Double-Edged Sword: The Case of Whistleblower Laws

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Keywords: Whistleblower laws, Anti-takeover provisions, False Claims Acts, Corporate governance, Managerial entrenchment.

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ABSTRACT

Using states' staggered passage of the False Claims Act (FCA) as an exogenous increase in whistleblower risk, we find that firms are more likely to adopt antitakeover provisions (ATPs) following FCA legislation. The effect of the FCA is more pronounced for firms with a higher awareness of and propensity for whistleblowing activities and where managers derive substantial private benefits. This ATP adoption negatively impacts firm valuation and performance, suggesting that whistleblower laws inadvertently reinforce managerial entrenchment and exacerbate agency problems. This study sheds light on the unintended consequences of whistleblower laws on corporate governance.

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1. Introduction

After several high-profile financial frauds, such as Enron, HealthSouth, and WorldCom, regulators have strengthened whistleblower provisions. For example, the Sarbanes-Oxley Act (enacted in 2002) mandates increased whistleblower protections and compliance monitoring, and the Dodd-Frank whistleblower program (implemented by the US Securities and Exchange Commission (SEC) in 2011) provides financial rewards to whistleblowers who report financial fraud directly to the SEC. Although regulators argue and researchers have shown that these whistleblower laws can deter accounting fraud (e.g., Dyck, Morse, and Zingales, 2010; Wilde, 2017; Dey, Heese, and Perez-Cavazos, 2021; Berger and Lee, 2022), there is little evidence on how managers respond to their exposure to the threat from whistleblowing.

Do managers suffer personal consequences for engaging in fraud? Karpoff, Lee, and Martin (2008a) show that of the 2,206 individuals identified by regulators as culpable parties, 93% lost their jobs by the end of the regulatory enforcement period. The likelihood of removal is positively related to the size of the misconduct's harm to shareholders and the quality of the firm's governance. Furthermore, fraud detection relies more on whistleblowers (employees, media, and industry regulators) than on the standard players in corporate governance (investors, SEC, and auditors) (Dyck, Morse, and Zingales, 2010). Therefore, managers tend to select governance structures that entrench their position in response to their exposure to the threat of whistleblowing.

From a theoretical perspective, corporate managers might prefer a weak governance structure, such as adding antitakeover provisions (ATPs) to a corporate charter, even if this action hurts their shareholders (e.g., Bebchuk, Coates, and Subramanian, 2002; Gompers, Ishii, and Metrick, 2003; Masulis, Wang, and Xie, 2007; Appel, 2019). The rationale is that under a robust governance structure, shareholders will be more eager to remove the incumbent managers if they engage in financial fraud (e.g., Persons, 2006; Karpoff, Lee, and Martin, 2008a). Management turnover can come from a new board put in place after a change in control through a takeover (Agrawal, Jaffe, and Karpoff, 1999; Huson, Parrino, and Starks, 2001; Velikonja, 2012). Thus, managers could adopt more ATPs to help entrench their position when they face a more significant threat from whistleblowing. Consequently, the increased agency cost of ATPs associated with managerial entrenchment lowers the value and performance of firms (e.g., Gompers, Ishii, and Metrick, 2003; Bebchuk, Cohen, and Ferrell, 2009; Cremers, Nair, and John, 2009; John, Li, and Pang, 2017; Masulis, Wang, and Xie, 2007, 2020).

However, the bonding hypothesis makes a different empirical prediction. A firm that is subject to a whistleblower's allegations may experience a disruption to its operations. This would impose costs on primary customers because the reactions of the stock market after these allegations result in a significant loss in shareholder value and are likely to attract opportunistic takeovers (Jarrell, Brickley, and Netter, 1988; Shleifer and Vishny, 1990; Cremers, Nair, and John, 2009; Bowen, Call, and Rajgopal, 2010; Lee and Fargher, 2013). Nevertheless, ATPs could isolate the firm from opportunistic takeovers (Laffont and Tirole, 1988; Shleifer and Summers, 1988) and thus can be a powerful, value-increasing mechanism for bonding with key stakeholders such as primary customers (Johnson, Karpoff, and Yi, 2015; Cremers, Litov, and Sepe, 2017; Dey and White, 2021). We hypothesize that managers are more likely to adopt ATPs to bond with key stakeholders when they face a more significant threat from whistleblowing. Consequently, the use of

ATPs increases the value and performance of firms.

To identify the causality between whistleblowing and adopting ATPs, we use a quasi-natural experiment based on US state governments' staggered passage of the False Claims Act (hereafter state FCA). The FCA aims to protect whistleblowers who bring to light fraudulent activities and offer financial rewards to them. Thus, fraud at a firm that a state government invests in via the state's pension funds can be interpreted as defrauding the state government (Rapp, 2007). A firm engaging in fraud can be prosecuted under a state FCA when both of the following are true: (1) a state adopts (or has previously adopted) an FCA covering financial fraud, and (2) at least one state pension fund from that state begins to invest in (or already is invested in) the firm (regardless of whether or not the firm is headquartered in the same state as the pension fund). Thus, since states adopted the FCA at staggered times, the extent to which their FCAs influence a firm varies depending on the size of the investment in the firm by a state pension fund and the variation in the FCA of the state, such as the year of passage and the scope of coverage. In sum, we identify firms that have or have not been subject to a state FCA and examine whether these provisions have causal effects on the increase in ATPs. We expect managers to have stronger incentives to adopt ATPs when faced with increased risk from whistleblowing after exposure to state FCAs.

The state FCAs provide an appealing setting to empirically test the association between the threat from whistleblowing and ATPs. First, the staggered adoption by state governments allows us to study the relationship between the threat from whistleblowing and the increased use of ATPs in a difference-in-differences framework. We do not consider federal laws such as SOX and the SEC's Dodd-Frank whistleblower program because the government simultaneously applies them to all US public firms. Therefore, an appropriate control group is hard to find for these federal laws and isolates their effect from other concurrent events. Second, unlike SOX and the Dodd-Frank Act, which change corporate governance, the state FCAs do not impose any obligations on firms. Therefore, the observed effects on firms are attributable solely to the risk of whistleblowing under those provisions.

We measure ATPs using Bebchuk, Cohen, and Ferrell's (2009) entrenchment index (Eindex) to test our predictions. In the difference-in-differences (DiD) analysis, we use firm-year observations from 1990 to 2017 to examine the effect of adopting state FCAs on the *Eindex*. We find that firms with investment from at least one state pension fund located in a state with an FCA (treated firms) have a significantly higher *Eindex*. Specifically, given that the mean of *Eindex* is 2.976, the exogenous increase in the threat from whistleblowing leads to an average 3.02% to 5.04% increase in *Eindex*. Next, we conduct parallel trend tests and find that the *Eindex* only changes after states adopt FCAs and not before that, indicating no pre-existing difference in the trend of ATPs between treated and control firms.

We further decompose the *Eindex* into its components: a staggered board, a golden parachute, a poison pill, a limit to bylaw amendments, a limit to charter amendments, and requirements for supermajority votes. We estimate the likelihood of adopting one of these components as a function of the state FCAs and control variables using the DiD model. The results indicate that treated firms are more likely to adopt staggered boards, poison pills, and limits on charter amendments after being exposed to state FCAs.

To validate our FCA setting as an appropriate instrument for the increased risk from whistleblowing, we run a series of robustness tests. First, the effect could reflect state pension funds selecting to invest in firms about to increase their ATPs rather than their exposure to an FCA law motivating them to strengthen ATPs. If this is the case, it raises a concern that the changes to the state pension fund's portfolio by the managers are endogenous. To address this concern, we use two strategies. In the first strategy, we estimate the DiD model using a sample that removes the treatment variation from the changes in the pension funds' investments. Thus, we can isolate the treatment variation driven by states' adoptions of FCAs.[†] The result indicates that firms' exposures to state FCAs increase the *Eindex* relative to firms without any changes in exposure to state FCAs. In the second strategy, we conduct a falsification test where we replace the treatment variable with firms' exposures to state Medicaid FCAs (an FCA that protects against Medicaid fraud only); the FCA should not affect the amount of ATPs if the pension funds' investment does not explain the main effect. The result shows no impact on ATPs associated with the Medicaid FCAs.

Second, we consider the heterogenous treatment effect due to the staggered DiD approach. Recent studies have shown that the DiD can be biased when treatments occur at different times for different groups[‡] because the always-treated groups effectively act as control observations for later-treated groups, and these groups act as control observations before they become treated (Cengiz, Dube, Lindner, and Zipperer, 2019; Barrios, 2021; Callaway and Sant'Anna, 2021; Baker, Larcker, and Wang, 2022). To alleviate these concerns, we estimate the main tests by removing firms always treated during the sample

[†] We isolate the treatment effect driven by only the case where the states whose fund already invests in the firm pass an FCA by keeping only the observation with *SPF* = 1 and remove the variable from our model.
[‡] Baker, Larcker, and Wang (2022) categorized the groups of staggered treatment into four categories: always-treated group (when treatment happened before the sample period); earlier-treated group (when treatment happened in the earlier years of the sample period); later-treated group (when treatment happened in the later years; and non-treated group or never-treated group.

period (Callaway and Sant'Anna, 2021). In addition, we estimate the stacked DiD regression (Cengiz, Dube, Lindner, and Zipperer, 2019; Deshpande and Li, 2019). We find our results are robust to these two approaches.

Third, the spillover effects inherent in the Difference-in-Differences (DiD) setting may bias the results (Berg, Reisinger, and Streitz, 2021). Specifically, these biases may stem from firm competition or regional interdependencies. Consequently, the potential influence of the treatment group's exposure to state FCA laws could affect firms' adoption of ATPs. To mitigate this bias, we use Berg, Reisinger, and Streitz's (2021) approach by identifying industry peers as a plausible economic mechanism of the spillover effect arising from exposure to state FCA laws regarding ATPs. The findings demonstrate that the impact of state FCA laws on ATPs remains robust even after accounting for these spillover effects.

Finally, the literature shows that state pension funds have a home bias when they invest (Brown, Pollet, and Weisbenner, 2015). In other words, their investment may be a proxy for the location of a firm's headquarters state instead of the effect of the state's FCA, and thus, their adoption is endogenous. To address this issue, we repeat our main tests using only observations in which states whose pension funds invest in firms not from their headquarters state find the results remain unchanged.

In cross-sectional tests, we first investigate whether the impact of state FCAs on ATP adoption varies systematically with managerial awareness of whistleblowing risk. Managers must be aware of the threat of whistleblowing to implement ATPs in response effectively. Therefore, we conjecture a positive relationship between managerial awareness of whistleblowing threats and ATP adoption. Our results support this conjecture, showing that the increase in ATP adoption following exposure to state FCAs is more significant when firms are more aware of whistleblowing risks.

Then, we examine whether the implications of state FCAs for adopting ATPs vary systematically with the likelihood of whistleblowing. As suggested by the research, whistleblowers are insiders and outsiders (e.g., Smaili and Arroyo, 2019). The labor union's power (a proxy for insider whistleblowers) and financial analysts' coverage (a proxy for outsider whistleblowers) should be positively associated with the likelihood of whistleblowing (Barnett, 1992; Dyck, Morse, and Zingales, 2010; Armitage, Hou, Sarkar and Talaulicar, 2017), and thus incentivize managers to increase their ATPs following their exposure to state FCAs. Our results indicate that the increase in the adoption of ATPs following exposure to state FCAs is more significant when the firms face greater power from unions and are covered by more financial analysts.

Lastly, if managerial entrenchment is the motivation behind firms' decisions to adopt more ATPs, managerial private benefits play a crucial role in encouraging managerial entrenchment (Morck, Shleifer, and Vishny, 1988; Stulz, 1988; Weisbach, 1988; Shleifer and Vishny, 1997), could intensify the positive association between the state FCAs and ATPs. Following Shleifer and Vishny (1997) and Masulis, Wang, and Xie (2009), we use the excessive compensation of CEOs as a proxy for private benefits and find that firms with this compensation adopt more ATPs following exposure to state FCAs. The above results provide supporting evidence for the managerial entrenchment hypothesis.

However, the bonding hypothesis also predicts a positive relationship between the state FCAs and ATPs by arguing there is an effect of the human capital and stakeholder relationship on the association between the state FCAs and ATPs. Therefore, we use the R&D intensity of the customer's industry and principal customers to measure human capital and the stakeholder relationship. However, we find that the association between the state FCAs and ATPs does not vary systematically with the R&D intensity and principal customers, which indicates there is no evidence to support the bonding hypothesis.

Furthermore, testing the effect of increased ATPs due to the passage of state FCAs on the valuation and performance of firms enables us to further distinguish between these two hypotheses. Although the evidence is mixed, our results indicate that firms adopting more ATPs lead to lower valuation and performance after their states adopt FCAs. This evidence supports the managerial entrenchment hypothesis over the bonding hypothesis.

This study contributes to the literature in two ways. First, it contributes to the literature on whistleblowing. That literature focuses on the effectiveness of whistleblowing in the detection of corporate fraud (Karpoff, Lee, and Martin 2008a, 2008b; Dyck, Morse, and Zingales, 2010; Wilde, 2017; Berger and Lee, 2022). By contrast, we investigate how managers respond to their exposure to a threat from whistleblowing. Specifically, firms have more substantial incentives to adopt more ATPs when managers are more likely to whistleblow. In particular, the increase in ATPs induced by the threat from whistleblowing can hurt a firm's operating performance, which suggests a dark side to the adoption of whistleblowing laws.

Second, this study provides evidence on the determinants of corporate governance structures. Several empirical studies offer specific predictions regarding the determinants of these structures (e.g., Johnson, Karpoff, and Yi, 2015; Appel, 2019; Dey and White, 2021; Ahn, Patatoukas, and Solomon, 2022; Foroughi, Marcus, Nguyen, and Tehranian, 2022). For example, Appel (2019) and Foroughi, Marcus, Nguyen, and Tehranian (2022) find that the dilution of the litigation rights of shareholders (represented by the staggered adoption of universal demand laws) is associated with an increase in ATPs that is commonly opposed by shareholders. However, Ahn, Patatoukas, and Solomon (2022) find no evidence of the association between the universal demand laws and ATPs and thus question the existence of a cause-and-effect link between these two variables. In contrast to the above agency argument, Johnson, Karpoff, and Yi (2015) and Dey and White (2021) argue that the use of ATPs is likely to entrench the firm's management by reducing the outside interferences that will change the firm's operating strategy and impose costs on its stakeholders. This study complements this line of research by testing the managerial entrenchment hypothesis and the bonding hypothesis through state FCAs.

The rest of the study is organized as follows: In Section 2, we review the literature and develop testable hypotheses. We discuss the data and method in Section 3. Section 4 presents the main empirical results, and Section 5 presents the cross-sectional tests. In Section 6, we discuss the competing hypotheses and conduct a set of robustness checks in Section 7. Section 8 contains the conclusion.

