A rigidity is greater than that in COGS rigidity (with the difference between the coefficients of *OPList**Δ*lnSales* in columns 1 and 2 being significant at the 1% level), which suggests that the change in overall cost rigidity after options listing mainly stems from substituting fixed SG&A for variable SG&A.

Firms adjust their cost structure by adjusting real activities. The extant literature has suggested that firms can outsource in-house activities to transfer some of the risk arising from financial constraints, demand uncertainty, and increased cost elasticity (e.g., Holzhacker, Krishnan, and Mahlendorf, 2015; Moon and Phillips, 2021; Choi, Ju, Trigeorgis, and Zhang, 2021). Therefore, if a firm commits more fixed costs to its operations after options listing, its demand for outsourcing may decrease. To test our prediction, we use *PC_raw* (the natural logarithm of 1 plus a firm's estimated payment amount of the purchase contracts) and *PC_COGS* (a firm's estimated payment amount of the purchase contracts normalized by the cost of goods sold) from the outsourcing data of Moon and Phillips (2021) to proxy for a firm's reliance on outsourcing contracts.[9] The results in columns 1 and 2 of Panel B show that after options listing, manufacturing firms' estimated raw payment amounts (*PC_raw*) and intensities (*PC_COGS*) of outsourcing activities significantly decrease ($p < 0.05$). Nevertheless, a flexible cost structure is reflected not only in the outsourcing decisions of manufacturing firms. In column 3, following Holzhacker, Krishnan, and Mahlendorf (2015), we examine the changes in firms' decisions to lease versus purchase assets after options listing. The results show that after options listing, firms prefer purchasing to leasing equipment ($p < 0.01$), with purchasing more equipment indicating more fixed resource commitments. Collectively, the results in Panel B provide further evidence for our baseline result from the perspective of changes in real business activities.

[Insert Table 4 here]

4.2.2 Alternative independent variables

In the baseline model, we primarily focus on the impact of the initial listing of options on the underlying firms' cost structure. However, if options trading indeed plays a significant informational role in firms' cost decisions, not only options listing but also options trading volume should significantly affect cost rigidity, as more active options trading could be more informative

[9] The test is performed on a small sample of manufacturing firms, because Moon and Phillips (2021) include only manufacturing firms in their study. We thank the authors for making the data available.

(Roll, Schwartz, and Subrahmanyam, 2010; Johnson and So, 2012; Hu, 2014). Therefore, in Panel C of Table 4, we replace the key independent variable *OPList* with common measures of options trading volume used in previous studies and interact them with Δ*lnSales* to examine the impact of options trading volume on cost structure.[10]

In column 1, we use the natural logarithm of 1 plus the annual dollar options trading volume (*lnVolume*) as the independent variable (Naiker, Navissi, and Truong, 2013; Blanco and Wehrheim, 2017; Chen, Ng, and Yang, 2021). The regression results show that *lnVolume**Δ*lnSales* is significantly negative ($p < 0.01$), indicating that an increase in options trading volume further increases cost rigidity. In column 2, we measure relative options trading volume (*OtS*) using the proportion of annual options trading volume to annual stock trading volume (Roll, Schwartz, and Subrahmanyam, 2010; Johnson and So, 2012; Hu, 2014), and the result is similar to that in column 1.

In column 3, we consider an event that provides an exogenous shock to options trading volume—the Penny Pilot Program. This program was initiated by the SEC in early 2007 and ran until 2020, with new firms being added annually as pilot firms. This program involved reducing tick sizes for selected options classes, thereby lowering trading costs and exogenously boosting trading volumes, liquidity, and informational efficiency (Cao, Goyal, Ke, and Zhan, 2024; Anagnostopoulou, Trigeorgis, and Tsekrekos, 2023). In this context, we construct a DiD estimator (*PPP*), which takes the value of 1 in the post-program years for the pilot firms and 0 otherwise.[11] Again, the regression results show that when options trading volume exogenously increases due to the Penny Pilot Program, firms' cost rigidity significantly rises ($p < 0.01$). Collectively, the results

[10] Samples in all regressions of Panel C, Table 4, include only firm-years with positive options trading volume, as previous studies on options trading volume generally argue that firms with and without positive options trading volume differ significantly (Blanco and Wehrheim, 2017; Chen, Ng, and Yang, 2021; Hsu, Ke, Ma, and Ruan, 2024). Nevertheless, even if we include firm-years with no options trading in the regressions, the results in Panel C remain highly consistent.

[11] Specifically, we obtain the list of firms added to the program each year from CBOE announcements (see https://www.cboe.com/us/options/market_statistics/historical_data/penny_class/ and https://www.cboe.com/us/options/notices/product_update/) and manually match them with our sample based on their ticker symbols. A total of 201 pilot firms are matched, which is consistent with Anagnostopoulou, Trigeorgis, and Tsekrekos (2023).

in Panel C indicate the significant impact of options trading on cost structure (whether it is the initial listing or the later trading volume), further demonstrating the robustness of our findings.

4.3 Further mitigating endogeneity

Prior research showed that options listing decisions are made by options exchanges and are largely exogenous to firm decisions (Blanco and Wehrheim, 2017; Bernile, Hu, Li, and Michaely, 2023; Brockman, Subasi, Wang, and Zhang, 2024; Hsu, Ke, Ma, and Ruan, 2024). Therefore, in the baseline model, we treat firms' options listing as quasi-natural experiments and construct the DiD estimator to get causal inference. To further mitigate endogeneity concerns, we conduct additional tests and report the results in Table 5.

4.3.1 Parallel trend assumption

For the DiD research design, an underlying assumption is a parallel trend between the treatment group and the control group; that is, there should be no significant difference in the cost structure dynamics of firms with and without options listing before the options listing events. We use an event study to test this key assumption. Specifically, we replace the indicator *OPList* in the baseline model with seven indicator variables—*pre_3*, *pre_2*, *pre_1*, *current*, *post_1*, *post_2*, and *post_3*—which represent three years before, two years before, and one year before the year of the options listing, and one year after, two years after, and three or more years after the initial listing of options. The regression results in column 1 show that the interaction terms between the pre-options listing indicators and Δ*lnSales* are all insignificant, indicating no significant difference in cost structure dynamics between the treatment group (options listing firms) and the control group (no options listing firms). Starting from the year of the initial listing of options, however, the cost rigidity of the treatment group samples begins to significantly increase compared with the control group (the regression coefficients of the post-options listing indicators*Δ*lnSales* are all significantly negative at the 10% level or higher), a change that is illustrated in Figure 1. Therefore, the validity of the parallel trend assumption further strengthens our confidence in the causal inference.

