CLASSIFIED BOARDS AND FIRM VALUE

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ABSTRACT

Classified boards constitute one of the most potent takeover defenses for U.S. firms today. However, as with takeover defenses more generally, economic theory offers an ambiguous prediction as to the effect that classified boards have on bottom-line firm value. A resolution of this ambiguity will require sound and convincing empirical methodology. In an effort to address limitations in the existing empirical literature, this article approaches the relationship between corporate governance and firm value while taking various measures to account for unobserved sources of heterogeneity across firms. Using the instrumental variables model developed by Hausman and Taylor,¹ I find evidence of a negative and statistically significant association between classified board status and firm value. I confirm these findings using a variation of the difference-in-difference-in-difference model recently employed by Rauh.² However, using quantile regressions, I find evidence suggesting that this negative association may be concentrated along the upper tail of the distribution of firm value.

I. INTRODUCTION

Over the last several decades, academics and practitioners have devoted significant attention to the benefits and costs of takeover defenses. The question remains, however, as to what effect such defenses ultimately have on the value of the firm. Given the variation in legal rules across jurisdictions and the discretion that firms possess in structuring their corporate governance profiles, firms do vary notably with respect to their ability to fend off unsolicited takeover bids. This variation offers the

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potential for empirical identification of the relationship between corporate governance and firm value.

Absent strong empirical guidance, analysts will likely remain uncertain as to the nature and extent of the relationship between governance and value. Theoretical analysis alone leaves academics and practitioners in numerous camps. Some claim that the threat of takeovers may alleviate ex ante agency costs and properly align incentives by offering a form of discipline for ineffective management.\(^3\) Where managers and directors face little threat of removal by outsiders and inadequate monitoring by shareholders, those managers and directors may exploit the consequent breathing room by engaging in numerous forms of self-serving behavior: extracting unreasonable compensation and perquisites, building inefficient "empires," rejecting value-enhancing acquisition offers, or redirecting business operations towards self-interested ventures.\(^4\) With these possibilities in mind, such analysts claim that takeover defenses will only serve to remove the ability of the market to limit destructive behavior on the part of managers.\(^5\)

On the other hand, many claim that takeover defenses allow management to avoid unnecessary distractions and remain focused on current operations and long-term planning.\(^6\) The threat of a hostile takeover may cause managers to over-discount future periods and focus excessively on short-term results.\(^7\) Additionally, takeover defenses may provide managers and directors with an important source of bargaining leverage in friendly acquisition discussions, allowing them to negotiate higher acquisition premiums.\(^8\)

Recognizing the net theoretical ambiguity, academics have devoted copious efforts to empirical testing of the impacts of takeover defenses. Empirical investigations have taken two approaches: (1) measure the impact of defenses on the various intermediate stages (e.g., the effect on premium bids in friendly acquisitions),\(^9\) and (2) account for all such


\(^4\)Id. at 994.

\(^5\)See, e.g., id. at 991-95.

\(^6\)See, e.g., id. at 1011.

\(^7\)See, e.g., Bebchuk, supra note 3, at 1011.

\(^8\)See Guhan Subramanian, *Bargaining in the Shadow of Takeover Defenses*, 113 YALE L.J. 621, 629-30 (2003). For a general review of the numerous arguments made in favor of allowing directors to block acquisition efforts, see Bebchuk, supra note 3.

\(^9\)See, e.g., Subramanian, supra note 8, at 667-81 (estimating the effect of the strength of takeover defenses on bid premiums received by shareholders in negotiated acquisitions).
behaviors simultaneously by measuring the net effect on firm value.\textsuperscript{10} Despite sacrificing a more structural specification, this article follows the second of these approaches and employs a reduced form model to estimate the impact of classified boards on firm value. Given the difficulties involved in modeling the aggregation of the intermediate channels, the reduced-form approach offers perhaps the only hope for evaluating the question in which we are ultimately interested: does governance enhance bottom-line shareholder value? With the proper methodology, this empirical exercise may prove quite instructive to both firms and policymakers alike.

A growing body of empirical literature has taken the second approach and tackled estimation of the relationship between governance and firm value.\textsuperscript{11} The current literature, however, focuses excessively on measuring the effect of an index of governance provisions.\textsuperscript{12} Such studies suffer from numerous shortcomings, most significant of which is the fact that a given index level\textsuperscript{13} does not represent a unique measure of entrenchment. In all likelihood, it is the precise composition of that index that matters. Firms with poison pills and limits to amend charters may earn an index value of 2. Nonetheless, such firms may be more embedded than others with antigreenmail provisions and limitations on director liability. Treating these separate firms as having identical governance profiles will cloud any attempt to measure a precise link between corporate governance and firm value. Partially avoiding these limitations, this article looks within these indices and focuses on the specific impact of one of the strongest takeover defenses available: classified boards. In firms with classified boards (also known as "staggered boards"), only a fraction of the directors stand for election each year, thereby forcing stockholders to wait for at least two annual elections before they can acquire control of the board of directors.