2. Literature Review and Hypothesis Development

2.1. Review of whistleblower laws

Whistleblowers face tremendous social and economic pressures when providing evidence of fraud and wrongdoings by their employers to regulators (Rapp, 2012; Call, Martin, Sharp, and Wilde, 2018). Therefore, laws and regulations on whistleblowers have been put in place to provide a certain level of protection and/or economic reward to encourage whistleblowing. The protection of whistleblowers is mainly in the form of laws that prohibit any demotion, discrimination, or punishment against them (such as 31 U.S.C. 3730(h) of the federal FCA and anti-retaliation provisions in state-level FCAs; Section 806 of Sarbanes Oxley Act (SOX) of 2002; 15 U.S.C. 7201 of Dodd-Frank Act). Meanwhile, financial incentives to report wrong-doings come from the *qui tam* law that allows those who report to claim a percentage of the recovered fraud money as a reward should the lawsuits be successful. While protection against retaliation is important, monetary incentives play a more critical role in motivating whistleblowers (Dyck, Morse, and Zingales, 2010; Andon, Free, Jidin, Monroe, and Turner, 2018).

The oldest whistleblower law in the US is the federal False Claims Act adopted in 1863. Its *qui tam* bounty provision and a "dual plaintiff" structure allowing a citizen to file a lawsuit on behalf of the government and to receive a portion of the recovered funds (specifically, 15% - 25%) were adopted in 1986. Nevertheless, at the federal level, the FCA cases are mostly Medicaid ones (Dyck, Morse, and Zingales, 2010) since securities fraud against the shareholders was not considered harmful to the federal government. Although the passage of SOX in 2002 did include corporate frauds, it only implemented an anti-retaliation clause, which has received criticism as not protective enough for the employee to risk their careers to become whistleblowers (Dworkin, 2007; Dyck, Morse, and Zingales, 2010; Rapp, 2012). The Dodd-Frank Act of 2010 took a step further by adding a *qui tam* bounty provision at the federal level for the first time in Section 21F that provides up to 30% of recoveries from financial frauds to the whistleblowers who report to the SEC.

State governments view financial fraud differently. Unlike the federal government, state governments can invest their funds in publicly traded companies through state retirement or pension funds. Financial fraud involving the state pension funds is subject to false claims against the state government (Rapp, 2007, 2010). In this case, whistleblowers can claim financial rewards under that state's FCA by providing evidence of securities fraud involving the state pension funds' investments as long as the state has an appropriate qui tam provision.

States have adopted their own versions of FCAs modeled after the federal FCA at staggered times since 1987, and their FCAs vary in terms of fraud coverage. As of 2015, 19 states and the District of Columbia had FCAs that covered fraud in general, including financial fraud (hereby called general FCAs), whereas 15 states have FCAs that protect against only Medicaid fraud (hereafter Medicaid FCAs).[§] The remaining 16 states have not yet adopted FCAs. This study focuses on the general FCAs because whistleblowing on financial fraud can be rewarded only when a state has a qui tam provision.[**] By contrast, a Medicaid FCA is limited to Medicaid fraud, such as paying kickbacks to pharmacies or doctors using money funded through the Medicaid program of the state government. The state FCAs provide an ideal setting for us to examine the effect of whistleblower threats on firms' ATPs. It provides a cleaner identification because the court's positions regarding the adoption of the federal FCA vary over time and across states, which generates a plausibly exogenous source of variation in the whistleblower threats.

2.2. Whistleblower threat and management turnover

Studies have shown that whistleblowing is an effective component of corporate governance to detect and report the suspicious misconduct of firms (Karpoff, Lee, and Martin, 2008a, 2008b; Dyck, Morse, and Zingales, 2010; Wilde, 2017; Call, Martin, Sharp, and Wilde, 2018; Dey, Heese and Perez-Cavazos, 2021; Berger and Lee, 2022). Firms that

[§] See Appendix B for the details.

[**] General FCAs cover both Medicaid fraud and non-Medicaid fraud including financial fraud. Bucy, Diesenhaus, Raspanti, Chestnut, Merrell, and Vacarella (2010) report that California, which adopted a general FCA in 1987, has recovered approximately $254 million ($353 million) from (non-)Medicaid cases since 1999.

experience whistleblowing could have remarkable financial losses and suffer adverse consequences for them and their managers. Following financial frauds, firms experience severe declines in stock valuations, and their CEOs face higher dissatisfaction from shareholders (Biggerstaff, Cicero, and Puckett, 2015; Eckbo, Thorburn, and Wang, 2016). These negatives make a firm more likely to be a takeover target (Fishman and Hagerty, 1992; Dow and Gorton, 1997; Goldstein and Guembel, 2008; Levit and Malenko, 2011; Edmans, Goldstein, and Jiang, 2012; and Bereskin, Campbell, and Kedia, 2020) that increases the likelihood of the new board replacing the CEO.

2.3. Whistleblower threat and antitakeover provisions

The entrenchment hypothesis views takeover deterrence as a self-interest, value-decreasing activity for entrenching the management and preventing a hostile takeover. DeAngelo and Rice (1983) develop the managerial entrenchment hypothesis, stating that ATPs deteriorate the principal-agent conflict between shareholders and managers. Entrenched managers adopt ATPs for corporate control (Scharfstein, 1988) to protect their high level of compensation (Borokhovich, Brunarski, and Parrino (1997) and to enjoy a quiet life (Bernstein, 2015).

To measure takeover deterrence, Gompers, Ishii, and Metrick (2003) and Bebchuk, Cohen, and Ferrell (2009) establish their versions of indexes based on ATPs. Specifically, Gompers, Ishii, and Metrick (2003) create the governance index (Gindex) established via 24 ATPs. Bebchuk, Cohen, and Ferrell (2009) identify six ATPs composed of four provisions related to the limiting of shareholders’ rights and two provisions on potential takeovers. These two indexes are widely used in measuring the effect of takeover deterrence.

Since the adoption of an FCA increases the threat from whistleblowing, its effect on management turnover is likely to be positive, thereby increasing the manager's incentives to entrench their position which could prove harmful to firm performance (dark-side)

On the other hand, *the bonding hypothesis* argues for the bright-side effect of takeover defenses (Karolyi, 2012; Johnson, Karpoff, and Yi, 2015). Alteration of firms' operating strategy (e.g., from takeovers) could potentially harm their performance and value by preventing them from continuing long-term relationship-related investment with key stakeholders. Hence, ATP adoption is the effective way for the firm to "bond" their contractual performance with such counterparties (Knoeber, 1986; Shleifer and Summers, 1988; Karolyi, 2012; Cen, Dasgupta, and Sen, 2015; Johnson, Karpoff, and Yi, 2015; Tsang, Yang and Zheng, 2022).

Increased whistleblowing threats from FCAs could potentially alter the firm's business strategy through management turnover or takeovers. Thus, the bonding hypothesis also suggests increased ATPs adoption of exposed firms. However, such an increase could not result from managerial entrenchment but from firms' commitment to their stakeholders and potential whistleblowers of better governance. In this case, such ATP adoption should have a positive impact on firm value subsequently.

From the previous argument, we propose our main hypothesis:

H1: Firms facing a higher risk from whistleblowing increase their antitakeover provisions (ATPs).

3. Data and Methodology

3.1. Data and variables

The sample consists of all firms in the COMPUSTAT and CRSP databases from 1990 to 2017. Equity holdings of state pension funds are collected from Thomson Reuters's 13F institutional holdings. We exclude financial firms (SIC codes 6000-6999) and firms in the healthcare industry (SIC codes 2830-2839, 3693, 3840-3859, and 8000-8099) to ensure that the healthcare companies that were subjected to Medicaid FCAs do not drive the treatment effect of state FCAs.

Data on ATPs are collected from the Institutional Shareholder Services (ISS) governance database (formerly RiskMetrics). Following Bebchuk, Cohen, and Ferrell (2009), we use the entrenchment index (*Eindex*) as a proxy for corporate governance. *Eindex* consists of the six provisions most frequently targeted by non-binding shareholder proposals.[††] The value of zero for the index depicts a low entrenchment level and the absence of all ATPs, while a maximum value of six for *Eindex* demonstrates an extremely high level of entrenchment for firm executives and the existence of all ATPs.

3.2. Research design

We examine the causal effect of the FCAs on corporate governance by exploiting the level of firms' exposure to state FCAs. Using panel data, we estimate the following regression model:

$$Eindex_{i,t} = \alpha_0 + \alpha_1 FCA\_G_{i,t} + \alpha_2 SPF_{i,t-1} + \beta' Controls_{i,t-1} + \gamma_i + \pi_t + \varepsilon_{i,t}, \qquad (1)$$

where *Eindex* is the index of the six ATPs as in Bebchuk, Cohen, and Ferrell (2009). *FCA_G* is a dummy variable that is set to one when a firm is influenced by at least one state's general FCA due to investment from that state's pension funds. *SPF* is an indicator variable that is assigned a value of one if a firm's shares were held by any state pension

[††] Four of the provisions (staggered boards, super-majority voting requirements for mergers, limits on shareholder bylaw amendments, and limits on shareholder charter amendments) restrict shareholders' voting power. The remaining provisions, namely poison pills and golden parachutes, limit the size of blockholders' positions and insulate managers from the economic risks associated with takeovers, respectively.

fund in the previous year. For firms under treatment, *FCA_G* increases from 0 to 1 either when the initial state pension fund investing in the previous year is from a state that had already enacted a general FCA, or when a state where the firm's shares are already held by pension funds passes a general FCA in the current year. In the first scenario, both *FCA_G* and *SPF* switch from 0 to 1 simultaneously when an FCA state pension fund makes a purchase, whereas in the second scenario, *FCA_G* changes subsequent to *SPF*. Consequently, including the *SPF* variable can help distinguish the impact of being exposed to an FCA from the influence of shifts in pension fund ownership. *Controls* are the vector of control variables; γ_i and π_t denote firm and year fixed effects, respectively; and $\varepsilon_{i,t}$ is the error term.

We define *SPF* based on the lagged year's investment to address the selection issue that pension funds might change their portfolio in the expectation of FCA adoption in their states. We assume that in the year prior to the adoption of state FCAs, pension fund managers did not expect the rule change. We drop firm-year observations that were initially affected by the FCAs but become unaffected because existing state pension funds sell their investment in the firm.[‡‡]

We add control variables that could affect the ATPs for reasons unrelated to the risk of whistleblowing. Following Straska and Waller (2010), Chen, Chen, Schipper, Xu, and Xue (2012), Appel (2019), and Dey and White (2021), the control variables are the size (*Size*), cash ratio (*Cash ratio*), age (*Age*), growth opportunities (*Q*), return on assets (*ROA*), research and development expenditures (*R&D*) of firms. We also control for some internal

[‡‡] For robustness, we keep such observations and code them as treated firm-years, because we expect that firms remain subject to FCAs once state pension funds invest in them even after the funds leave the firm. We find the results are similar.

corporate governance variables. These are the previous year's board characteristics, namely the number of directors (Board size), the ratio of female directors (Female dir ratio), and the ratio of independent directors (*Indep dir ratio*) as in Chemmanur, Jordan, Liu, and Wu (2010). Also, we used CEO characteristics such as the tenure of the CEO (*CEO tenure*), a CEO who is also the chairman of the board of directors (*CEO duality*), and a dummy of having a female CEO (*Female CEO*) as in Hermalin (2005) and Li (2014). Both firm and year-fixed effects are added to control for unobserved time and entity-invariant variables.[§§] Firm fixed effects control for time-invariant characteristics associated with corporate governance that allow for the exploitation of within-firm variation in the pension fund's investment through which firms are exposed to whistleblowing threats. Year-fixed effects control for changes over time in factors other than whistleblower laws that affect corporate governance. All variables are defined in Appendix A and winsorized at the 1st and 99th percentiles.

3.3. Descriptive Statistics

Table 1 summarizes the statistics for the main variables in this study. The mean value of Eindex is 2.976, with a median of 3. Overall, an average firm in our sample has around three out of the six ATPs. The mean values of FCA_G and SPF are 0.843 and 0.976, respectively. These values mean around 84% of firm-year observations from 1990 to 2017 were affected by state FCAs.

Regarding firm characteristics, the mean natural logarithm of age is 3.097, which translates into an average of 22 years in operation for firms in our sample. On average, sample firms have total assets of approximately 2.4 billion USD with a nearly 10% cash-

[§§] As a robustness check, we reestimate the regressions using firms' headquarter state × year fixed effects instead of year fixed effects and find similar results.

to-assets ratio. They also spend a mean value of 2.8% of total assets on research and development and are profitable with a ROA of 3.4% at the mean (8.68% in median).

[Insert Table 1 here]

Regarding board and CEO characteristics, the normal board size for sample firms is nine directors (natural logarithm of 2.205) with an average of 12.4% female directors and nearly 75% independent directors (with a standard deviation of 0.103 and 0.147, respectively). The mean value for the CEO's tenure is about nine years. Of the firms, 77.3% have a CEO who is also the chairman of the board of directors, and 2.2% of firms have a female manager.

Table 2 depicts the Pearson-Spearman correlation matrix of the main variables in this study. Pearson (Spearman) correlations are below (above) the diagonal. Overall, the positive and significant correlation between *FCA_G* and *Eindex* is consistent with our hypothesis. Another noteworthy point is that none of the correlation coefficients among the explanatory variables are exceptionally high, which mitigates the concern for multicollinearity in our work.

[Insert Table 2 here]

4. Empirical Results

4.1. Difference-in-differences (DiD) approach

To study the effect of the FCA and ATPs, we estimate Equation (1) and report the results in Table 3. Columns (1) and (2) show the results for the two main binary variables *FCA_G* and *SPF*, respectively. The coefficient for *FCA_G* is positive and statistically significant at the 1% level, which indicates firms adopt more ATPs after states pass FCAs. After adding the full set of control variables as well as the firm and year fixed effects to column (3), the coefficient for *FCA_G* is 0.090 and significant at the 5% level. This coefficient is consistent with our hypothesis that firms that are exposed to the state FCA adopt more ATPs. In terms of economic significance, given that the mean of *Eindex* is 2.976, an exogenous increase in the threat from whistleblowing leads to an increase in *Eindex* by 3.02%.

[Insert Table 3 here]

In our empirical setting, treatment firms are defined as either those in states whose pension funds invest in them and pass an FCA law or those in a state that has already passed the FCA and then pension funds invest in them. The concern arises in the latter case as the decision to invest and change the state pension fund's portfolio relies mostly on its manager's endogenous choice. Therefore, one might argue that the change in ATPs may not reflect the true causal effect of the state FCA. To solve this problem, we isolate the treatment effect driven by only the first case by keeping only the observation with *SPF* = 1. We report the results in column (4) that indicate the coefficient for *FCA_G* remains significantly positive.