[Insert Table 5 here]

4.3.2 Placebo test

If our results were driven by random factors or confounding events, then randomly selecting some firms as the treatment group and randomly assigning their options listing time might also yield treatment effects similar to our baseline result. Based on this intuition, we conduct a placebo test. Specifically, we randomly select firms as options listing firms and randomly assign sample years as the years of the initial listing of their options to construct a simulated treatment indicator, *sim_OPList*. Then, we replace the actual treatment variable *OPList* in the baseline model with *sim_OPList* and obtain the placebo treatment effects. Next, we repeat the previous two steps 1000 times. Finally, we plot the empirical distribution of the 1000 placebo treatment effects in Figure 2. Clearly, the simulated treatment effects are concentrated around 0 (i.e., the placebo treatment effects are not significantly different from 0). Further, the real treatment effect, marked by the solid vertical line, differs greatly from the placebo treatment effects (the difference between the placebo treatment effects and the real treatment effect is significant at the 1% level). Therefore, the result of the placebo test further helps us rule out potential confounding effects from unobserved random factors or events.

4.3.3 Additional fixed effects

We incorporate additional fixed effects to further mitigate concerns about omitted variables. Although we have controlled for year and firm fixed effects in the baseline model, some industry characteristics, such as the industry life cycle, may be time-varying and cannot be captured by static firm fixed effects. Therefore, in column 2, we add *Year*Industry* high-dimensional fixed effects to control for more unobservable factors that vary with time and industries. The regression coefficient of *OPList**Δ*lnSales* in column 2 is significantly negative at the 1% level, indicating that our baseline result remains valid after controlling for more potentially omitted variables.

4.3.4 Propensity score matching (PSM) and entropy balancing (EB)

Firms with options listing could be significantly different from those with no options listing in various firm characteristics, which may also influence cost structure. Therefore, following Blanco and Wehrheim (2017) and Hsu, Ke, Ma, and Ruan (2024), we adopt two matching methods to control for observed differences in firms' cost structure between firms with and without options listing. In PSM, we use a logit regression with all the control variables in the baseline model to estimate the probability of a firm having options listing. Next, we match (with replacement) each firm with options listing to a firm without options listing and require the difference in the estimated probabilities to be no greater than 0.05.[12] Column 3 reports the regression result using PSM-matching samples. Unlike PSM, we do not drop any observations with the EB method but give each observation an entropy balancing weight to obtain a near-perfect covariate balance. We divide samples into treatment and control groups and then use EB with the first-order moment (i.e., mean) of all control variables in the baseline model to calculate the entropy balancing weights.[13] Column 4 reports the regression result of the weighted sample. The coefficients of *OPList**Δ*lnSales* remain significant in both columns 3 and 4 ($p < 0.01$), which suggests that our baseline result is robust to accounting for observable differences between firms with options listing and firms without options listing.[14]

4.4 Cross-sectional analysis

In the previous sections, we concluded that options listing increases the underlying firms' cost rigidity, a result that remains consistent across a series of robustness checks and after addressing endogeneity concerns. In this section, we further examine whether the impact of options listing on firms' cost structure aligns with our efficient options trading narrative that suggests that

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[12] The results remain largely consistent regardless of what matching estimators we choose—for example, matching with replacement versus without replacement, matching one firm with options listing to one versus two firms without options listing, and caliper = 0.005 versus 0.01 versus 0.05.

[13] Untabulated results indicate that EB significantly reduces the differences in the mean, variance, and skewness of the control variables between the two groups. We observe similar results if we use EB with the second-order moment (i.e., variance) or the third-order moment (i.e., skewness) of all control variables to calculate the weights.

[14] Although PSM and EB help to reduce observable differences between the treatment group and the control group, matching based on covariates cannot eliminate differences in the other unobservable characteristics. A more direct test is to regress with only observations of firms with options listings, as these samples are likely to be more similar in both observable and unobservable characteristics. In this case, there is no firm whose options are never listed during the sample period (i.e., no "never-treated" group). Instead, the later-listed firms serve as the control group for the earlier-listed firms. Untabulated results show that our baseline result is highly robust to this subsample.

options trading mitigates financial constraints and managers' bad-news hoarding incentive by enhancing the value-relevant information flows.

4.4.1 Financial constraints

First, higher information asymmetry is often associated with a higher cost of capital and more limited access to external financing (Armstrong, Core, Taylor, and Verrecchia, 2011; Lambert, Leuz, and Verrecchia, 2012). Therefore, if options listing alleviates financial constraints and thus motivates firms to undertake more fixed investments, we should observe that the impact of options listing on cost rigidity is more pronounced in firms with a higher level of information asymmetry, in which the marginal effect of incremental information flows from options listing in reducing information asymmetry and thereby lowering financial constraints is stronger. We use two variables to measure the degree of information asymmetry between firms and the market, including the (absolute value of) analyst forecast error and bid-ask spread in stock prices (Zhang, 2006; Cho, Lee, and Pfeiffer, 2013). The variables of interest are the three-way interaction terms, *error*OPList*ΔlnSale* and *spread*OPList*ΔlnSale.* The regression results in Panel A, Table 6, demonstrate that our baseline result is stronger in firms with a higher degree of information asymmetry (i.e., firms with a higher analyst forecast error and a larger bid-ask spread in stock prices) (p < 0.01 in both tests). These results support the notion that options listing enhances information flows in firms with higher information asymmetry and alleviates firms' financial constraints, subsequently increasing cost rigidity.