In addition to offering a more structural econometric model, by looking within the governance index, this approach offers more useful guidance to judges, legislators, directors, and other parties involved in shaping corporate governance profiles. Such parties do not face the decision of whether or not to encourage more or less of the amorphous

\textsuperscript{10}See, e.g., Paul Gompers et al., Corporate Governance and Equity Prices, 118 Q. J. ECON. 107, 125-29 (2003) (estimating the effect of an index of governance provisions on firm value).

\textsuperscript{11}See, e.g., id.

\textsuperscript{12}See, e.g., id.

\textsuperscript{13}In such studies, index values for each firm are constructed by adding one point for every corporate governance provision maintained by such firm (out of the available provisions comprising the index). See, e.g., id at 109.
thing called a "takeover defense." Rather, they must know how to address each provision independently. With this in mind, it becomes important to estimate the isolated effects of provisions like classified boards.

One must also recognize that governance provisions are not random characteristics of firms. In many cases, firms select the structures they desire, and in the process, they may trade off other governance provisions. In addition to illustrating the limitations of an index approach, these trade-offs highlight the fact that the composition of governance structures is changing over time. Business combination statutes may have been meaningful in earlier times. However, devices such as classified boards and other takeover provisions that limit one's ability to wage a proxy contest are of far greater significance today and may even render most other defenses irrelevant. After Delaware's validation of the classified board/poison pill combination in the mid-1990s, one could argue that classified boards now offer the most potent protection available. Consequently, despite addressing a seemingly dated issue, by focusing on classified boards, this article offers novel contributions to the existing corporate governance literature.

In a recent study, Bebchuk and Cohen respond to this transformation in the corporate governance landscape and take an intensive look at the relationship between firm value and classified boards. Using pooled ordinary least squares (OLS) estimators on an unbalanced sample compiled by the Investor Responsibility Research Center (IRRC), they find that classified boards are associated with a reduction in firm value. By focusing on classified boards specifically, Bebchuk and Cohen's analysis

14Of course, policymakers also want to understand the interactions among various governance devices. Nonetheless, index approaches provide an imperfect means of studying interactive mechanisms.
15For instance, firms with classified boards and poison pills may decide that they need little else in the way of protection.
16See John C. Coates IV, Takeover Defenses in the Shadow of the Pill: A Critique of the Scientific Evidence, 79 Tex. L. Rev. 271, 320-23 (2000); see also Gompers et al., supra note 10, at 111-12 (noting that the dynamics of modern takeover battles, which are shaped by blank check preferred stock provisions, classified board provisions, special meeting limitations, and provisions limiting stockholder actions by written consent, may have "rendered all other defenses superfluous").
17See Guhan Subramanian, The Disappearing Delaware Effect, 20 J. L., Econ., & Org. 32, 52-53 (2004) (noting that the outcome of three takeover contests involving Delaware firms in the mid-1990s may have solidified the ability of a target with a staggered board to maintain its poison pill after having lost the first round of a proxy contest against a hostile bidder).
19Id. at 420-24.
20Id. at 424.
improves upon those studies concerned exclusively with governance indices. Like most of the relevant literature, however, their analysis suffers from various econometric limitations. Given the endogenous nature of classified board status, Bebchuk and Cohen's model raises the concern that other factors missing from the specification are responsible for the estimated relationship. Particularly, the failure to account for unobservable firm-fixed effects may be generating a significant bias in the estimated coefficients.

The aim of this article is to build upon the empirical groundwork laid by Bebchuk and Cohen and the corresponding studies. I explore various econometric models in an attempt to generate unbiased coefficients that account for unobservable factors. First, unlike much of the recent governance literature, I embrace the panel nature of the IRRC data. With endogeneity and omitted variables concerns and a lack of potential instruments, it may be imprudent to rely solely on pooled OLS specifications. Containing both cross-sectional and time-series variation, panel data offers powerful econometric tools that may allow one to eliminate the influence of unobservable factors.

However, standard panel data fixed-effects models, which exploit variation in firms over time, cannot identify the relationship between classified boards and firm value, given the time-invariant nature of the classified board variable. To avoid this limitation and to allow for the use of valuable cross-sectional information, I employ the instrumental variables model developed by Hausman and Taylor. If the identification conditions are met, the Hausman-Taylor model controls for firm-fixed effects while still identifying the coefficient of the time-invariant variable of interest. Using the same IRRC dataset as Bebchuk and Cohen, I estimate a negative and statistically significant relationship between classified boards and firm value.