In column (5), we use a matched sample to control for underlying differences between the treated and control firms. We generate the matched sample by matching each treated firm at the fiscal year-end immediately before the passage of state FCA laws with a control firm based on the control variables as specified in Equation (1) with no replacement. The coefficient for *FCA_G* is 0.15 and statistically significant at the 1% level. The estimation result indicates that the effect of the state FCA laws on firms' ATPs is not driven by observable heterogeneity. Economically, with a mean *Eindex* of 2.976, an exogenous increase in whistleblower risk raises the *Eindex* by 5.04%.

For the control variables, the results indicate that older firms, firms with higher research and development expenditures, and firms with a larger board size are more likely to adopt more ATPs. Meanwhile, firms with long-tenured CEOs adopt fewer ATPs. The evidence is consistent with the findings in other studies (e.g., Straska and Waller, 2010; Chemmanur and Tian, 2018).

4.2. Decomposition of the entrenchment index

The use of the *Eindex* to measure corporate governance has some weaknesses. For instance, aggregating dissimilar forms of ATPs lacks a legal rationale (Klausner, 2013; Catan and Kahan, 2016). Furthermore, Larcker, Reiss, and Xiao (2015) find measurement errors in the ISS summaries, especially for golden parachutes and supermajority voting provisions, two key components of the *Eindex*. Given the potential weaknesses suggested by these studies, we examine the effect of state FCAs on each of the *Eindex*'s constituent components. Focusing on each component circumvents aggregation concerns and reduces the concerns that our findings may be simply due to measurement error.

Since all the components in the *Eindex* are binary, we estimate a probit model as follows: [***]

$$\Pr\left(Eindex\ component_{i,t} = 1\right) =$$

$$\Phi(\alpha_0 + \alpha_1 FCA\_G_{i,t} + \alpha_2 SPF_{i,t-1} + \beta' Controls_{i,t-1} + \gamma_i + \pi_t), \qquad (2)$$

where *Eindex component* is a given component in the *Eindex* for firm *i* in year *t*. Specifically, *Eindex* contains six components: a staggered board, a golden parachute, a poison pill, a limit to bylaw amendments, a limit to charter amendments, and requirements for supermajority voting. Φ is the cumulative distribution function of the standard normal

[***] We also rerun the probability analysis using logistic models and the results are the same.

distribution. The other variables are defined in Equation (1).

[Insert Table 4 here]

Table 4 presents the estimates from the probit regression. We find that *FCA_G* is positively correlated with a staggered board, a poison pill, and a limit to charter amendments and is significant at the 5% level or better. We report a marginal effect for *FCA_G* that represents the change in the probability of adopting an ATP after a change in whistleblowing risk from before to after the passage of an FCA, holding all other variables constant at their mean values. A discrete change in *FCA_G* from zero to one increases the probability of adopting staggered board, poison pill, and limit to charter amendments by 7.1%, 3.1%, and 8.4%, respectively. The relations between *FCA_G* and the other three provisions are not significant at the conventional statistical levels.

Based on Table 4, we conclude that our findings are not limited to a single component in the *Eindex* but are evident for a staggered board, a poison pill, and a limit to charter amendments. Moreover, these findings make it unlikely that our overall conclusions are driven by aggregation concerns or possible errors in measuring the components in the *Eindex*.

4.3. Robustness of DiD approach.

In this subsection, we undertake a set of robustness tests to ensure the validity of our DiD results. First, to assess the validity of the assumption of parallel trends, we test whether our results are driven by the explicit trend in the ATPs of the treated firms before the passage of the state FCAs. Second, we conduct a falsification test to rule out the alternative explanation that state pension funds invested in firms that were about to increase their ATPs rather than those increasing their ATPs because of FCA exposure. Third, we use two strategies for the issue of a heterogenous treatment effect due to the staggered DiD: 1) estimating the DiD by removing firms that are always treated during the sample period, and 2) estimating the staggered DiD. Finally, we follow Berg, Reisinger, and Streitz's (2021) suggestion to address the spillover effects that arise with the DiD setting.

4.3.1. Parallel trends assumption

The parallel trends assumption assumes that in the absence of events, treated and control firms should have common trends for the outcome variable (Roberts and Whited, 2012). However, the results may be driven by the explicit trend in the treated firms' ATPs before the state FCAs. Thus, following Bertrand and Mullainathan (2003) and Sun and Abraham (2021), we address this concern by estimating a dynamic DiD model as follows:

$$Eindex_{i,t} = \sum_{\substack{j=-5 \\ j\neq -1}}^{j=+5} \alpha_j FCA\_G_j + \theta SPF_{i,t-1} + \beta' Controls_{i,t-1} + \gamma_i + \pi_t + \varepsilon_{i,t}, \quad (3)$$

where *Eindex* is the entrenchment index constructed by Bebchuk, Cohen, and Ferrell (2009) of firm *i* in year *t*. FCA_G_j is an interaction variable between *FCA_G* and time dummy indicating *j*-year before or after the passage of state FCA law. Since we include the time and entity fixed effect, we remove the dummy in the *t – 1* period to avoid perfect multicollinearity. *SPF* is a binary variable that equals one if at least one state pension fund invests in the firm in the lagged year. *Controls* are the vector of control variables; γ and π denote firm and event year fixed effects, respectively; and ε is the error term. We report the results in Table 5.

[Insert Table 5 here]

In Table 5, the coefficients of FCA_G_j for the years preceding state FCA exposure are not statistically significant, while those for the first and fourth years following the exposure are both positive and statistically significant. Joint tests confirm that the effect of state FCA laws on the entrenchment index manifests only after the passage of these laws. Overall, the results align with the parallel trend assumption. Additionally, the estimated coefficients of state FCA laws from Equation (3) are reported in Figure 1.

[Insert Figure 1 here]

Panel A of Figure 1 shows the difference in the entrenchment index between the treatment and control groups over an 11-year event window surrounding the passage of state FCA laws. It shows that the difference between the treatment and control groups is stable in the three years leading up to the laws, suggesting that there are no pre-trends present for ATPs. Panel B reports the results using the sample conditional on SPF = 1. We observe a similar trend in this subsample. Overall, this piece of evidence indicates that the parallel trend assumption is satisfied.

4.3.2. Falsification test

From the definition of treatment firms, *FCA_G* equals one not only when states whose pension funds invest in a firm pass an FCA but also when their pension funds invest in the firm while having an FCA. Therefore, a selection problem may arise if state pension funds choose to invest in firms due to the characteristics of the firms themselves (e.g., high risk but high investment rate in capital expenditure or R&D) that may lead to a future change in the adoption of ATPs. To mitigate this concern, we use the Medicaid FCA. Further, only the general FCA can be a threat to entrenched managers by exposing financial wrongdoings and fraud. Therefore, Medicaid FCAs should not be able to affect firms' adoption of ATPs. As we apply the same empirical setting to Medicaid FCAs, if the characteristics of firms drive the adoption of ATPs, we expect to see the same treatment effect in both General and Medicaid FCAs. On the contrary, if it is truly the General FCA that drives the results, we will not see any significant effect on firms with exposure to Medicaid FCAs. The results are shown in Table 6.

[Insert Table 6 here]

In column (1), we report the benchmark result for general FCAs from our baseline model in Table 3. In columns (2) and (3), we retain only those observations that experienced changes in exposure to both general and Medicaid FCAs during our sample period in different years to isolate the effect of each type of FCA. Although this isolation greatly reduces our sample size to only 2,263 observations (in comparison to 15,535 in the baseline model), we are able to compare the effect of different types of law on a constant sample. We use a new binary variable, *FCA_M*, that equals one for firms that are exposed to a Medicaid FCA. We use the Chi-square test for the equality of the coefficients to test the null hypothesis that the treatment effects of both general and Medicaid FCAs are the same.

In column (2), the coefficient for *FCA_G* continues to be positive and statistically significant at the 5% level with a higher effect at 0.185. Meanwhile, the coefficient for *FCA_M* in column (3) is not statistically or significantly different from zero. A Chi-square test rejects the null hypothesis at the 10% significance level (*p*-value of 0.062) that there are equal coefficients for *FCA_G* and *FCA_M*. Overall, the results support our main findings that the FCA (*FCA_G*) significantly affects the adoption of ATPs.

4.3.3. Heterogenous treatment effect

The recent literature has argued that an estimation of a staggered DiD could be biased because treatments occur at different times for different groups (which could be categorized into always-treated, early adoption, later adoption, and non-treatment groups as in Baker, Larcker, and Wang, 2022). The bias happens when the always-treated group takes the role of control for the later adoption group, which violates the parallel trend assumption (Cengiz, Dube, Lindner, and Zipperer, 2019; Barrios, 2021; Callaway and Sant'Anna, 2021; Baker, Larcker, and Wang, 2022). To account for this heterogeneous treatment effect, we use two strategies. First, we re-estimate the regression models specified in Table 3 by removing firms always treated during the sample period (Callaway and Sant'Anna, 2021). The results are reported in Table 7. The coefficients for *FCA_G* are positive and statistically significant at the 5% level or better across models.

[Insert Table 7 here]

Second, we follow Cengiz, Dube, Lindner, and Zipperer (2019) and Deshpande and Li (2019) to create our Stacked difference-in-differences (DiD) approach. Notably, we first stack cohorts of treatment and control firms in an event study style. Treatment firms are firms that have been exposed to a general FCA for the first time during the sample period (always-treated firms during the period are excluded), while the control firms have not experienced a general FCA (only non-treatment firms are selected as clean controls as in Sun and Abraham, 2021). Next, we perform propensity score matching (PSM) to match the control and treatment firms using data from the year immediately before the treatment events. We use the nearest propensity scores and a caliper of 0.01 to select control firms in each cohort that are comparable to the control variables used in Equation (1) for the treatment firm. The control firms are matched with replacement. We run the two-way fixed effect DiD approach using the stacked data with an event year fixed effect and report the results in Table 8. The coefficients for *FCA_G* are significantly positive across the models, which indicates that firms respond to exogenous increases in the risk of whistleblowing by adopting more ATPs.

[Insert Table 8 here]

Based on the findings in Tables 7 and 8, our results are robust to the heterogeneous treatment effect.

4.3.4. Spillover effects

A recent paper from Berg, Reisinger, and Streitz (2021) documented that the spillover effects arise naturally with the difference-in-differences (DiD) setting. Specifically, the spillover-induced bias arises through firm competition or local interdependencies among firms. Therefore, the potential spillover effect of the treatment group for exposure to state FCA laws could influence firms' adoption of antitakeover provisions. To address this bias, we follow their suggestion and identify the plausible economic mechanism and dimension of the potential spillover effect from the exposure of state FCA laws on antitakeover provisions. Industry peers seem to be a common dimension, as suggested by the previous literature (Servaes and Tamayo, 2014; Bhojraj, Sengupta, and Zhang, 2017; Dey and White, 2021). Then, we construct a full spillover effect model as follows:

$$Eindex_{i,g,t} = \alpha_0 + \alpha_1 FCA_{G_{i,g,t}} + \alpha_T \overline{FCA_G}_g \times FCA_{G_{i,g,t}} + \alpha_C \overline{FCA_G}_g \times \left(1 - FCA\_G_{i,g,t}\right) + \alpha_2 SPF_{i,t-1} + \beta' Controls_{i,t-1} + \gamma_i + \pi_t + \varepsilon_{i,g,t}, \quad (4)$$

where $Eindex_{igt}$ is the entrenchment index of firm i in the spillover group g at time t. $\overline{FCA\_G_g}$ is the average treatment effect of all other firms located in the same spillover group g, excluding the firm i itself. Other variables are defined in the previous equations. γ and π denote firm and year fixed effects, respectively, and ε is the error term. In this model, α_1 depicts the direct treatment effect. α_T shows the spillover effect on treated units. If the whistleblowing risk is a mechanism through which the state FCA laws change ATPs, we expected both coefficients to be statistically significant. α_C illustrates the spillover effect on control units. Similarly, a significant coefficient indicates that the industry competition mechanism exists. We report the results of the spillover analysis in Table 9:

[Insert Table 9 here]

In Table 9, columns (1) and (1) report the results of full spillover model settings with a full sample and a sample with SPF=1 only, respectively. In all specifications, the coefficients of *FCA_G* are statistically significant at the 5% level, suggesting that the state FCA laws directly impact ATPs. In addition, both spillover effect variables are statistically significant at the 10% level. The results thus indicate the impact of state FCA laws on ATPs is robust in controlling for the spillover effect of the laws.

5. Cross-Sectional Tests

In this section, we explore the factors likely to facilitate or mitigate the effect of a state FCA on ATPs. In particular, we examine how the managerial awareness of the whistleblowing threat, the likelihood of whistleblowing, managerial private benefits, and stakeholder relationships influence that effect.

5.1. Managerial awareness of whistleblowing threat

Managers can implement ATPs in their firm because whistleblower risk arises. The key to this hypothesis is that managers need to be aware of the threat of whistleblowing. Thus, it would be helpful to discuss the degree to which managers know about the threat of whistleblowing from a given state.

In this study, we use state pension fund ownership as a necessary condition when we define treatment firms because they could file financial fraud lawsuits on behalf of the state government. Therefore, it is reasonable to expect the managers to be more responsive to whistleblowing risk with ATP adoption if the firms have a noticeable state pension fund shareholder. To test this conjecture, we use a 1% pension fund ownership threshold to divide the sample into two subsamples: (1) high-awareness firms, where the state pension fund holds at least 1% of the firm's shares, and (2) low-awareness firms, where the state pension fund holds less than 1% of the firm's shares. The 1% threshold determines a shareholder's eligibility to receive notice, vote on proposals at annual and special meetings, and implement majority voting resolutions (Securities Exchange Act Rule 14a-8(b)). Consequently, surpassing this threshold grants shareholders significant influence in the company's proxy statements (Fairfax, 2009). We estimate the baseline regression separately for the high- and low-awareness subsamples, and the results are reported in Table 10.

[Insert Table 10 here]

In Table 10, the coefficients of *FCA_G* are positive for both subsamples. However, only the coefficient in the high-awareness subsample is statistically significant and is much larger than that of the low-awareness subsample. The Chi-square test for equality of coefficients reveals a strong and significant difference, indicating that the impact of state FCA laws is significantly greater for firms with higher managerial awareness of whistleblower threats.

5.2. Likelihood of whistleblowing

As suggested in the research, whistleblowers are not only insiders but also outsiders (e.g., Smaili and Arroyo, 2019). Executives of organizations with union employees anticipate a higher overall level of internal whistleblowing because employees consider unionization as instrumental in the resolution of their dissatisfaction with various aspects of their jobs (Deshpande and Fiorito, 1989). Insider whistleblowers can thus be measured by unionization. Moreover, analysts have more experience and expertise than individual investors, which allows them to track firm performance, identify abnormal patterns, offer early warnings, and even act as whistleblowers about value-destroying managerial misconduct (Lang, Lins, and Miller, 2004; Yu, 2008; Dyck, Morse, and Zingales, 2010). Accordingly, financial analysts' coverage can be a suitable proxy for outsider whistleblowers. We expect that these two proxies will be positively associated with the likelihood of whistleblowing, thus incentivizing managers to adopt more ATPs following exposure to state FCAs.