We then conduct additional cross-sectional tests by employing variables that are widely used in the literature to measure firms' financial constraints. The proxies we use include the *WW index* (Whited and Wu, 2006) and the *SA index* (Hadlock and Pierce, 2010). The regression results in columns 1 and 2 of Panel B, Table 6, show that our baseline result is more pronounced in firms with higher financial constraints (i.e., a higher *WW index* and a higher *SA index*) (p < 0.05 in both columns 1 and 2). In column 3, we measure firms' financial constraints according to whether a firm has a long-term credit rating. Following previous studies, we classify firms as having no credit rating, *Norating*, if their long-term debt is not rated by Standard & Poor's or their debt is in default (Almeida, Campello, and Weisbach, 2004; Denis and Sibilkov, 2010). In column 4, we measure firms' financial status according to the dividend payout ratio, *Payout*. The findings in columns 3 and 4 suggest that the effect of options listing on cost rigidity is stronger in firms without a credit rating and with a lower dividend payout ratio—that is, firms that are financially constrained (p < 0.01 in both columns 3 and 4). Hence, the results in Table 6 provide consistent evidence that reducing financial constraints is a possible mechanism through which options listing increases cost rigidity.

[Insert Table 6 here]

4.4.2 Managers' bad-news hoarding incentive

In this section, we test whether options listing leads to a higher cost rigidity by mitigating managerial career concern and thus their incentives to hoard bad news. We set *Loss_Avoidance* to one when a firm-year's income before extraordinary items (scaled by beginning assets) falls into a small positive interval (i.e., [0, 0.01]) and 0 otherwise (Kama and Weiss, 2013).[15] Kama and Weiss (2013) argue that managers with career concerns would avoid market punishment by barely avoiding losses. We therefore use a small positive return-on-assets ranging between 0 and 0.01 as a proxy for managers' incentive to avoid reporting losses. We then test the impact of options listing on cost rigidity conditional on managers' loss avoidance incentive, captured by the coefficient of *Loss_Avoidance*OPList*ΔlnSale*. Column 1 in Table 7 shows a significantly negative coefficient of *Loss_Avoidance*OPList*ΔlnSales* (p < 0.01), indicating that options listing increases firms' cost rigidity more in firms in which managers have a stronger incentive to avoid loss.

We perform similar tests in columns 2 and 3 of Table 7. If managers choose an elastic cost structure to achieve a better and more stable performance result, they are more likely to do so in firms with higher short-sale constraints in which subsequent stock price crash risk is higher.

[15] The results are similar regardless of whether we define the upper limit of the small positive interval as 0.01–0.05.

However, when options trading allows investors to profit from negative information by circumventing the short-sale constraints, managers are more likely to switch to a rigid cost structure. In other words, when options listing increases cost rigidity by promoting information flow, such an effect should be stronger in firms with *ex ante* higher short-sale constraints, in which managers have a stronger incentive to report stable and favorable performance.

We first measure short-sale constraints using *Institutional_Ownership* (the ratio of total shares owned by institutional investors to total shares outstanding), because lower institutional ownership indicates a lower supply of lendable stock and therefore higher short-sale constraints (Hu, 2014). Furthermore, short-sale constraints may be exacerbated if short-sale demand increases, given the limited supply. Therefore, our second measure of short-sale constraints, *Short-sale constraints*, is the gap between short interest (the annual average ratio of shares sold short for a given firm in a given month to the monthly shares outstanding) and institutional ownership, with a larger gap indicating that lendable stock is more in short supply and that short-sale constraints are more severe (Chen, Hong, and Stein, 2002; Asquith, Pathak, and Ritter, 2005).[16] The results are presented in columns 2 and 3 of Table 7. We find that options listing significantly increases cost rigidity in firms subject to higher short-sales constraints, as reflected by the significantly positive coefficient of *Institutional_Ownership*OPList**Δ*lnSales* (column 2) ($p < 0.05$) and the significantly negative coefficient of *Short-sale_Constraints*OPList**Δ*lnSales* (column 3) ($p < 0.05$).

[Insert Table 7 here]

5. Additional Analyses Related to the Cost Structure Literature

In this section, we perform additional tests to examine the impact of options listing on firms' cost structure choices. Banker, Byzalov, and Plehn-Dujowich (2014) propose demand shock as an important determinant of cost rigidity. In particular, they identify a "congestion cost" when firms

[16] We do not use short interest alone as an alternative measure because higher short interest without considering the supply of lendable stock does not reliably indicate higher short-sale constraints (Chen, Hong, and Stein, 2002; Autore, Boulton, and Braga-Alves, 2015).

cannot quickly expand production capacity.[17] They argue that firms with greater demand uncertainty should adopt a rigid cost structure to reduce the congestion cost, because an unusually positive demand shock is more likely to be realized when demand uncertainty is larger. While we explained the effect of options listing on cost structure in terms of enhanced information flow in the previous analysis, we examine in this section how options listing affects cost structure when considering firms' demand uncertainty. Our earlier analyses document that options trading improves firms' information environment, reduces firms' financing constraints, and encourages firms to commit more internal resources and increase cost rigidity. Following Banker, Byzalov, and Plehn-Dujowich (2014), we expect such an effect to be stronger in firms with higher demand uncertainty to the extent that firms would reduce congestion cost and capitalize on the positive demand shock. In column 1 of Table 8, we measure demand uncertainty with sales uncertainty, *Sales_Uncertainty*, which is the standard deviation of log-changes in sales for a firm in the sample period (Banker, Byzalov, and Plehn-Dujowich, 2014).

To substantiate the findings with the sales uncertainty measure, we redo the test in column 2 of Table 8 with another measure of demand uncertainty, *Similarity*. This is the product market similarity measure constructed by Hoberg and Phillips (2016).[18] Chang, Hall, and Paz (2021) argue that a competitive supplier market will shift the bargaining power to strong customers. To secure their relationship with the customers, the suppliers in the industry will make more relationship-specific investments, which are largely fixed. In sum, the results from Table 8 indicate that firms facing more demand shocks (i.e., when sales are more variable and/or when products are more replaceable in the product market) undertake more fixed investments, consistent with Banker,

[17] Nevertheless, Kallapur and Eldenburg (2005) have a counter argument here. Based on a real-options theory of investment, Kallapur and Eldenburg (2005) infer that uncertainty leads firms to prefer production with low fixed and high variable costs. Managers should have greater flexibility to respond to changes in business conditions when upfront (fixed) costs are lower. Because the value of such flexibility increases with uncertainty, managers will prefer production technologies with high variable and low fixed costs when uncertainty increases. Or put in a simple way, real-options theory implies that the ratio of variable to fixed costs in the cost structure should be higher in firms with higher uncertainty.

[18] *Similarity* is a recently developed, text-based product market competition proxy, derived from firms' 10-K product descriptions (Hoberg and Phillips, 2016). This is a firm-level, time-varying measure capturing changes in competitive landscapes more quickly than traditional industry-based measures.

Byzalov, and Plehn-Dujowich (2014). More importantly, we find that options listing strengthens this result, as evidenced by significantly negative coefficients of *Sales_Uncertainty*OPList*ΔlnSales* and *Similarity*OPList*ΔlnSales* (p < 0.01 and p < 0.05, respectively). It is possible that options listing makes external financing more accessible and therefore renders firms' fixed investments more likely.

[Insert Table 8 here]

6 Conclusion

Understanding how managers adjust their firms' cost structures is a fundamental issue in cost accounting (Garrison, Noreen, and Brewer., 2012; Horngren, Datar, and Rajan, 2012). We use options listing as a test setting and find that firms increase the proportion of fixed costs relative to variable costs when they have options listed on the stock exchange, and this holds especially for the SG&A cost component. The results are robust to modifications in the fixed effects included in the tests, the sample constructions, and the empirical methods used to control for the endogeneity concern in the association between cost structure and options listing. Cross-sectional tests suggest that the effect of options listing on firms' cost structure is stronger in firms with higher financial constraints and/or higher career concerns (i.e., the propensity to report stable and favorable operational results). The findings indicate that options listing encourages firms to replace variable investments with fixed ones by relaxing financial constraints and/or reducing the likelihood of managers being penalized by the market because of temporary shortfalls in operational performance.

Highlighting the implications that options listing can have for firms' operations, our results can be important to investors. The improvement in transparency and information flow following options listing partly induces an increase in firms' operational leverage. Consequently, firms may be more inclined to engage in risk-taking behaviors, pursuing strategies that could lead to higher rewards but also greater potential downsides. For investors, recognizing these dynamics is crucial, as it can influence the assessment of a firm's financial health and strategic direction, so investors should navigate their investments accordingly.

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Appendix A: Definitions of Variables

Variables	Definition	Data source
Dependent variables		
Δ*lnOPR*	Change in the natural logarithm of operating costs in year *t* relative to year *t*-1. Operating costs equal the sum of the cost of goods sold and SG&A costs.	CRSP/Compustat Merged
Δ*lnCOGS*	Change in the natural logarithm of the cost of goods sold in year *t* relative to year *t*-1.	CRSP/Compustat Merged
Δ*lnSG&A*	Change in the natural logarithm of SG&A costs in year *t* relative to year *t*-1.	CRSP/Compustat Merged
PC_raw	Natural logarithm of 1 plus a firm's estimated payment amount within the closest fiscal year under all of the purchase contracts.	Moon and Phillips (2021)
PC_COGS	A firm's estimated payment amount within the closest fiscal year under all of the purchase contracts normalized by the cost of goods sold.	Moon and Phillips (2021)
LeasevsBuy	Lease and rental costs divided by the sum of depreciation/amortization and lease and rental costs.	CRSP/Compustat Merged
Independent variables		
Δ*lnSales*	Change in the natural logarithm of sales revenue in year *t* relative to year *t*-1.	CRSP/Compustat Merged
OPList	An indicator variable that takes the value of 1 in the post-options-listing years for firms with listed options and 0 otherwise. We define a firm's options listing date as the first date on which it appears in the OptionMetrics database.	OptionMetrics
lnVolume	Natural logarithm of 1 plus the annual dollar options trading volume. We multiply the daily trading volume with the midpoint of the end-of-day bid and ask prices for each options contract on a stock. Then, we aggregate all listed options on a stock across all trading days annually to get the annual dollar options trading volume.	OptionMetrics
OtS	Annual options trading volume to annual stock trading volume.	OptionMetrics CRSP
PPP	An indicator variable that takes the value of 1 in the post-PPP (Penny Pilot Program) years for the pilot firms and 0 otherwise.	CBOE announcements
error	Analyst forecast error measured by the absolute value of the difference between a firm's actual EPS and its consensus (median) forecast EPS, normalized by the stock closing price.	IBES
spread	Annual average of the daily high-low bid-ask spread of stocks normalized by the stock closing price.	CRSP
SA	The SA index is calculated as $SA = -0.737*Size + 0.043*Size^2 - 0.040*Age$, where *Size* equals the log of inflation-adjusted book assets and *Age* is the number of years the firm is listed with a non-missing stock price on Compustat. In calculating this index, *Size* is winsorized (i.e., capped) at (the	CRSP/Compustat Merged

	log of) $4.5 billion, and *Age* is winsorized at 37 years, following Hadlock and Pierce (2010).	
WW	The *WW* index is calculated as *WW = -0.091*CF - 0.062*DIVPOS + 0.021*TLTD - 0.044*LNTA + 0.102*ISG - 0.035*SG*, where *CF* is the ratio of cash flows to total assets, *DIVPOS* is an indicator that takes the value of 1 if the firm pays cash dividends, *TLTD* is the ratio of long-term debt to total assets, *LNTA* is the natural logarithm of total assets, *ISG* denotes the firm's three-digit industry sales growth, and *SG* represents the firm's sales growth.	CRSP/Compustat Merged
Norating	Firms are classified as *Norating* if they have not had their long-term debt rated by Standard & Poor's (the initial credit rating date is not available in the S&P credit ratings database or beyond our sample period) or their debt is in default (rating of D or SD).	S&P Credit Ratings
Payout	Dividend payout ratio calculated as the total amount of dividends declared on the common/ordinary capital of the company divided by net income.	CRSP/Compustat Merged
Sales_Uncertainty	Standard deviation of log-changes in sales for all valid observations of a firm in the sample period.	CRSP/Compustat Merged
Similarity	Product similarity measure developed by Hoberg and Phillips (2016) using textual analysis of firms' product descriptions in their 10-K filings.	Hoberg and Phillips (2016)
Loss_Avoidance	An indicator variable that takes the value of 1 when a firm's annual income before extraordinary items to beginning asset is in the interval [0, 0.01] and 0 otherwise.	CRSP/Compustat Merged
Institutional_Ownership	Ratio of total shares owned by institutional investors to total shares outstanding.	Thomson Reuters Institutional (13f) Holdings
Short_Interest	Annual average of shares sold short for a given firm in a given month divided by the monthly shares outstanding.	Compustat
Short-sale_Constraints	Calculated as *Short_Interest - Institutional_Ownership*.	Thomson Reuters Institutional (13f) Holdings; Compustat
Control variables		
Size	Natural logarithm of a firm's total assets.	CRSP/Compustat Merged
Age	Natural logarithm of the months between the current month and the first month that a stock appears in CRSP.	CRSP/Compustat Merged
MtB	Firm's market value (shares outstanding*closing price) to book value (stockholders' equity + deferred taxes + investment tax credit - preferred stock).	CRSP/Compustat Merged
AI	Asset intensity, calculated as the ratio of total assets to sales revenue.	CRSP/Compustat Merged
EI	Employee intensity, calculated as the total number of employees scaled by sales revenue.	CRSP/Compustat Merged