To test this hypothesis, we estimate Equation (1) separately for the subsamples of firms facing high and low levels of the likelihood of whistleblowing by insiders and outsiders. The high or low likelihood of insider (outsider) whistleblowing is defined as a firm's unionization rate (analyst coverage) in the top and bottom half of its empirical distribution.

Following Klasa, Maxwell, and Ortiz-Molina (2009) and Chen, Kacperczyk, and Ortiz-Molina (2011), the unionization rate is measured as the percentage of workers in a firm's industry covered by collective bargaining agreements as a proxy for labor unions' ability to affect the firm's behavior. The data on the unionization rate come from the Union Membership and Coverage Database.

We follow Doukas, Kim, and Pantzalis (2005) and use two measures of excess analyst coverage that are defined as the actual coverage of a firm beyond the expected coverage for similar firms. Positive (negative) excess coverage indicates strong (weak) coverage above (below) the market standard for such a firm. The first measure is a simple difference between the firm's actual number of analysts following and the expected coverage, that is, the size-adjusted average analyst coverage of the firm's industry. The second measure of excess analyst coverage is estimated as the residuals of the regression model (as in Hong. Lim and Stein, 2000) in each year:

$$\ln(AC)_{i,t} = \gamma_0 + \gamma_1 \ln(MVE)_{i,t} + \sum_j \gamma_{j,i,t}(IND)_{j,t} + \varepsilon_{i,t}, \tag{5}$$

where $AC_{i,t}$ is the number of analysts' coverage for firm *i* in year *t*; *MVE* is the market value of the firm's equity; *IND* is the two-digit SIC industry dummy.

[Insert Table 11 here]

We report the results in Table 11. Panels A and B give those for the likelihood of insider and outsider whistleblowing represented by the unionization rate and the number of analysts covering, respectively. The even- (odd-) numbered columns show the results for the association between the state FCAs and ATPs for the high (low) likelihood of whistleblowing. Panel A shows that the coefficient for *FCA_G* is 0.102 (*t*-statistic = 2.00) for the high unionization subsample and 0.076 (*t*-statistic = 1.25) for the low unionization subsample. The difference in *FCA_G* coefficients between the two subsamples is statistically significant at the conventional level, as demonstrated using an *F*-test.

In Panel B, the coefficients for *FCA_G* are more positive for a high number of analysts covering than for a low number. The differences in the coefficients for these two subsamples are statistically significant at the 5% level or better, as demonstrated by *F*-tests. Overall, the results in Table 11 indicate that the effect of state FCAs on ATPs is stronger for firms facing a greater likelihood of whistleblowing than for those facing a smaller one.

5.3. Managerial private benefits

If managerial entrenchment motivates firms' decisions to adopt ATPs; then managerial private benefits should intensify the positive association between the state FCA and ATPs.

Following Shleifer and Vishny (1997) and Masulis, Wang, and Xie (2009), we use the excessive compensation of the CEO as a proxy for private benefits and predict that firms with more excessive compensation for CEOs adopt more ATPs following exposure to state FCAs. We adopt the method in Core, Guay, and Larcker (2008) and run the following model to obtain the excessive compensation as the residuals:

$$\begin{aligned}\ln(CEO\ TDC)_{i,t} = {} & \beta_0 + \beta_1 \ln(Tenure)_{i,t} + \beta_2 \ln(Sales)_{i,t-1} + \beta_3 S\&P500_{i,t-1} \\ & + \beta_4 BM_{i,t-1} + \beta_5 Stockreturns_{i,t} + \beta_6 Stockreturns_{i,t-1} \\ & + \beta_7 ROA_{i,t} + \beta_8 ROA_{i,t-1} + IND_{j,t} + \varepsilon_{i,t},\end{aligned} \tag{6}$$

where $CEO\ TDC_{i,t}$ is CEO total compensation of firm *i* in year *t*. *Tenure* is the CEO's tenure as of year *t*. *Sales* is the sales. *S&P500* is an indicator for the firm included in the S&P 500 index. *BM* is the firm's book-to-market ratio. *Stockreturn* is the firm's stock returns, and *ROA* is the return-on-assets ratio of the firm. The explanatory variables in the models are those that could explain the compensation and represent the predicted or expected total pay of a CEO given such determinants.

To test this hypothesis, we estimate Equation (1) separately for the subsamples of CEOs with high or low excessive compensation. A high (low) excessive compensation is defined as a CEO's excessive compensation at the top (bottom) half of its empirical distribution.

[Insert Table 12 here]

Table 12 presents the results of this analysis. While the coefficient for *FCA_G* is positive and statistically significant at the 1% level for the high excessive compensation subsample, the coefficient for *FCA_G* for the low excessive compensation subsample is negative and insignificant. The *F*-test shows that the state FCA laws produce a stronger effect on the adoption of ATPs in firms with high managerial private benefits compared to that in firms with low managerial private benefits. The result provides supporting evidence for the managerial entrenchment hypothesis.

6. Discussions

The evidence so far supports the entrenchment hypothesis that CEOs strategically use ATPs after exposure to state FCAs, which increases their entrenchment. However, it is also possible that adopting ATPs increases the CEO's commitment to a business strategy that cannot easily be reversed via an outside takeover; that is, the bonding hypothesis (Knoeber, 1986; Shleifer and Summers, 1988; Cremers, Nair, and Peyer, 2008; Johnson, Karpoff, and Yi, 2015; Dey and White, 2021). To distinguish these two hypotheses, we propose three empirical tests. First, the effect of relationship-specific investments on the association between state FCAs and ATPs. Then, how increased ATPs induced by the passage of state FCA laws affect firm valuation and performance. Lastly, is disclosure quality affected by increased ATPs induced by the passage of state FCA laws?

6.1. Relationship-specific investments (RSIs)

The bonding hypothesis suggests that ATPs can be beneficial for a firm's shareholders by allowing managers to focus on long-term value creation rather than short-term gains to fend off potential takeovers, implying that the bonding of a CEO to stakeholders that have made RSIs (Cremers, Nair, and Peyer, 2008; Johnson, Karpoff, and Yi, 2015; Dey and White, 2021). To test this hypothesis, we estimate Equation (1) separately for the subsamples of firms with high and low RSI levels. High (low) RSI is defined as a firm's RSI in the top (bottom) half of its empirical distribution. We use two proxies for RSI motivated by the literature: (1) the R&D intensity of the customer's industry (Cust R&D)

and that of the supplier's industry (Supp R&D) that are measured as the weighted-average R&D intensity across all industries (Ahern and Harford, 2014), and (2) the core customers who represent at least 10% of the total revenue of the firm or if the sales to the customer are material to the business of the firm (Cen, Dasgupta, and Sen, 2016; Phua, Tham, and Wei, 2018).[†††]

[Insert Table 13 here]

Panels A and B of Table 13 provide the regression results for the RSIs as measured by the R&D intensity of the customer's industry and the core customers, respectively. The even- (odd-) numbered columns show the results for the association between the state FCAs and ATPs for high (low) RSI. For brevity, the results for the control variables are not reported. In Panel A, the R&D intensity of the customer's (supplier's) industry is the weighted average of R&D intensity across a firm's customer's (supplier's) industries. The R&D intensity of an industry is defined as the mean ratio of R&D to either total assets or sales across firms in a 3-digit SIC industry. The results indicate that the association between the state FCAs and ATPs does not vary systematically with the R&D intensity of the customer's industry. In Panel B, the RSI is represented by whether a firm has a core customer. The results are qualitatively similar to those reported in Panel A, which indicates that the effect of state FCAs on corporate ATPs is not sensitive to whether or not firms have core customers. Overall, our results show no evidence to support the bonding hypothesis.

[†††] Firm's customer and supplier industries are obtained from the Bureau of Economic Analysis's Input-Output tables. Principal customers are identified using Compustat's segment customer files. In the customer industry, weights are calculated based on the proportion of sales to each industry relative to total sales across all industries. In the supplier industry, weights are determined by the ratio of purchases from each industry to total purchases.

6.2. Firm valuation and performance

The managerial entrenchment hypothesis predicts that ATPs serve as tools for managers to entrench themselves in their positions, safeguarding their jobs and personal benefits at the expense of shareholder value and corporate efficiency. By contrast, the bonding hypothesis posits that ATPs Preserve value by limiting the potential for opportunistic whistleblowing to disrupt a firm's long-term operational strategy and relationship-specific investments. To test the two competing hypotheses, we deploy a model that accounts for firms that increase ATPs after exposure to state FCA laws. Specifically, we use a stacked DiD approach with an event window of five years before and after the passing of the state FCA laws. The sample contains the treated firms (firms exposed to state FCAs) and the control firms from the propensity score matching (PSM). We match the treated firms with the control firms using one-to-one nearest neighbor propensity score matching without replacement. The propensity scores are estimated from a logit model that regresses a treated firm indicator on the control variables specified in Equation (1). Then, we estimate the following regression model:

$$Firm_{i,t} = \alpha_0 + \alpha_1 FCA\_G_{i,t} + \alpha_2 Pst\_Eindex_{i,t} + \alpha_3 FCA\_G_{i,t} \times Pst\_Eindex_{i,t} + \beta' Controls_{i,t-1} + \gamma_i + \omega_t + \varepsilon_{i,t}, \tag{7}$$

where $Firm_{i,t}$ is measured by either Tobin's Q or operating performance for firm *i* in event year *t*. Following the literature, we use ROA, net margin, and gross margin to measure operating performance. Pst_Eindex is a binary variable that equals one if a firm experiences an increase in the *Eindex* after the passage of state FCA law. The other variables are defined in Equation (1). The γ and ω denote firm and event year fixed effects, respectively, and ε is the error term.

[Insert Table 14 here]

In Table 14, the coefficient for the interaction *FCA_G×Pst_Eindex* captures the channel through which the state FCAs affect firm valuation and performance. The coefficient estimates for the interaction are negative and statistically significant across the models, indicating that the state FCAs may lead to a lower firm's valuation and performance because the firm strategically adopts ATPs. This result is consistent with the managerial entrenchment hypothesis.

6.3. Disclosure quality

Under the managerial entrenchment hypothesis, ATPs allow managers to secure their positions and reduce the threat of external oversight or takeover. Without the pressure of potential takeover threats, entrenched managers might opt for less transparency in an attempt to withhold suboptimal decision-making (Leuz, Nanda, and Wysocki, 2003; Armstrong, Balakrishnan, and Cohen, 2012; Hermalin and Weisbach, 2012; Lin, Wei, and Xie, 2020). On the contrary, the bonding hypothesis posits that ATPs allow managers to undertake investments and strategies with longer horizons, potentially leading to more substantial and sustained firm growth. To communicate these long-term plans and their expected benefits effectively, firms may improve the quality of their disclosures, offering more comprehensive, forward-looking information that provides insight into the company's long-term strategic direction (Youmans and Tomlinson, 2018; Kotsantonis, Rehnberg, Serafeim, Ward, and Tomlinson, 2019).

To test the previously mentioned hypotheses, we use a stacked DiD approach as the specification in Equation (6) by replacing firm performance with disclosure quality. Then, we estimate the following regression model:

$$Disclosure\ Quality_{i,t} = \alpha_0 + \alpha_1 FCA\_G_{i,t} + \alpha_2 Pst\_Eindex_{i,t} + \alpha_3 FCA\_G_{i,t} \times Pst\_Eindex_{i,t} + \beta' Controls_{i,t-1} + \gamma_i + \omega_t + \varepsilon_{i,t}, \quad (8)$$

where $Disclosure\ Quality_{i,t}$ is defined by Chen, Miao, and Shevlin (2015) as the disaggregation quality for firm *i* in year *t*. This measure captures the level of detail in a firm's annual reports by counting nonmissing Compustat line items from both the balance sheet and income statement. For the balance sheet, the disclosure quality score is calculated as the value-weighted average ratio of nonmissing to total items across 11 account groups, where weights are determined by the proportion of each group's assets to the firm's total assets. For the income statement, the score is the equal-weighted ratio of nonmissing items across 7 account groups.[‡‡‡] The overall disclosure quality score combines these two scores, taking their simple average to represent the disclosure quality of both the balance sheet and the income statement. The other variables are defined in Equations (1) and (6). The γ and ω denote firm and event year fixed effects, respectively, and ε is the error term.

[Insert Table 15 here]

Table 15 shows that the coefficients on $FCA\_G \times Pst\_Eindex$ are negative and statistically significant at the 10% level. This indicates that firms adopting more antitakeover provisions due to increased whistleblowing risk lower their disclosure quality. This result provides supporting evidence for the managerial entrenchment hypothesis.

Overall, the findings from the tests conducted in this section offer corroborative evidence for the managerial entrenchment hypothesis, suggesting that firms adopting ATPs following the enactment of state FCAs are likely driven by motives of managerial entrenchment.

[‡‡‡] The detailed classification of account groups can be referred to Chen, Miao, and Shevlin's (2015).

7. Robustness Checks

7.1. Miscellaneous tests

One might argue that adopting antitakeover provisions in response to the higher whistleblowing threat from FCA laws based on the managerial hypothesis would only make sense if it is the same (presumably) entrenched manager who initiates such action. Therefore, we rerun our baseline models using a limited sample of firms with the same CEOs during the event window of FCA exposure. In Table IA1 of the Internet Appendix, the results are qualitatively similar to those reported in the previous tables.

Another potential concern is that the whistleblower provisions adopted under the Dodd-Frank Reform Act of 2010 may confound the interpretation of our results. To ensure that the state FCA effect is independent of the Dodd-Frank Act, we limit our sample to the period ending in 2009 and report the results in Table IA2 of the Internet Appendix. Overall, the results are qualitatively similar to those reported in the previous tables.

7.2. Alternative governance index

For robustness, we estimate the regressions as specified in the previous equations by replacing *Eindex* with the governance index proposed by Gompers, Ishii, and Metrick (2003), composed of 24 antitakeover provisions. The number of observations for the analysis become remarkably reduced as the index's data and its components from the ISS governance database are only available up to 2006. Nevertheless, we yield the same conclusions for a more extensive set of antitakeover provisions. The coefficients for *FCA_G* are positively and statistically significant at the 1% level. The results are available upon request.