FCF	Free cash flows, calculated as cash flows from operating activities minus common and preferred dividends scaled by the total assets.	CRSP/Compustat Merged

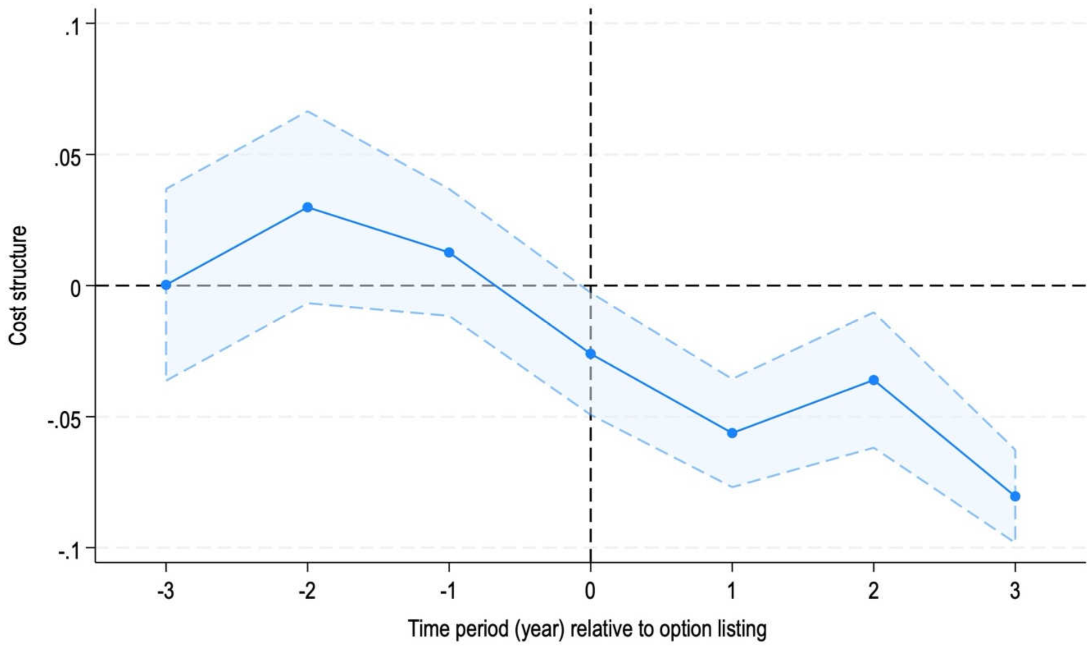

Figure 1: Parallel trend test

This figure depicts the results of the parallel trend test. The horizontal axis represents the interval of sample years relative to the firms' options listing years, while the vertical axis represents the dynamics of the firms' cost structure. We treat period -4 and earlier as the baseline group, and the period after 3 is winsorized to period 3. The dashed area represents the 90% confidence interval.

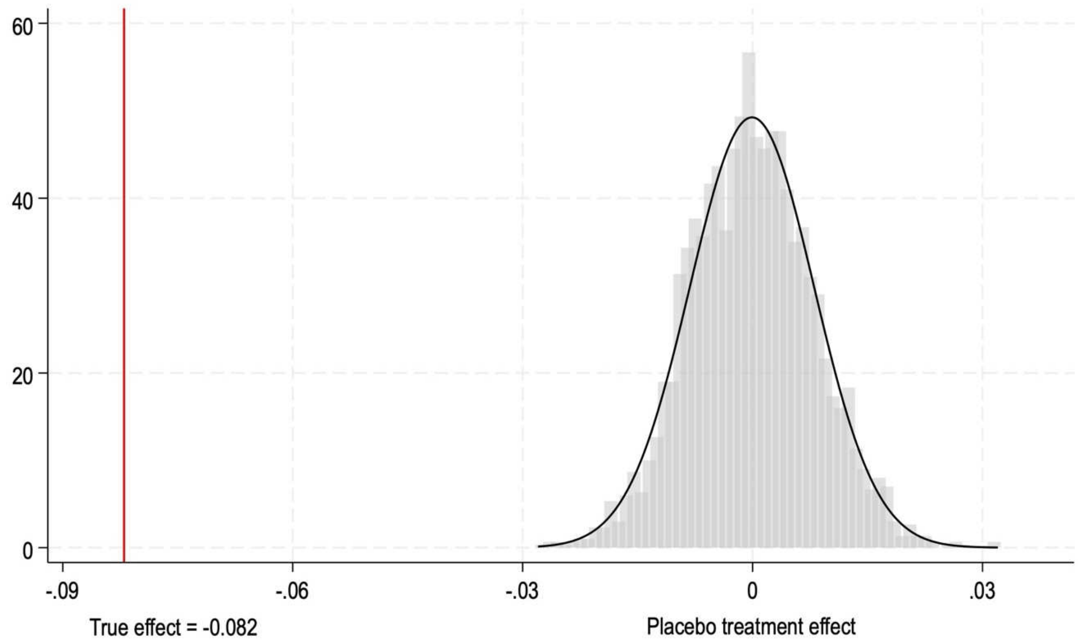

Figure 2: Placebo test

This figure displays the empirical distribution of the 1,000 placebo treatment effects obtained from the placebo test as detailed in Section 4.3.2 and the real treatment effect. The empirical distribution of the placebo treatment effect is displayed in the histogram overlaid with a normal curve. The solid vertical line is added to mark the real treatment effect derived from the baseline regression result in Table 3.

Table 1: Sample distribution

This table presents the sample distribution. Panels A and B detail the distribution of the sample by fiscal year and Fama–French 48 industries, respectively.

Panel A: Sample distribution by fiscal year

Fiscal year	Frequency	Percent	Fiscal year	Frequency	Percent
1996	2,035	4.68	2010	1,482	3.41
1997	2,401	5.53	2011	1,399	3.22
1998	2,435	5.60	2012	1,290	2.97
1999	2,380	5.48	2013	1,267	2.92
2000	2,109	4.85	2014	1,238	2.85
2001	2,113	4.86	2015	1,151	2.65
2002	2,073	4.77	2016	1,088	2.5
2003	2,137	4.92	2017	1,053	2.42
2004	2,194	5.05	2018	973	2.24
2005	2,160	4.97	2019	898	2.07
2006	2,115	4.87	2020	843	1.94
2007	1,985	4.57	2021	834	1.92
2008	1,632	3.76	2022	696	1.6
2009	1,464	3.37			
Total				43,445	100

Panel B: Sample distribution by Fama–French 48 industries

Industry code	Industry name	Frequency	Percent
1	Agriculture	212	0.49
2	Food Products	984	2.26
3	Candy and Soda	130	0.3
4	Beer and Liquor	226	0.52
5	Tobacco Products	57	0.13
6	Recreation	437	1.01
7	Entertainment	713	1.64
8	Printing and Publishing	316	0.73
9	Consumer Goods	731	1.68
10	Apparel	835	1.92
11	Healthcare	1,055	2.43
12	Medical Equipment	1,776	4.09
13	Pharmaceutical Products	1,632	3.76
14	Chemicals	869	2
15	Rubber and Plastic Products	454	1.04
16	Textiles	203	0.47
17	Construction Materials	1,099	2.53
18	Construction	743	1.71
19	Steelworks, Etc.	742	1.71
20	Fabricated Products	188	0.43
21	Machinery	1,886	4.34
22	Electrical Equipment	893	2.06
23	Automobiles and Trucks	797	1.83
24	Aircraft	335	0.77
25	Shipbuilding, Railroad Equipment	145	0.33
26	Defense	78	0.18
27	Precious Metals	238	0.55
28	Non-metallic and Industrial Metal Mining	271	0.62
29	Coal	142	0.33
30	Petroleum and Natural Gas	1,254	2.89
32	Communication	1,656	3.81
33	Personal Services	711	1.64
34	Business Services	7,089	16.32
35	Computers	1,811	4.17
36	Electronic Equipment	3,120	7.18
37	Measuring and Control Equipment	1,073	2.47
38	Business Supplies	595	1.37

39	Shipping Containers	146	0.34
40	Transportation	1,613	3.71
41	Wholesale	2,016	4.64
42	Retail	2,501	5.76
43	Restaurants, Hotels, Motels	1,102	2.54
48	Almost Nothing	571	1.31
Total		43,445	100

Table 2: Summary statistics

This table presents the descriptive statistics of the key variables in the baseline model. Δ*lnOPR* is the change in the natural logarithm of operating costs in year *t* relative to year *t*-1. Operating costs equal the sum of the cost of goods sold and SG&A costs. Δ*lnSales* is the change in the natural logarithm of sales revenue in year *t* relative to year *t*-1. *OPList* is an indicator variable that takes the value of 1 in the post-options-listing years for firms with listed options and 0 otherwise. *Size* is the natural logarithm of a firm's total assets. *Age* is the natural logarithm of the months between the current month and the first month that a stock appears in CRSP. *MtB* is a firm's market value divided by the book value. *AI* is the asset intensity, calculated as the ratio of total assets to sales revenue. *EI* is the employee intensity, calculated as the total number of employees scaled by sales revenue. *FCF* is free cash flows, calculated as the cash flows from operating activities minus the common and preferred dividends scaled by the total assets. Detailed definitions of the variables can be found in Appendix A.

	Obs	Mean	SD	P25	Median	P75
Δ*lnOPR*	43,445	0.076	0.201	-0.025	0.060	0.167
Δ*lnSales*	43,445	0.072	0.211	-0.032	0.057	0.174
OPList	43,445	0.450	0.497	0.000	0.000	1.000
Size	43,445	5.481	1.841	4.165	5.350	6.620
Age	43,445	5.174	0.700	4.682	5.124	5.704
MtB	43,445	2.909	3.457	1.119	1.850	3.250
AI	43,445	1.325	1.968	0.675	1.004	1.532
EI	43,445	0.008	0.011	0.004	0.006	0.009
FCF	43,445	0.063	0.110	0.019	0.070	0.117

Table 3: Baseline regression results

This table presents the baseline results. Δ*lnOPR* is the change in the natural logarithm of operating costs in year *t* relative to year *t*-1. Operating costs equal the sum of COGS and SG&A costs. Δ*lnSales* is the change in the natural logarithm of sales revenue in year *t* relative to year *t*-1. *OPList* is an indicator variable that takes the value of 1 in the post-options-listing years for firms with listed options and 0 otherwise. Detailed variable definitions are provided in Appendix A. Standard errors are clustered on the firm-year level. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. *t*-statistics are reported in parentheses.