7.3. Home bias in the investment of state pension fund

The literature raises a concern that states' adoptions of FCAs may be endogenous. State pension funds are more likely to invest in within-state firms (Brown, Pollet, and Weisbenner, 2015), representing a home bias. If such a bias is severe, then the state pension fund's investment may be a proxy for the location of a firm's headquarters instead of the effect of the state's FCA. To address this concern, we re-estimate our main regressions using only firm-years in which states with pension fund investment in the firm differ from the state of the firm's headquarters and find the results remain unchanged.

8. Conclusion

While numerous studies have explored how whistleblower laws aid in uncovering accounting fraud, the potential drawbacks of these regulations have received less attention. Our research focuses on whether the threat of whistleblowing prompts managers to implement antitakeover provisions, thus strengthening their control over the company. We utilize the staggered introduction of FCAs across U.S. states as a quasi-natural experiment to examine this question. Our difference-in-differences analysis reveals a notable increase in the entrenchment index for firms receiving investments from state pension funds in jurisdictions with enacted FCA legislation. This study dissects the entrenchment index further, highlighting the adoption of staggered boards, poison pills, and restrictions on charter amendments as the most favored antitakeover measures following the introduction of FCAs.

To solidify the association between the risk of whistleblowing and the adoption of antitakeover provisions, we conduct several robustness checks. These include analyzing samples excluding variations from pension fund investments, performing parallel trend and falsification tests, examining heterogeneous treatment effects and spillover effects, and adjusting for the home biases of investment. Our findings consistently underscore the robustness of our initial results.

Further analysis strengthens the link between FCAs and antitakeover provisions among firms that are more aware of whistleblowing risk, more susceptible to whistleblowing, and where managers derive significant private benefits. This supports the notion of managerial entrenchment. We also rule out the alternative explanation that antitakeover measures are adopted for bonding purposes through two additional tests. These examine the impact of relationship-specific investments and the effect of increased antitakeover measures, post-FCA enactment, on firm valuation and performance. Notably, we observe a more pronounced negative correlation between FCAs, firm valuation, and performance among firms that heighten their use of antitakeover provisions, reinforcing the premise that such actions are driven more by a desire for managerial entrenchment than by efforts to bond with stakeholders.

Appendix A: Definitions of Variables

This table shows the definitions of the main variables used in this study.

Variable	Definition	Data source
	Corporate governance	
Eindex	The entrenchment index created by Bebchuk, Cohen, and Ferrell (2009) comprises six entrenchment provisions: staggered boards, poison pills, golden parachutes, supermajority requirement for charter amendments, supermajority requirement for bylaw amendments, and supermajority requirement for mergers.	ISS Corporate Governance database
Staggered Board	Dummy variable equals one if the firm has a staggered/classified board structure in which directors are divided into separate classes elected to overlapping terms and zero otherwise.	ISS Corporate Governance database
Golden Parachute	Dummy variable equals one if the firm has a provision that provides benefits to managers/board members in the firing/demotion following a change in control event and zero otherwise.	ISS Corporate Governance database
Poison Pill	Dummy variable that equals one if the firm has a non-shareholder-approved poison pill provision and zero otherwise.	ISS Corporate Governance database
Limit to Change Bylaws	Dummy variable equals one if the firm has a provision limiting the shareholder's ability to amend the firm bylaws without a majority vote and zero otherwise.	ISS Corporate Governance database
Limit to Change Charter	Dummy variable equals one if the firm has a provision limiting the shareholder's ability to amend the firm charter without a majority vote and zero otherwise.	ISS Corporate Governance database
Supermajority	Dummy variable equals one if the firm has a supermajority vote requirement to approve a merger and zero otherwise.	ISS Corporate Governance database
	Main independent variables	
FCA_G	Dummy variable equals one if at least one state pension fund is located in a state with a general FCA invested in the firm and zero otherwise.	13F filings by pension funds
SPF	Dummy variable equals one if at least one state pension fund in the lagged year invested in the firm and zero otherwise.	13F filings by pension funds
	Firm characteristics	
Firm Age	Natural logarithm of firm age.	COMPUSTAT
Size	Natural logarithm of firm's total assets.	COMPUSTAT
Cash ratio	The ratio of cash to total assets.	COMPUSTAT
R&D	The ratio of research and development investment to total assets.	COMPUSTAT

ROA	Returns on assets ratio.	COMPUSTAT
Capex	Capital expenditures to total assets ratio.	COMPUSTAT
Leverage	Total debt to total assets ratio.	COMPUSTAT
Inst Own	Total institutional shares ownership in a firm.	COMPUSTAT
	Board characteristics	
Board size	Natural logarithm of the number of directors.	EXECUCOMP
Female dir ratio	The ratio of female directors to total director number.	EXECUCOMP
Indep dir ratio	The ratio of independent directors to the total directors.	EXECUCOMP
	CEO characteristics	
CEO tenure	CEO tenure in years.	EXECUCOMP
CEO duality	Dummy variable equals one if the CEO serves as both CEO and board chair and zero otherwise.	EXECUCOMP
Female CEO	Dummy variable equals one if the CEO is female and zero otherwise.	EXECUCOMP

Appendix B: State False Claims Acts (FCA)

This table summarizes the state-by-state FCA provisions as of 2015. The column "FCA type" indicates the coverage of state FCAs: either general fraud (General) or Medicaid fraud (Medicaid). The states not listed in the table are states that did not adopt their own FCA. States with * (reported in the column "Use") are the FCA states used in the empirical tests.

Use	State	Adoption year	FCA type	Qui tam provision
	California	1987	General	Yes
*	Illinois	1992	General	Yes
*	Florida	1994	General	Yes
	Texas	1995	Medicaid	Yes
	Nebraska	1996	Medicaid	No
	Louisiana	1997	Medicaid	Yes
*	D.C.	1998	General	Yes
*	Nevada	1999	General	Yes
*	Delaware	2000	General	Yes
*	Massachusetts	2000	General	Yes
*	Hawaii	2001	General	Yes
*	Tennessee	2001	General	Yes
*	Virginia	2003	General	Yes
*	New Mexico	2004	General	Yes
*	Indiana	2005	General	Yes
*	Montana	2005	General	Yes
	New Hampshire	2005	Medicaid	Yes
	Georgia	2007	Medicaid	Yes
	Missouri	2007	Medicaid	No
*	New York	2007	General	Yes
*	Oklahoma	2007	General	Yes
	Michigan	2008	Medicaid	Yes
*	New Jersey	2008	General	Yes
*	Rhode Island	2008	General	Yes
	Wisconsin	2008	Medicaid	Yes
	Arizona	2009	Medicaid	No
	Arkansas	2009	Medicaid	No
*	North Carolina	2009	General	Yes
	Colorado	2010	Medicaid	Yes
	Utah	2011	General	No
*	Iowa	2010	General	Yes
*	Minesota	2010	General	Yes
	Washington	2012	Medicaid	Yes
	Maryland	2015	Medicaid	Yes
*	Vermont	2015	General	Yes

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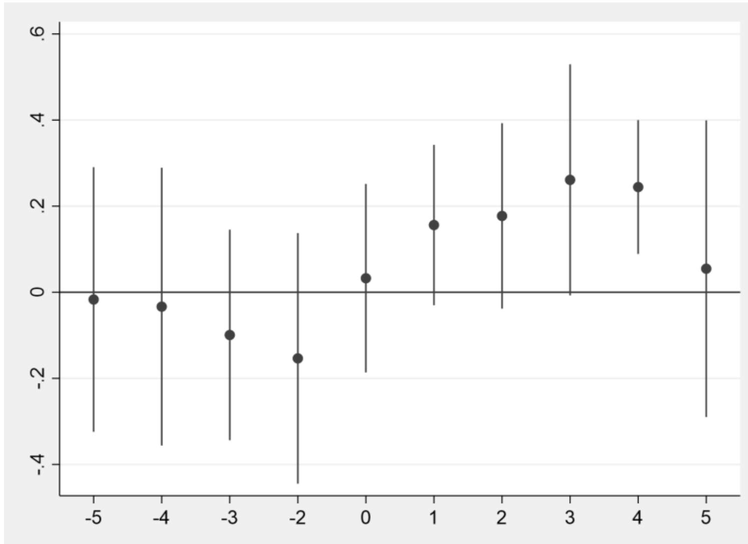
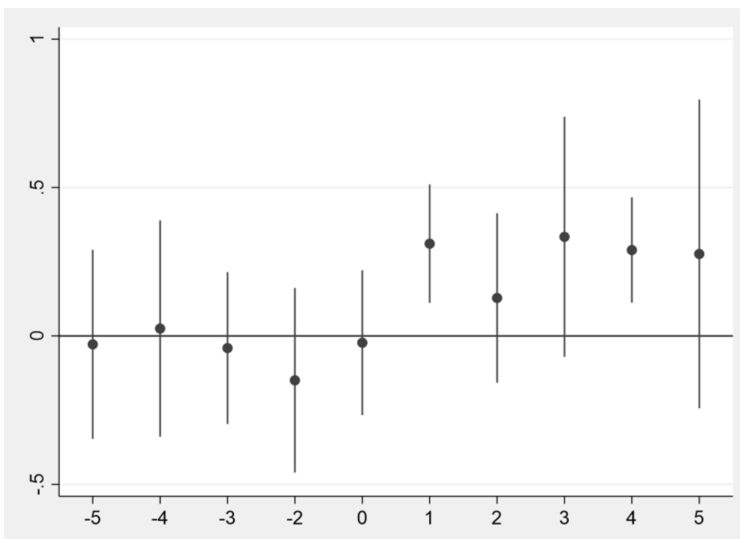


Panel A: Full sample



Panel B: SPF = 1 only

Figure 1. Trends of ATPs in the Treatment Sample (Net of Control)

This figure shows the trend of the E-index for the treatment sample net of the control group five years before and after the event (the passage of state FCA laws) year. Panel A shows the results using the full sample, and Panel B shows the results using the sample conditional on firms' shares held by any state pension fund.

Table 1. Descriptive Statistics

This table presents the number of observations (*N*), means (Mean), 25th percentiles (P25), medians (Median), 75th percentiles (P75), and standard deviations (STD). All variables are lagged by one year relative to the antitakeover index. Variables are defined in Appendix A and winsorized at the 1st and 99th percentiles.

Variable	*N*	Mean	P25	Median	P75	STD
Eindex	23,972	2.976	2.000	3.000	4.000	1.587
Gindex	9,559	9.012	7.000	9.000	11.000	2.717
FCA_G	22,865	0.843	1.000	1.000	1.000	0.364
SPF	22,865	0.976	1.000	1.000	1.000	0.152
Firm Age	23,971	3.097	2.639	3.136	3.689	0.694
Size	23,926	7.769	6.534	7.633	8.863	1.727
Cash	23,417	0.095	0.017	0.054	0.131	0.115
R&D	23,925	0.028	0.000	0.000	0.024	0.069
ROA	23,898	0.034	0.010	0.042	0.087	0.121
Board size	18,202	2.205	2.079	2.197	2.398	0.264
Female director	18,202	0.124	0.000	0.111	0.182	0.103
Indep director	18,202	0.750	0.667	0.778	0.875	0.147
CEO tenure	19,500	8.713	3.000	6.000	12.000	8.195
CEO duality	19,500	0.773	1.000	1.000	1.000	0.419
Female CEO	23,972	0.022	0.000	0.000	0.000	0.146

Table 2. Pearson/Spearman Correlation Matrix

This table depicts the Pearson/Spearman correlation matrix. Pearson correlations are below the diagonal; Spearman correlations are above the diagonal. All variables are defined in Appendix A and winsorized at the 1st and 99th percentiles. Statistical significance at the 10%, 5%, and 1% levels are denoted by *, **, and ***, respectively.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)	(15)
(1) *Eindex*	–	0.69***	0.03**	0.03*	0.06***	0.03*	-0.08***	-0.05***	-0.07***	0.14***	0.09***	0.15***	-0.11***	-0.01	-0.01
(2) *Gindex*	0.68***	–	0.10***	0.04**	0.32***	0.21***	-0.12***	-0.07***	-0.06***	0.31***	0.18***	0.22***	-0.16***	-0.02	-0.02
(3) *FCA_G*	0.03**	0.10***	–	0.32***	0.09***	0.31***	0.05***	0.01	0.09***	0.13***	0.13***	0.14***	-0.01	0.00	-0.02
(4) *SPF*	0.03*	0.04**	0.32***	–	0.01	0.01	-0.01	0.02	0.03**	0.03*	0.01	0.07***	0.00	0.01	-0.00
(5) *Firm age*	0.05***	0.32***	0.09***	0.01	–	0.37***	-0.15***	-0.09***	0.00	0.42***	0.24***	0.23***	-0.10***	0.00	-0.00
(6) *Size*	0.00	0.17***	0.29***	0.01	0.33***	–	-0.29***	-0.27***	-0.10***	0.60***	0.32***	0.18***	-0.06***	0.01	-0.06***
(7) *Cash*	-0.09***	-0.12***	0.03*	0.01	-0.18***	-0.31***	–	0.42***	0.18***	-0.30***	-0.05***	-0.02	0.04***	-0.03*	0.04***
(8) *R&D*	-0.09***	-0.14***	-0.01	0.02	-0.20***	-0.27***	0.42***	–	0.04***	-0.27***	-0.13***	0.09***	-0.04***	-0.01	-0.04**
(9) *ROA*	-0.02	0.01	0.10***	0.07***	0.07***	0.05***	0.01	-0.28***	–	-0.08***	0.03**	-0.05***	0.09***	0.04***	-0.01
(10) *Board size*	0.14***	0.29***	0.13***	0.03*	0.40***	0.60***	-0.32***	-0.29***	0.04***	–	0.31***	0.13***	-0.09***	0.01	-0.06***
(11) *Female director*	0.08***	0.16***	0.12***	0.02	0.22***	0.31***	-0.10***	-0.13***	0.07***	0.29***	–	0.23***	-0.12***	-0.01	0.12***
(12) *Indep director*	0.15***	0.23***	0.15***	0.08***	0.20***	0.17***	-0.03*	0.03*	-0.01	0.12***	0.21***	–	-0.17***	-0.02	-0.01
(13) *CEO tenure*	-0.11***	-0.16***	-0.02	-0.01	-0.06***	-0.09***	0.07***	-0.02	0.07***	-0.10***	-0.11***	-0.22***	–	0.02	-0.02
(14) *CEO duality*	-0.01	-0.02	0.00	0.01	-0.00	0.00	-0.03*	-0.01	0.03**	0.01	-0.01	-0.03*	0.02	–	0.01
(15) *Female CEO*	-0.01	-0.02	-0.02	-0.00	0.00	-0.06***	0.05***	-0.02	-0.01	-0.06***	0.16***	-0.00	-0.02	0.01	–

Table 3. Effect of FCA Laws on Antitakeover Provisions

This table shows the results of the following regression model:

$$Eindex_{i,t} = \alpha_0 + \alpha_1 FCA\_G_{i,t} + \alpha_2 SPF_{i,t-1} + \beta' Controls_{i,t-1} + \gamma_i + \pi_t + \varepsilon_{i,t},$$

where $Eindex_{i,t}$ is the entrenchment index proposed by Bebchuk, Cohen, and Ferrell (2009) of firm i in year t; *FCA_G* is a binary variable that equals one if a firm is exposed to a general FCA and zero otherwise. *SPF* is a binary variable that equals one if a firm's shares were owned by at least one state pension fund in the lagged year. *Control* is the vector of control variables; γ and π denote firm and year fixed effects, respectively, and ε is the error term. All variables are defined in Appendix A. The *t*-statistics with clustered robust standard errors at the firm level (Petersen, 2009) are in parentheses. Statistical significance at the 10%, 5%, and 1% levels are denoted by *, **, and ***, respectively.

	No Control		Control Included	*SPF* = 1 Only	Matched Sample
	(1)	(2)	(3)	(4)	(5)
FCA_G	0.861***	0.100***	0.090**	0.094**	0.150***
	(19.05)	(3.30)	(2.35)	(2.34)	(3.22)
SPF	-0.209**	0.075*	0.061		0.072
	(-2.24)	(1.72)	(1.19)		(0.83)
Firm age			0.470***	0.485***	0.587***
			(5.12)	(5.27)	(5.33)
Size			0.045	0.037	-0.008
			(1.37)	(1.11)	(-0.20)
Cash			0.063	0.060	-0.131
			(0.53)	(0.50)	(-0.78)
R&D			0.794**	0.759**	0.809*
			(2.29)	(2.19)	(1.81)
ROA			-0.068	-0.068	-0.080
			(-0.73)	(-0.71)	(-0.64)
Board size			0.184***	0.189***	0.260***
			(2.78)	(2.83)	(2.82)
Female director			0.043	0.061	0.245
			(0.26)	(0.37)	(1.07)
Indep director			-0.024	-0.021	0.000
			(-0.25)	(-0.21)	(0.00)
CEO tenure			-0.004**	-0.004**	-0.006***
			(-2.09)	(-2.14)	(-2.67)
CEO duality			-0.017	-0.011	0.005
			(-0.38)	(-0.25)	(0.07)
Female CEO			0.059	0.035	-0.025
			(0.67)	(0.40)	(-0.18)
Constant	2.505***	1.349***	-0.604*	-0.536	-0.743*
	(23.37)	(25.98)	(-1.80)	(-1.62)	(-1.93)
Firm fixed effect	NO	YES	YES	YES	YES
Year fixed effect	NO	YES	YES	YES	YES
N	21,736	21,736	15,535	15,191	6,838
Adj. R^2	0.039	0.719	0.709	0.710	0.719

Table 4. Decomposition of Antitakeover Provisions

This table shows the results of the following probit model:

$$\Pr(Eindex\ component_{i,t} = 1) = \Phi(\alpha_0 + \alpha_1 FCA_{G_{i,t}} + \alpha_2 SPF_{i,t-1} + \beta' Controls_{i,t-1} + \gamma_i + \pi_t),$$

where *Eindex component* is a given component of the entrenchment index proposed by Bebchuk, Cohen, and Ferrell (2009) of firm *i* in year *t*. These components include six binary variables: *Staggered Board*, *Golden Parachute*, *Poison Pill*, *Limit to Bylaw Amendments, Limit to Change Amendments,* and *Supermajority*. Φ is the cumulative distribution function of the standard normal distribution. *FCA_G* is a binary variable that equals one if a firm is exposed to a general FCA and zero otherwise. *SPF* is a binary variable that equals one if a firm's shares were owned by at least one state pension fund in the lagged year. *Control* is the vector of control variables; γ_i and π_t denote firm and year fixed effects, respectively, and ε is the error term. All variables are defined in Appendix A. The *t*-statistics with clustered robust standard errors at the firm level (Petersen, 2009) are in parentheses. Statistical significance at the 10%, 5%, and 1% levels are denoted by *, **, and ***, respectively.

	Staggered Board	*Golden Parachute*	*Poison Pill*	*Limit to Bylaw Amendments*	*Limit to Charter Amendments*	*Supermajority*
	(1)	(2)	(3)	(4)	(5)	(6)
FCA_G	0.671***	-0.140	0.573***	-0.083	0.646**	-0.044
	(2.77)	(-0.63)	(2.79)	(-0.33)	(2.33)	(-0.16)
SPF	-0.155	0.027	-0.029	-0.011	0.340	-0.564**
	(-0.42)	(0.11)	(-0.08)	(-0.05)	(0.72)	(-2.29)
Firm age	-0.796	1.910***	1.127**	1.304	1.619	2.741***
	(-0.89)	(3.16)	(2.19)	(1.58)	(1.52)	(3.14)
Size	0.009	0.225	-0.109	-0.103	0.319	0.249
	(0.03)	(1.25)	(-0.61)	(-0.49)	(1.20)	(0.91)
Cash	0.562	-0.898	-1.054*	-0.328	0.082	0.837
	(0.58)	(-1.43)	(-1.74)	(-0.53)	(0.09)	(0.89)
R&D	4.122	2.771	-2.138	1.654	5.851**	3.333**
	(1.01)	(0.90)	(-1.49)	(1.27)	(2.14)	(1.96)
ROA	-0.989	0.632	0.254	-0.092	-1.056	0.812
	(-1.17)	(1.16)	(0.56)	(-0.21)	(-1.33)	(1.43)
Board size	1.555***	0.250	0.020	0.431	0.391	0.407
	(4.16)	(0.71)	(0.05)	(1.05)	(0.73)	(1.03)
Female director	-1.085	-0.471	-0.436	-1.361	-2.790*	1.176
	(-1.06)	(-0.62)	(-0.49)	(-1.31)	(-1.79)	(1.02)
Indep director	-0.640	0.127	1.777***	0.499	1.208	-0.318
	(-0.93)	(0.21)	(3.91)	(0.88)	(1.09)	(-0.51)
CEO tenure	-0.019	-0.018**	-0.008	-0.011	-0.023	0.007
	(-1.16)	(-2.18)	(-0.74)	(-1.10)	(-1.17)	(0.64)

CEO duality	0.092	0.196	-0.122	-0.295	0.204	-0.132
	(0.25)	(1.00)	(-0.58)	(-1.44)	(0.64)	(-0.65)
Female CEO	0.562	0.365	-0.249	0.371	-0.059	0.098
	(0.97)	(1.13)	(-0.50)	(0.86)	(-0.11)	(0.30)
Constant	3.083	-11.103***	0.866	-7.622***	-12.329***	-10.919***
	(1.01)	(-3.59)	(0.58)	(-3.26)	(-4.64)	(-4.84)
Marginal effect of *FCA_G*	0.071***	-0.006	0.031***	-0.001	0.084**	-0.004
Firm fixed effect	YES	YES	YES	YES	YES	YES
Year fixed effect	YES	YES	YES	YES	YES	YES
N	15,535	15,535	15,535	15,535	15,535	15,535
Pseudo R^2	0.644	0.553	0.578	0.761	0.902	0.765

Table 5. Parallel Trend Assumption: Dynamic Regression

This table shows the results of the dynamic regression model of whistleblower law exposure and ATP adoption as suggested by Sun and Abraham (2021) as follows:

$$Eindex_{i,t} = \sum_{\substack{j=-5 \\ j\neq -1}}^{j=+5} \alpha_j FCA\_G_j + \theta SPF_{i,t-1} + \beta' Controls_{i,t-1} + \gamma_i + \pi_t + \varepsilon_{i,t},$$

where *Eindex* is the entrenchment index constructed by Bebchuk, Cohen, and Ferrell (2009) of firm *i* in year *t*. FCA_G_j is an interaction variable between FCA_G and time dummy indicating *j*-year before or after the passage of state FCA law. *SPF* is a binary variable that equals one if at least one state pension fund invests in the firm in the lagged year. *Controls* are the vector of control variables; γ_i and π_t denote firm and event year fixed effects, respectively; and $\varepsilon_{i,t}$ is the error term.

	Event window [-5, +5]	
Dependent variable:	Full sample	*SPF* = 1 only
E-index	(1)	(2)
FCA_G_{-5}	-0.017	-0.028
	(-0.11)	(-0.17)
FCA_G_{-4}	-0.033	0.025
	(-0.20)	(0.13)
FCA_G_{-3}	-0.099	-0.041
	(-0.80)	(-0.31)
FCA_G_{-2}	-0.153	-0.149
	(-1.03)	(-0.94)
FCA_G_{0}	0.033	-0.023
	(0.29)	(-0.18)
FCA_G_{1}	0.156*	0.311***
	(1.65)	(3.05)
FCA_G_{2}	0.177	0.128
	(1.61)	(0.88)
FCA_G_{3}	0.261*	0.334
	(1.91)	(1.62)
FCA_G_{4}	0.244***	0.289***
	(3.09)	(3.20)
FCA_G_{5}	0.055	0.276
	(0.31)	(1.04)
SPF	0.090	
	(1.06)	
Firm age	2.129***	2.132***
	(19.60)	(19.44)
Size	0.417***	0.409***
	(8.09)	(7.82)
Cash	0.823***	0.818***
	(4.28)	(4.23)
R&D	3.014***	2.967***
	(4.31)	(4.24)
ROA	0.068	0.041
	(0.46)	(0.27)
Board size	-0.135	-0.138
	(-1.28)	(-1.30)
Female director	0.589**	0.623**
	(2.32)	(2.45)
Indep director	1.510***	1.493***
	(9.86)	(9.66)
CEO tenure	-0.007**	-0.007**
	(-2.41)	(-2.29)
CEO duality	-0.197***	-0.192***
	(-2.83)	(-2.66)
Female CEO	-0.030	-0.053
	(-0.19)	(-0.34)

Constant	-7.889***	-7.708***
	(-19.15)	(-19.05)
Test coefficients FCA_G$_j$ (j < 0): F-stats	0.40	0.26
(p-value)	(0.807)	(0.902)
Test coefficients FCA_G$_j$ ($j \geq 0$): F-stats	2.26**	2.92***
(p-value)	(0.035)	(0.008)
Firm fixed effect	YES	YES
Event Year fixed effect	YES	YES
N	11,453	11,218
Adj. R^2	0.704	0.704

Table 6. Falsification Test

This table presents the regression results. The dependent variable is the *Eindex* proposed by Bebchuk, Cohen, and Ferrell (2009). *FCA_G* is a binary variable that equals one if a firm is exposed to a general FCA and zero otherwise. *FCA_M* is a binary variable that equals one if a firm is exposed to a Medicaid-only FCA and zero otherwise. *SPF* is a binary variable that equals one if a firm's shares were owned by at least one state pension fund in the lagged year. All variables are defined in Appendix A. The *t*-statistics with clustered robust standard errors at the firm level (Petersen, 2009) are in parentheses. Statistical significance at the 10%, 5%, and 1% levels are denoted by *, **, and ***, respectively.

	Table 3 Column (3)	Sub-sample	
	(1)	(2)	(3)
FCA_G	0.090**	0.185**	
	(2.35)	(2.02)	
FCA_M			-0.208
			(-1.25)
SPF	0.061	0.129	0.162
	(1.19)	(1.12)	(1.33)
Firm age	0.470***	0.604***	0.555***
	(5.12)	(2.95)	(2.64)
Size	0.045	0.070	0.072
	(1.37)	(0.88)	(0.91)
Cash	0.063	0.040	0.026
	(0.53)	(0.12)	(0.08)
R&D	0.794**	2.026**	2.061**
	(2.29)	(2.10)	(2.18)
ROA	-0.068	-0.160	-0.144
	(-0.73)	(-0.66)	(-0.60)
Board size	0.184***	0.155	0.199
	(2.78)	(1.03)	(1.28)
Female director	0.0430	-0.221	-0.166
	(0.26)	(-0.66)	(-0.52)
Indep director	-0.025	0.117	0.138
	(-0.25)	(0.48)	(0.59)
CEO tenure	-0.004**	-0.003	-0.003
	(-2.09)	(-0.55)	(-0.52)
CEO duality	-0.017	-0.184*	-0.153
	(-0.38)	(-1.68)	(-1.46)
Female CEO	0.059	0.417	0.418*
	(0.67)	(1.60)	(1.70)
Constant	-0.604*	-1.299*	-1.179
	(-1.80)	(-1.73)	(-1.60)
Chi-square test for equality of coefficients		3.49*	
(*p*-value)		(0.062)	
Firm fixed effect	YES	YES	YES
Year fixed effect	YES	YES	YES
N	15,535	2,263	2,263
Adj. R^2	0.709	0.737	0.738

Table 7. Heterogenous Treatment Effect: Limited Sample

This table shows the results of regression models as specified in Equation (1) using a limited sample after removing firms that are always treated during the sample period. The dependent variable is the *Eindex* proposed by Bebchuk, Cohen, and Ferrell (2009). *FCA_G* is a binary variable that equals one if a firm is exposed to a general FCA and zero otherwise. *SPF* is a binary variable that equals one if a firm's shares are owned by at least one state pension fund in the lagged year. All variables are defined in Appendix A. The *t*-statistics with clustered robust standard errors at the firm level (Petersen, 2009) are in parentheses. Statistical significance at the 10%, 5%, and 1% levels are denoted by *, **, and ***, respectively.

	No Control		Control included	*SPF* = 1 only
	(1)	(2)	(3)	(4)
FCA_G	0.862***	0.110***	0.092**	0.092**
	(17.85)	(3.34)	(2.23)	(2.13)
SPF	-0.168*	0.067	0.053	
	(-1.75)	(1.44)	(0.96)	
Firm age			0.442***	0.457***
			(4.42)	(4.54)
Size			0.058	0.049
			(1.52)	(1.29)
Cash			0.053	0.047
			(0.36)	(0.32)
R&D			0.752*	0.720
			(1.69)	(1.61)
ROA			-0.061	-0.065
			(-0.54)	(-0.57)
Board size			0.201***	0.208***
			(2.67)	(2.72)
Female director			-0.061	-0.040
			(-0.34)	(-0.22)
Indep director			0.019	0.028
			(0.17)	(0.25)
CEO tenure			-0.004*	-0.004*
			(-1.88)	(-1.89)
CEO duality			-0.026	-0.024
			(-0.53)	(-0.47)
Female CEO			0.106	0.078
			(1.01)	(0.73)
Constant	2.487***	1.391***	-0.636*	-0.571
	(22.39)	(25.09)	(-1.69)	(-1.54)
Firm fixed effect	NO	YES	YES	YES
Year fixed effect	NO	YES	YES	YES
N	18,494	18,494	13,221	12,912
Adj. R^2	0.041	0.709	0.698	0.698

Table 8. Heterogenous Treatment Effect: Stacked DiD Approach

This table presents the results of the stacked DiD approach related to the *Eindex* proposed by Bebchuk, Cohen, and Ferrell (2009). We first stack cohorts of treatment and control firms in an event study style. In each general FCA event cohort, treatment firms are firms that were exposed to a general FCA for the first time, and clean control firms are firms that have not experienced a general FCA during the whole sample period. We perform propensity score matching (PSM) to match the control and treatment firms using data from the year immediately before the treatment events. We use the nearest propensity scores and a caliper of 0.01 to select control firms in each cohort comparable to the treatment firms by using the control variables used in Equation (1). The control firms are matched with replacement. *FCA_G* is a binary variable that equals one if a firm is exposed to a general FCA and zero otherwise. *SPF* is a binary variable that equals one if a firm's shares were owned by at least one state pension fund in the lagged year. All variables are defined in Appendix A. The *t*-statistics with clustered robust standard errors at the firm level (Petersen, 2009) are in parentheses. Statistical significance at the 10%, 5%, and 1% levels are denoted by *, **, and ***, respectively.