Variables	(1)	(2)
	Δ*lnOPR*	Δ*lnOPR*
Δ*lnSales*	0.803***	0.689***
	(154.71)	(22.18)
OPList	0.011***	0.006***
	(5.15)	(2.79)
OPList*ΔlnSales	-0.021**	-0.082***
	(-2.53)	(-8.56)
Size		0.028***
		(14.72)
*Size**Δ*lnSales*		0.036***
		(12.97)
Age		-0.035***
		(-10.29)
*Age**Δ*lnSales*		-0.009
		(-1.52)
MtB		-0.001**
		(-2.22)
*MtB**Δ*lnSales*		-0.003**
		(-2.52)
AI		-0.003
		(-1.07)
*AI**Δ*lnSales*		-0.007***
		(-2.60)
EI		0.816***
		(5.00)
*EI**Δ*lnSales*		0.752**
		(2.09)
FCF		-0.144***
		(-15.23)
*FCF**Δ*lnSales*		0.261***
		(6.22)
Constant	0.014***	0.054***
	(12.78)	(2.88)
Firm FEs	YES	YES
Year FEs	YES	YES
N	50,534	43,445
Adj. R^2	0.75	0.78

Table 4: Robustness checks

This table presents the results of robustness checks. Detailed variable definitions are provided in Appendix A. *Controls* include all the control variables and their interactions with Δ*lnSales* in the baseline model. Standard errors are clustered at the firm-year level. We obtain the *p*-value for the difference in *OPList**Δ*lnSales* between groups using a Z-test. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. *t*-statistics are reported in parentheses.

Panel A: Different cost components as alternative dependent variables

Variables	(1)	(2)
	Δ*lnCOGS*	Δ*lnSG&A*
Δ*lnSales*	1.038***	0.470***
	(24.25)	(8.87)
OPList	-0.000	0.008**
	(-0.11)	(2.14)
OPList*ΔlnSales	-0.026*	-0.091***
	(-1.77)	(-6.40)
Controls	YES	YES
Firm FEs	YES	YES
Year FEs	YES	YES
p-value of coefficient difference	0.001***	
N	40,899	40,900
Adj. R^2	0.68	0.42

Panel B: Outsourcing as alternative dependent variable

Variables	(1)	(2)	(3)
	PC_raw	*PC_COGS*	*LeasevsBuy*
OPList	-0.097**	-0.010**	-0.006***
	(-2.05)	(-2.04)	(-2.74)
Size	0.587***	0.015***	-0.043***
	(13.72)	(2.96)	(-28.67)
Age	0.544***	0.038***	0.001
	(4.27)	(2.88)	(0.24)
MtB	0.013**	0.002**	0.001***
	(2.28)	(2.54)	(4.15)
AI	-0.195***	0.010	0.000
	(-3.82)	(1.31)	(0.36)
EI	-0.940	1.724	-0.043
	(-0.12)	(1.62)	(-0.23)
FCF	-0.096	-0.027	-0.003
	(-0.73)	(-1.43)	(-0.39)
Constant	-4.280***	-0.217***	0.575***
	(-5.98)	(-2.84)	(27.69)
Firm FEs	YES	YES	YES
Year FEs	YES	YES	YES
N	7,223	7,325	39,400
Adj. R^2	0.85	0.68	0.84

Panel C: Options trading volume as alternative independent variable

Variables	(1)	(2)	(3)
	Δ*lnOPR*	Δ*lnOPR*	Δ*lnOPR*
Δ*lnSales*	0.477***	0.476***	0.675***
	(8.74)	(8.67)	(11.10)
lnVolume	0.012***		
	(9.78)		
lnVolume*ΔlnSales	-0.027***		
	(-6.19)		
OtS		0.088***	
		(5.33)	
OtS*ΔlnSales		-0.227***	
		(-3.77)	
PPP			-0.003
			(-0.51)
PPP*ΔlnSales			-0.148***

			(-3.94)
Controls	YES	YES	YES
Firm FEs	YES	YES	YES
Year FEs	YES	YES	YES
N	18,028	18,028	26,135
Adj. R^2	0.77	0.77	0.71

Table 5: Further strengthening causal inference

This table presents the results of further tests to strengthen causal inference. The dependent variable in all columns is Δ*lnOPR*. *pre_3*, *pre_2*, *pre_1*, *current*, *post_1*, *post_2*, and *post_3* represent three years before, two years before, and one year before the year of options listing, and one, two, and three or more years after options listing, respectively. Other variable definitions are provided in Appendix A. *Controls* include all the control variables and their interactions with Δ*lnSales* in the baseline model. Industry fixed effects are based on the Fama-French 48 classification. Standard errors are clustered at the firm-year level. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. *t*-statistics are reported in parentheses.

Variables	(1)	(2)	(3)	(4)
	Parallel trend	High-dimensional FEs	PSM-DiD	Entropy balancing
Δ*lnSales*	0.668***	0.672***	0.633***	0.707***
	(20.82)	(21.60)	(12.07)	(19.95)
OPList		0.007***	0.007*	0.007***
		(3.17)	(1.82)	(2.82)
OPList*ΔlnSales		-0.080***	-0.068***	-0.073***
		(-8.45)	(-4.50)	(-6.98)
pre_3	0.001			
	(0.13)			
pre_2	-0.001			
	(-0.25)			
pre_1	-0.013***			
	(-3.07)			
current	0.003			
	(0.65)			
post_1	0.013***			
	(3.65)			
post_2	-0.004			
	(-1.13)			
post_3	-0.010***			
	(-3.15)			
pre_3*ΔlnSales	0.000			
	(0.01)			
pre_2*ΔlnSales	0.030			
	(1.34)			
pre_1*ΔlnSales	0.013			
	(0.86)			
current*ΔlnSales	-0.026*			
	(-1.83)			
post_1*ΔlnSales	-0.056***			
	(-4.48)			
post_2*ΔlnSales	-0.036**			
	(-2.29)			
post_3*ΔlnSales	-0.080***			
	(-7.45)			
Controls	YES	YES	YES	YES
Firm FEs	YES	YES	YES	YES
Year FEs	YES	NO	YES	YES
Year*Industry FEs	NO	YES	NO	NO
N	43,445	43,424	15,653	43,445
Adj. R^2	0.78	0.79	0.78	0.77

Table 6: Cross-sectional tests on financial constraints

This table presents the results of the cross-sectional tests on financial constraints. Variable definitions are provided in Appendix A. *Controls* include all the control variables and their interactions with Δ*lnSales* in the baseline model. Standard errors are clustered at the firm-year level. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. *t*-statistics are reported in parentheses.