	Event window [-5, +5]			
	No Control		Control included	*SPF* = 1 only
	(1)	(2)	(3)	(4)
FCA_G	1.001***	0.093***	0.088**	0.096**
	(25.90)	(2.91)	(2.09)	(2.16)
SPF	-0.430***	0.064	0.029	
	(-4.49)	(1.24)	(0.47)	
Firm age			0.433***	0.466***
			(4.11)	(4.43)
Size			0.042	0.034
			(1.07)	(0.87)
Cash			0.052	0.059
			(0.40)	(0.45)
R&D			0.728*	0.692*
			(1.91)	(1.82)
ROA			-0.074	-0.074
			(-0.74)	(-0.73)
Board size			0.164**	0.174**
			(2.18)	(2.29)
Female director			0.188	0.214
			(1.01)	(1.15)
Indep director			-0.062	-0.067
			(-0.56)	(-0.60)
CEO tenure			-0.005**	-0.005**
			(-2.21)	(-2.20)
CEO duality			0.018	0.033
			(0.38)	(0.66)
Female CEO			-0.084	-0.107
			(-0.84)	(-1.04)
Constant	2.348***	1.350***	-0.425	-0.466
	(22.75)	(23.10)	(-1.09)	(-1.21)
Firm fixed effect	NO	YES	YES	YES
Event year fixed effect	NO	YES	YES	YES
N	16,973	16,973	11,921	11,630
Adj. R^2	0.035	0.724	0.718	0.719

Table 9. Spillover Effects of the Treatment Group

This table reports the results of the spillover model as suggested by Berg, Reisinger, and Streitz (2021). The spillover model is estimated as follows:

$$Eindex_{i,g,t} = \alpha_0 + \alpha_1 FCA\_G_{i,g,t} + \alpha_T \overline{FCA\_G}_g \times FCA\_G_{i,g,t} + \alpha_C \overline{FCA\_G}_g \times \left(1 - FCA\_G_{i,g,t}\right) + \alpha_2 SPF_{i,t-1} + \beta' Controls_{i,t-1} + \gamma_i + \pi_t + \varepsilon_{i,g,t},$$

where $Eindex_{i,g,t}$ is the entrenchment index of firm i in the spillover group g at time t. $\overline{FCA\_G}_g$ is the average treatment effect of all other firms located in the same spillover group g, excluding the firm i itself. Other variables are defined in the previous equations. γ and π denote firm and year fixed effects, respectively, and ε is the error term.

	Full sample	*SPF* = 1 only
	(1)	(2)
FCA_G	0.087**	0.084**
	(2.28)	(2.17)
$\overline{FCA\_G} \times FCA\_G$	0.190*	0.203**
(Treatment spillover)	(1.92)	(2.03)
$\overline{FCA\_G} \times (1 - FCA\_G)$	0.344*	0.331*
(Control spillover)	(1.78)	(1.75)
SPF	0.058	
	(1.12)	
Firm age	0.470***	0.480***
	(5.12)	(5.24)
Size	0.046	0.037
	(1.39)	(1.15)
Cash	0.058	0.063
	(0.49)	(0.53)
R&D	0.796**	0.723**
	(2.31)	(2.11)
ROA	-0.071	-0.067
	(-0.76)	(-0.71)
Board size	0.179***	0.179***
	(2.72)	(2.72)
Female director	0.047	0.064
	(0.29)	(0.39)
Indep director	-0.028	-0.021
	(-0.28)	(-0.21)
CEO tenure	-0.004**	-0.004**
	(-2.06)	(-2.07)
CEO duality	-0.017	-0.010
	(-0.38)	(-0.23)
Female CEO	0.061	0.038
	(0.70)	(0.43)
Constant	-0.689**	-0.601*
	(-2.02)	(-1.82)
Firm fixed effect	YES	YES
Year fixed effect	YES	YES
N	15,535	15,191
Adj. R^2	0.710	0.711

Table 10. Cross-Sectional Variations: Managerial Awareness of Whistleblowing Threat

This table provides regression estimates as specified in Equation (1) for subsamples of low and high managerial awareness of whistleblower threat. We use state pension fund ownership to proxy for managerial awareness with the 1% cutoff point. High (Low) awareness is a firm's state pension fund ownership of at least (less than) 1%. All variables are defined in Appendix A. The *t*-statistics with clustered robust standard errors at the firm level (Petersen, 2009) are in parentheses. Statistical significance at the 10%, 5%, and 1% levels are denoted by *, **, and ***, respectively.

	Low Awareness	High Awareness
	(1)	(2)
FCA_G	0.046	0.107**
	(0.84)	(2.08)
SPF	-0.026	0.175**
	(-0.44)	(2.07)
Firm age	0.261*	0.561***
	(1.85)	(4.70)
Size	0.060	0.037
	(0.96)	(0.94)
Cash	-0.098	0.092
	(-0.41)	(0.68)
R&D	1.715**	0.552
	(2.26)	(1.38)
ROA	-0.057	-0.051
	(-0.33)	(-0.46)
Board size	0.136	0.195**
	(1.31)	(2.30)
Female director	0.015	0.056
	(0.06)	(0.27)
Indep director	-0.145	0.030
	(-0.84)	(0.25)
CEO tenure	0.001	-0.006**
	(0.22)	(-2.51)
CEO duality	0.008	-0.030
	(0.12)	(-0.53)
Female CEO	0.070	0.056
	(0.54)	(0.47)
Constant	0.058	-0.923**
	(0.10)	(-2.24)
Chi-square test for equality of coefficients	10.35***	
(*p*-value)	(0.001)	
Firm fixed effect	YES	YES
Year fixed effect	YES	YES
N	5,536	9,999
Adj. R^2	0.676	0.725

Table 11. Cross-Sectional Variations: Likelihood of Whistleblowing

This table provides regression estimates as specified in Equation (1) for subsamples of high and low likelihood of whistleblowing. We use insider whistleblower (represented by the unionization rate) and outsider whistleblower (represented by financial analyst coverage) to measure the likelihood of whistleblowing. The results of insider and outsider whistleblowers are reported in Panels A and B, respectively. All variables are defined in Appendix A. The *t*-statistics with clustered robust standard errors at the firm level (Petersen, 2009) are in parentheses. Statistical significance at the 10%, 5%, and 1% levels are denoted by *, **, and ***, respectively.

Panel A: Insider Whistleblower

	Unionization Rate	
	Low	High
	(1)	(2)
FCA_G	0.076	0.102**
	(1.25)	(2.00)
SPF	-0.047	0.068
	(-0.53)	(0.84)
Firm age	0.482***	0.439***
	(3.36)	(3.60)
Size	0.066	0.039
	(1.17)	(0.93)
Cash	0.068	0.007
	(0.36)	(0.04)
R&D	1.286**	0.505
	(2.34)	(1.07)
ROA	-0.076	-0.088
	(-0.48)	(-0.68)
Board size	0.269***	0.124
	(2.70)	(1.29)
Female director	0.305	-0.083
	(1.27)	(-0.34)
Indep director	-0.068	0.045
	(-0.44)	(0.33)
CEO tenure	-0.007**	-0.006**
	(-2.45)	(-2.19)
CEO duality	-0.016	-0.009
	(-0.24)	(-0.14)
Female CEO	-0.062	0.160
	(-0.49)	(1.29)
Constant	-0.853*	-0.301
	(-1.77)	(-0.68)
Chi-square test for equality of coefficients	3.49*	
(*p*-value)	(0.06)	
Firm fixed effect	YES	YES
Year fixed effect	YES	YES
N	6,188	7,384
Adj. R^2	0.718	0.726

Table 11 (Continued)

Panel B: Outsider Whistleblower

	Excess Analyst Coverage (Difference)		*Excess Analyst Coverage (Residuals)*	
	Low	High	Low	High
	(1)	(2)	(3)	(4)
FCA_G	0.019	0.106*	0.020	0.151***
	(0.36)	(1.85)	(0.43)	(2.70)
SPF	0.099	0.004	0.073	0.087
	(1.22)	(0.07)	(1.11)	(1.13)
Firm age	0.693***	0.290**	0.376***	0.509***
	(4.81)	(2.44)	(2.89)	(3.90)
Size	0.114**	0.044	0.071	0.045
	(2.19)	(0.98)	(1.49)	(0.97)
Cash	0.179	0.137	0.128	-0.137
	(1.13)	(0.84)	(0.73)	(-0.93)
R&D	1.072**	1.229***	1.074	0.402
	(2.16)	(2.70)	(1.34)	(1.14)
ROA	-0.404***	0.073	-0.158	-0.029
	(-2.98)	(0.61)	(-1.02)	(-0.26)
Board size	0.103	0.187**	0.184**	0.094
	(0.97)	(2.26)	(1.99)	(1.03)
Female director	-0.274	0.362*	0.008	0.249
	(-1.11)	(1.81)	(0.03)	(1.20)
Indep director	0.013	-0.015	-0.070	0.014
	(0.09)	(-0.11)	(-0.49)	(0.10)
CEO tenure	-0.007**	-0.006**	-0.004	-0.005*
	(-2.25)	(-2.31)	(-1.43)	(-1.93)
CEO duality	0.036	-0.061	0.039	-0.074
	(0.53)	(-1.01)	(0.66)	(-1.16)
Female CEO	-0.048	0.214*	0.042	0.132
	(-0.42)	(1.80)	(0.36)	(1.22)
Constant	0.019	0.106*	0.020	0.151***
	(0.36)	(1.85)	(0.43)	(2.70)
Chi-square test for equality of coefficients	6.29**		10.04***	
(*p*-value)	(0.01)		(<.01)	
Firm fixed effect	YES	YES	YES	YES
Year fixed effect	YES	YES	YES	YES
N	6,027	9,508	7,919	7,616
Adj. R^2	0.706	0.698	0.700	0.713

Table 12. Cross-Sectional Variations: Managerial Private Benefits

This table provides regression estimates specified in Equation (1) for subsamples of high and low managerial private benefits. We use the excess compensation for CEOs to measure the managerial private benefits. CEOs' excessive compensation is measured as the residuals of the regression model of their total and expected pay, as in Core, Guay, and Larcker (2008). All variables are defined in Appendix A. The *t*-statistics with clustered robust standard errors at the firm level (Petersen, 2009) are in parentheses. Statistical significance at the 10%, 5%, and 1% levels are denoted by *, **, and ***, respectively.

	Excessive Compensation for CEOs	
	Low	High
	(1)	(2)
FCA_G	-0.002	0.166***
	(-0.03)	(3.06)
SPF	0.045	0.063
	(0.45)	(1.03)
Firm age	0.541***	0.385***
	(4.08)	(3.15)
Size	0.042	0.057
	(0.88)	(1.39)
Cash	0.189	-0.051
	(1.06)	(-0.33)
R&D	0.450	1.008**
	(0.62)	(2.51)
ROA	-0.170	-0.012
	(-1.21)	(-0.09)
Board size	0.192*	0.162*
	(1.94)	(1.88)
Female director	-0.156	0.182
	(-0.66)	(0.90)
Indep director	0.082	-0.091
	(0.55)	(-0.71)
CEO tenure	-0.005*	-0.002
	(-1.91)	(-0.89)
CEO duality	-0.025	-0.033
	(-0.37)	(-0.57)
Female CEO	-0.018	0.157
	(-0.14)	(1.28)
Constant	-0.853*	-0.301
	(-1.77)	(-0.68)
Chi-square test for equality of coefficients	10.45***	
(*p*-value)	(<.01)	
Firm fixed effect	YES	YES
Year fixed effect	YES	YES
N	7,895	7,640
Adj. R^2	0.718	0.683

Table 13. Cross-Sectional Variations: Relationship-Specific Investments

This table provides regression estimates specified in Equation (1) for subsamples of high and low relationship-specific investments (RSI). Panel A reports the results for RSI measured by a weighted average of the R&D intensities of the firm's customer (supplier) industries (see Ahern and Harford (2014)). Panel B reports the results for RSI measured by whether the firm has core customers who contribute to at least 10% of the total revenue or whose sales are material to the firm's business. All variables are defined in Appendix A. The *t*-statistics with clustered robust standard errors at the firm level (Petersen, 2009) are in parentheses. Statistical significance at the 10%, 5%, and 1% levels are denoted by *, **, and ***, respectively.

Panel A: Industry R&D Intensity

	Cust R&D (Weighted average R&D/Assets)		*Cust R&D (Weighted average R&D/Sales)*		*Supp R&D (Weighted average R&D/Assets)*		*Supp R&D (Weighted average R&D/Sales)*	
	Low	High	Low	High	Low	High	Low	High
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
FCA_G	0.098*	0.102*	0.102*	0.101*	0.089*	0.095*	0.089*	0.088*
	(1.69)	(1.94)	(1.74)	(1.93)	(1.68)	(1.76)	(1.65)	(1.65)
SPF	0.156*	0.014	0.157*	0.016	0.131*	-0.017	0.144**	-0.016
	(1.86)	(0.21)	(1.83)	(0.23)	(1.92)	(-0.23)	(2.02)	(-0.23)
Firm age	0.539***	0.403***	0.556***	0.387***	0.493***	0.436***	0.498***	0.454***
	(4.44)	(2.96)	(4.56)	(2.85)	(3.93)	(3.30)	(3.82)	(3.62)
Size	0.045	0.055	0.049	0.051	0.039	0.066	0.028	0.073
	(0.96)	(1.15)	(1.06)	(1.07)	(0.79)	(1.46)	(0.59)	(1.60)
Cash ratio	0.204	-0.106	0.219	-0.114	0.091	0.069	0.208	-0.044
	(1.29)	(-0.59)	(1.38)	(-0.64)	(0.53)	(0.41)	(1.13)	(-0.28)
R&D	0.961*	0.969**	0.958*	0.960**	0.268	1.341**	0.226	1.385**
	(1.71)	(2.25)	(1.69)	(2.24)	(0.78)	(2.00)	(0.66)	(2.08)
ROA	-0.261*	0.118	-0.282**	0.128	-0.201	0.092	-0.172	0.046
	(-1.90)	(0.82)	(-2.05)	(0.89)	(-1.45)	(0.68)	(-1.18)	(0.35)
Board size	0.218**	0.145	0.217**	0.146	0.171*	0.213**	0.132	0.236***
	(2.30)	(1.57)	(2.28)	(1.60)	(1.76)	(2.41)	(1.29)	(2.76)
Female dir ratio	-0.025	0.185	-0.047	0.193	-0.097	0.249	-0.127	0.261
	(-0.11)	(0.83)	(-0.20)	(0.87)	(-0.41)	(1.12)	(-0.50)	(1.27)
Indep dir ratio	-0.072	0.004	-0.076	0.001	-0.018	-0.007	-0.060	0.013
	(-0.50)	(0.03)	(-0.53)	(0.01)	(-0.13)	(-0.05)	(-0.40)	(0.10)
CEO tenure	-0.003	-0.005	-0.003	-0.005	-0.003	-0.006**	-0.003	-0.006**
	(-1.12)	(-1.57)	(-1.14)	(-1.55)	(-1.23)	(-2.07)	(-0.94)	(-2.34)
CEO duality	-0.038	0.002	-0.039	0.006	-0.108*	0.068	-0.086	0.036
	(-0.73)	(0.03)	(-0.75)	(0.08)	(-1.83)	(1.04)	(-1.38)	(0.58)
Female CEO	-0.074	0.159	-0.070	0.154	0.156	-0.032	0.325***	-0.148
	(-0.62)	(1.27)	(-0.59)	(1.23)	(1.31)	(-0.25)	(3.08)	(-1.15)

Constant	-0.941**	-0.390	-1.000**	-0.334	-0.587	-0.745	-0.472	-0.855*
	(-2.06)	(-0.79)	(-2.18)	(-0.68)	(-1.25)	(-1.59)	(-0.96)	(-1.89)
Chi-square test for equality of coefficients	0.03		0.11		0.56		0.06	
(*p*-value)	(0.87)		(0.74)		(0.46)		(0.81)	
Firm fixed effects	YES	YES	YES	YES	YES	YES	YES	YES
Year fixed effects	YES	YES	YES	YES	YES	YES	YES	YES
No. of obs.	7,788	7,747	7,695	7,840	8,444	7,091	7,474	8,061
Adj. R^2	0.723	0.697	0.721	0.699	0.698	0.727	0.698	0.725

Table 13 (Continued)

Panel B: Corporate Core Customers		
	Corporate Principal Customers	
	Without	With
	(1)	(2)
FCA_G	0.093**	0.106*
	(2.03)	(1.66)
SPF	0.067	-0.043
	(1.18)	(-0.41)
Firm age	0.427***	0.617***
	(3.65)	(3.94)
Size	0.043	0.050
	(0.99)	(0.95)
Cash	0.210	-0.033
	(1.28)	(-0.20)
R&D	0.676	0.790*
	(1.01)	(1.96)
ROA	-0.058	-0.072
	(-0.42)	(-0.55)
Board size	0.166**	0.144
	(1.99)	(1.38)
Female director	-0.182	0.504*
	(-0.95)	(1.82)
Indep director	-0.017	0.049
	(-0.14)	(0.29)
CEO tenure	-0.007**	0.002
	(-2.46)	(0.79)
CEO duality	-0.038	0.054
	(-0.76)	(0.64)
Female CEO	0.146	-0.158
	(1.44)	(-0.96)
Constant	-0.466	-0.907
	(-1.07)	(-1.57)
Chi-square test for equality of coefficients	0.30	
(p-value)	(0.58)	
Firm fixed effect	YES	YES
Year fixed effect	YES	YES
N	10,814	4,721
Adj. R^2	0.696	0.725

Table 14. Effect of FCA Laws on Firm Valuation and Performance

This table shows the results for the effect of positive change in the *Eindex* associated with the exposure to FCA laws on the valuation and performance of firms after using a stacked DiD approach with the event window of [-5, +5] and propensity score matching. The detailed method setting is defined in Table 8. The estimation model is as follows:

$$Firm_{it} = \alpha_0 + \alpha_1 FCA\_G_{it} + \alpha_2 Pst\_Eindex_{it} + \alpha_3 FCA\_G_{it} \times Pst\_Eindex_{it} + \alpha_4 SPF_{it-1} + \beta' Controls_{it-1} + \gamma_i + \omega_t + \varepsilon_{i,t},$$

where $Firm_{it}$ is measured by Tobin's Q or performance for firm *i* in event year *t*. Following the literature, we use ROA, net margin, and gross margin to measure firm performance. FCA_G is a binary variable that equals one if a firm is exposed to at least one state's general FCA through an investment by at least one of that state's pension funds. Pst_Eindex is a binary variable that equals one if a firm experiences an increase in the *Eindex* after the passage of the FCA law. *SPF* is a binary variable that equals one if at least one state pension fund invested in the firm's shares in the lagged year. The other variables are defined in Equation (1). γ and ω denote firm and event year fixed effect, respectively; and ε is the error term. All variables are defined in Appendix A. The *t*-statistics with clustered robust standard errors at the firm level (Petersen, 2009) are in parentheses. Statistical significance at the 10%, 5%, and 1% levels are denoted by *, **, and ***, respectively.

	Q	*ROA*	*Net Margin*	*Gross Margin*
	(1)	(2)	(3)	(4)
FCA_G	0.157**	-0.007	-0.004	-0.001
	(2.50)	(-1.61)	(-0.56)	(-0.21)
Pst_Eindex	-0.140***	0.001	0.000	-0.006***
	(-6.57)	(0.52)	(0.03)	(-2.77)
FCA_G×Pst_Eindex	-0.296***	-0.010***	-0.010*	-0.006***
	(-14.46)	(-2.88)	(-1.91)	(-2.66)
SPF	-0.148*	0.017**	0.023	0.000
	(-1.92)	(1.97)	(1.45)	(0.01)
Firm age	0.227*	0.023***	0.037**	-0.004
	(1.89)	(2.59)	(2.00)	(-0.42)
Size	-0.475***	-0.026***	-0.030***	0.011**
	(-8.68)	(-5.65)	(-3.50)	(2.03)
Cash ratio	0.941***	0.103***	0.104***	0.043**
	(3.90)	(5.36)	(2.93)	(2.20)
R&D	2.846**	0.027	-0.044	-0.031
	(2.35)	(0.33)	(-0.23)	(-0.31)
ROA	1.596***			
	(5.12)			
CF Ratio		0.372***	0.536***	0.304***
		(5.27)	(5.10)	(5.58)
Board size	-0.178*	-0.007	-0.002	0.000
	(-1.83)	(-0.81)	(-0.13)	(0.04)
Female dir ratio	0.360	0.004	0.002	0.006
	(1.62)	(0.21)	(0.06)	(0.36)
Indep dir ratio	-0.080	0.018	0.035**	0.005
	(-0.57)	(1.56)	(2.01)	(0.38)
CEO tenure	0.006**	0.000**	0.001***	0.000
	(2.08)	(2.43)	(2.74)	(0.97)
CEO duality	0.048	0.005	0.004	0.003
	(1.04)	(0.88)	(0.47)	(0.51)
Female CEO	0.048	-0.011	-0.010	0.005
	(0.44)	(-1.07)	(-0.76)	(0.53)
Constant	5.045***	0.104***	0.041	0.263***
	(10.71)	(3.29)	(0.71)	(6.34)
Firm fixed effect	YES	YES	YES	YES
Event year fixed effect	YES	YES	YES	YES
N	10,876	9,968	9,930	9,930
Adj. R^2	0.658	0.574	0.537	0.882

Table 15. Effect of FCA Laws on the Disclosure Quality of Firms

This table shows the results for the effect of positive change in the *Eindex* associated with the exposure to FCA laws on the Disclosure Quality of firms using a stacked DiD approach with the event window of [-5, +5] and propensity score matching. The detailed method setting is defined in Table 7. The estimation model is as follows:

$$Disclosure\ Quality_{it} = \alpha_0 + \alpha_1 FCA\_G_{it} + \alpha_2 Pst\_Eindex_{it} + \alpha_3 FCA\_G_{it} \times Pst\_Eindex_{it} + \beta' Controls_{it-1} + \gamma_i + \omega_t + \varepsilon_{i,t},$$

where $Disclosure\ Quality_{i,t}$ is defined by Chen, Miao, and Shevlin (2015) as the disaggregation quality for firm *i* in year *t*. Following the literature, we use ROA, net margin, and gross margin to measure firm performance. FCA_G is a binary variable that equals one if a firm is exposed to at least one state's general FCA through an investment by at least one of that state's pension funds. Pst_Eindex is a binary variable that equals one if a firm experiences an increase in the Eindex around the passage of the FCA law. *SPF* is a binary variable that equals one if at least one state pension fund invested in the firm's shares in the lagged year. The other variables are defined in Equation (1). γ and ω denote firm and event year fixed effect, respectively, and ε is the error term. All variables are defined in Appendix A. The *t*-statistics with clustered robust standard errors at the firm level (Petersen, 2009) are in parentheses. Statistical significance at the 10%, 5%, and 1% levels are denoted by *, **, and ***, respectively.

	All sample	*SPF=1 only*
	(1)	(2)
FCA_G	0.003	0.002
	(0.51)	(0.39)
Pst_Eindex	-0.000	-0.000
	(-0.01)	(-0.01)
FCA_G× Pst_Eindex	-0.002*	-0.002*
	(-1.71)	(-1.68)
SPF	-0.020***	
	(-3.16)	
Firm age	-0.003	-0.003
	(-0.37)	(-0.33)
Size	-0.004	-0.004
	(-1.03)	(-1.07)
Cash ratio	0.020**	0.020*
	(2.01)	(1.91)
R&D	0.020	0.018
	(0.50)	(0.45)
ROA	0.003	0.003
	(0.36)	(0.33)
Board size	-0.010	-0.010
	(-1.35)	(-1.37)
Female dir ratio	-0.010	-0.011
	(-0.70)	(-0.79)
Indep dir ratio	0.009	0.010
	(0.73)	(0.84)
CEO tenure	-0.000	-0.000
	(-0.20)	(-0.19)
CEO duality	-0.006	-0.006
	(-1.36)	(-1.35)
Female CEO	0.018**	0.018**
	(2.54)	(2.49)
Constant	0.854***	0.834***
	(25.59)	(24.88)
Firm fixed effect	YES	YES
Event year fixed effect	YES	YES
N	3,057	3,031
Adj. R^2	0.834	0.832

Internet Appendix

Table IA1. Limit the Sample with the Same CEO over the Event Window

This table shows the results of regression models as specified in Equation (1) using a limited sample after removing firms that replace their CEOs during the event window. The dependent variable is the *Eindex* proposed by Bebchuk, Cohen, and Ferrell (2009). *FCA_G* is a binary variable that equals one if a firm is exposed to a general FCA, and zero otherwise. *SPF* is a binary variable that equals one if a firm's shares are owned by at least one state pension fund in the lagged year. All variables are defined in Appendix A. The *t*-statistics with clustered robust standard errors at the firm level (Petersen, 2009) are in parentheses. Statistical significance at the 10%, 5%, and 1% levels are denoted by *, **, and ***, respectively.

	No Control		Control included	*SPF* = 1 only
	(1)	(2)	(3)	(4)
FCA_G	1.213***	0.101***	0.130**	0.128**
	(29.21)	(2.85)	(2.34)	(2.14)
SPF	-0.378***	0.101*	0.108	
	(-4.06)	(1.78)	(1.41)	
Firm age			0.570***	0.597***
			(4.45)	(4.65)
Size			0.099*	0.098*
			(1.89)	(1.85)
Cash			-0.216	-0.210
			(-1.18)	(-1.14)
R&D			1.621***	1.630***
			(2.72)	(2.68)
ROA			-0.073	-0.079
			(-0.49)	(-0.52)
Board size			0.093	0.089
			(0.85)	(0.80)
Female director			-0.234	-0.197
			(-0.92)	(-0.77)
Indep director			-0.136	-0.105
			(-0.94)	(-0.73)
CEO tenure			-0.008***	-0.008***
			(-2.61)	(-2.59)
CEO duality			0.037	0.048
			(0.46)	(0.59)
Female CEO			0.055	0.042
			(0.34)	(0.25)
Constant	2.284***	1.296***	-1.005**	-0.961*
	(23.19)	(19.53)	(-2.01)	(-1.95)
Firm fixed effect	NO	YES	YES	YES
Year fixed effect	NO	YES	YES	YES
N	12,503	12,503	7,666	7,417
Adj. R^2	0.041	0.704	0.696	0.697

Table IA2. Limit the Sample Period before the Dodd-Frank Act

This table shows the results of regression models as specified in Equation (1) using a limited sample to the period before the Dodd-Frank Act of 2010. The dependent variable is the *Eindex* proposed by Bebchuk, Cohen, and Ferrell (2009). *FCA_G* is a binary variable that equals one if a firm is exposed to a general FCA, and zero otherwise. *SPF* is a binary variable that equals one if a firm's shares are owned by at least one state pension fund in the lagged year. All variables are defined in Appendix A. The *t*-statistics with clustered robust standard errors at the firm level (Petersen, 2009) are in parentheses. Statistical significance at the 10%, 5%, and 1% levels are denoted by *, **, and ***, respectively.

	No Control		Control included	*SPF* = 1 only
	(1)	(2)	(3)	(4)
FCA_G	0.696***	0.083***	0.081**	0.101***
	(26.50)	(3.15)	(2.19)	(2.70)
SPF	0.267***	0.004	-0.002	
	(2.94)	(0.05)	(-0.02)	
Firm age			0.151	0.172
			(1.11)	(1.27)
Size			0.023	0.019
			(0.54)	(0.44)
Cash			0.013	0.020
			(0.08)	(0.12)
R&D			0.300	0.292
			(0.77)	(0.75)
ROA			-0.000	-0.001
			(-0.00)	(-0.00)
Board size			0.094	0.091
			(1.11)	(1.06)
Female director			0.079	0.104
			(0.33)	(0.43)
Indep director			-0.027	-0.025
			(-0.23)	(-0.22)
CEO tenure			-0.005**	-0.006**
			(-2.12)	(-2.14)
CEO duality			0.032	0.031
			(0.60)	(0.59)
Female CEO			0.058	0.022
			(0.37)	(0.14)
Constant	1.230***	1.369***	0.617	0.581
	(13.19)	(20.02)	(1.35)	(1.30)
Firm fixed effect	NO	YES	YES	YES
Year fixed effect	NO	YES	YES	YES
N	11,176	11,176	7,057	6,958
Adj. R^2	0.039	0.712	0.735	0.735