Panel A: Information asymmetry measures

Variables	(1)	(2)
	Δ*lnOPR*	Δ*lnOPR*
Δ*lnSales*	0.544***	0.677***
	(12.90)	(20.27)
OPList	0.013***	0.009***
	(4.99)	(3.94)
*OPList**Δ*lnSales*	-0.086***	-0.065***
	(-7.29)	(-6.15)
error	0.012	
	(0.58)	
*error*OPList*	0.059*	
	(1.90)	
*error**Δ*lnSales*	0.042	
	(0.46)	
error*OPList*ΔlnSales	-0.457***	
	(-2.83)	
spread		-0.014***
		(-9.04)
*spread*OPList*		-0.023***
		(-3.68)
*spread**Δ*lnSales*		-0.001
		(-0.10)
spread*OPList*ΔlnSales		-0.086***
		(-3.28)
Controls	YES	YES
Firm FEs	YES	YES
Year FEs	YES	YES
N	24,772	43,274
Adj. R^2	0.80	0.78

Panel B: Financial constraint measures

Variables	(1)	(2)	(3)	(4)
	Δ*lnOPR*	Δ*lnOPR*	Δ*lnOPR*	Δ*lnOPR*
Δ*lnSales*	0.760***	0.763***	0.677***	0.705***
	(20.99)	(23.71)	(17.34)	(21.90)
OPList	0.064***	0.018***	-0.002	0.008***
	(4.98)	(3.91)	(-0.40)	(3.72)
*OPList**Δ*lnSales*	-0.209***	-0.116***	-0.040*	-0.082***
	(-4.09)	(-5.40)	(-1.92)	(-7.85)
SA	-0.039***			
	(-4.84)			
*SA*OPList*	-0.070***			
	(-4.29)			
*SA**Δ*lnSales*	0.017***			
	(4.72)			
SA*OPList*ΔlnSales	-0.038**			
	(-2.45)			
WW		0.014		
		(1.28)		
*WW*OPList*		-0.399***		
		(-6.30)		
*WW**Δ*lnSales*		0.048***		
		(3.44)		
WW*OPList*ΔlnSales		-0.167**		

		(-2.13)		
Norating			0.004	
			(0.76)	
*Norating*OPList*			0.011**	
			(2.55)	
*Norating*ΔlnSales*			0.020	
			(1.04)	
Norating*OPList*ΔlnSales			-0.051**	
			(-2.20)	
Payout				-0.002
				(-1.48)
*Payout*OPList*				-0.001
				(-0.88)
*Payout*ΔlnSales*				-0.002
				(-0.34)
Payout*OPList*ΔlnSales				0.025**
				(2.43)
Controls	YES	YES	YES	YES
Firm FEs	YES	YES	YES	YES
Year FEs	YES	YES	YES	YES
N	42,668	43,282	43,445	32,153
Adj. R^2	0.78	0.78	0.78	0.86

Table 7: Cross-sectional tests on managers' bad-news hoarding incentive

This table presents the results of the cross-sectional tests on managers' incentive to avoid reporting losses. Variable definitions are provided in Appendix A. *Controls* include all the control variables and their interactions with Δ*lnSales* in the baseline model. Standard errors are clustered at the firm-year level. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. *t*-statistics are reported in parentheses.

Variables	(1)	(2)	(3)
	Δ*lnOPR*	Δ*lnOPR*	Δ*lnOPR*
Δ*lnSales*	0.666***	0.629***	0.552***
	(21.02)	(16.35)	(10.00)
OPList	0.004**	0.016***	0.011**
	(2.16)	(3.21)	(2.08)
*OPList**Δ*lnSales*	-0.050***	-0.139***	-0.136***
	(-5.67)	(-6.17)	(-5.19)
Loss_Avoidance	-0.006***		
	(-2.70)		
*Loss_Avoidance*OPList*	0.005		
	(1.31)		
*Loss_Avoidance**Δ*lnSales*	0.095***		
	(5.49)		
Loss_Aviodance*OPList*ΔlnSales	-0.080***		
	(-2.83)		
Institutional_Ownership		0.022***	
		(3.12)	
*Institutional_Ownership*OPList*		-0.017**	
		(-2.18)	
*Institutional_Ownership**Δ*lnSales*		-0.048*	
		(-1.86)	
Institutional_Ownership*OPList*ΔlnSales		0.093**	
		(2.45)	
Short-sale_Constraints			-0.009
			(-1.11)
*Short-sale_Constraints*OPList*			0.009
			(1.09)
*Short-sale_Constraints**Δ*lnSales*			0.012
			(0.37)
Short-sale_Constraints*OPList*ΔlnSales			-0.098**
			(-2.11)
Controls	YES	YES	YES
Firm FEs	YES	YES	YES
Year FEs	YES	YES	YES
N	37,542	31,078	23,566
Adj. R^2	0.83	0.80	0.79

Table 8: Cross-sectional tests on product market factors

This table presents the results of the cross-sectional tests on product market factors. Variable definitions are provided in Appendix A. *Controls* include all the control variables and their interactions with Δ*lnSales* in the baseline model. Standard errors are clustered at the firm-year level. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. *t*-statistics are reported in parentheses.

Variables	(1)	(2)
	Δ*lnOPR*	Δ*lnOPR*
Δ*lnSales*	0.918***	0.754***
	(23.37)	(20.52)
OPList	0.006	0.004
	(1.30)	(1.31)
*OPList*Δ*lnSales*	-0.004	-0.049***
	(-0.17)	(-4.19)
*Sales_Uncertainty*OPList*	-0.017	
	(-0.52)	
*Sales_Uncertainty*Δ*lnSales*	-0.395***	
	(-7.18)	
Sales_Uncertainty*OPList*ΔlnSales	-0.241***	
	(-2.76)	
Similarity		0.001
		(1.13)
*Similarity*OPList*		0.001
		(1.18)
*Similarity*Δ*lnSales*		-0.011***
		(-4.34)
Similarity*OPList*ΔlnSales		-0.008**
		(-2.33)
Controls	YES	YES
Firm FEs	YES	YES
Year FEs	YES	YES
N	43,445	37,747
Adj. R^2	0.78	0.79