As an alternative to the Hausman-Taylor model, I estimate a difference-in-difference-in-difference (DDD) model in an attempt to account for unobservable factors that are correlated with both classified board incidence and firm value. I utilize several dimensions of variation in structuring the DDD estimator, including the state of incorporation, pre- and post-1995, and the likelihood of having a classified board. Under this alternative approach, I again document a negative and statistically significant relationship between classified boards and firm value. The

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21Hausman & Taylor, supra note 1, at 1378-79, 1383-89.

22Id.

23Bebchuk & Cohen, supra note 18, at 418.
estimated effects are of a considerable magnitude, suggesting that the relationship is also economically meaningful.

As a final robustness check on the model and on the underlying theory, I run quantile regressions of firm value on classified board status. Quantile specifications are robust to the impact of outlying observations and allow for relaxation of the assumption of homogenous effects along the full distribution of the dependent variable.24 The results of this quantile approach indicate that classified boards have either no effect or a slight positive effect on firm value for the least valuable firms. However, for firms at the upper end of the distribution of firm value, the results indicate a stronger negative effect. This pattern suggests that classified boards may not be wholly negative for firms, despite the estimation of a negative relationship on average. Rather, classified boards may generate positive contributions to firm value in certain circumstances.

This article is divided into seven parts. Part II surveys prior empirical work regarding the link between governance and value. Part III describes the data and provides summary statistics. Part IV presents the methodology and results for the Hausman-Taylor estimator. Part V explains the methodology and results for the triple-differences (DDD) estimator. Part VI discusses the application of quantile regressions. Finally, Part VII concludes.

II. LITERATURE REVIEW

A wealth of literature has explored the consequences that attach to corporate governance mechanisms. Many studies take a focused approach and examine the effects of corporate governance on specific managerial behaviors and specific measures of firm performance. Such studies are best viewed as "intermediate" inquiries, given that they leave ambiguous the net effect on firm valuation.

The empirical results derived from this "intermediate" literature present mixed findings with respect to the impact of governance structures on firm behavior. Bertrand and Mullainathan find a rise in CEO compensation as a result of state antitakeover laws, consistent with the "skimming" hypothesis, whereby executives provide themselves greater

compensation as a result of entrenchment.\textsuperscript{25} Using a sample of all negotiated acquisitions of U.S. public company targets, Subramanian rejects the hypothesis that the enhanced bargaining power in friendly negotiations attributable to takeover defenses leads to an increase in negotiated bid premiums.\textsuperscript{26} Garvey and Hanka find that "second generation" antitakeover laws significantly reduced debt loads taken on by firms, providing evidence in support of the views that takeover threats discipline management and that managerial discretion, as influenced by the takeover environment, can impact a firm's capital structure.\textsuperscript{27} Using a sample of firms between 1979 and 1985, Johnson and Rao find no significant evidence of a deleterious effect of antitakeover charter amendments on (1) operating income, (2) research and development, (3) capital expenditures, and (4) operating and overhead costs.\textsuperscript{28}

Other studies that would be classified as "intermediate" include those that address the trade-offs that firms make among alternative governance structures. In one such study, Rauh finds that Delaware's validation of the classified board/poison pill combination in the mid-1990s was associated with a significant negative effect on employee ownership of own-company stock, supporting an inference that management's inducement of employees to acquire own-company stock in defined compensation plans is a form of takeover defense.\textsuperscript{29}

Results from this "intermediate" literature have motivated numerous analysts to estimate the aggregate effect of corporate governance mechanisms on firm value. Some studies take a global approach and explore such relationships using data from a cross section of countries. For instance, La Porta and his colleagues find evidence of higher valuation of firms in countries that provide greater protection for minority


\textsuperscript{26} Subramanian, supra note 8, at 669-79.


\textsuperscript{29} Rauh, supra note 2, at 401.
shareholders. While perhaps limited in terms of generality, intra-country studies avoid many of the limitations of cross-country regressions, such as sample size and omitted variables. Fortunately, corporate governance profiles vary enough across U.S. corporations to identify the effects of these structures.

Originating a line of intra-country literature, Morck, Shleifer, and Vishny document a significant non-monotonic relationship between firm value and managerial ownership, with firm value increasing, then decreasing, and finally increasing again as the proportion of shares held by management rises. Other intra-country approaches employ event-study methodologies, examining market responses to the occurrence of certain events. For instance, Ryngaert examines changes in stock prices following the adoption of poison pills. Karpoff and Malatesta find that the announcement of state antitakeover legislation is associated with a small but statistically significant decline in the stock prices of affected firms.

Results derived from such event studies, however, suffer from potential estimation bias, given that the relevant announcement (e.g., the

\[\text{Rafael La Porta et al., Investor Protection and Corporate Valuation, 57 J. FIN. 1147 (2002).}\]


\[\text{Michael Ryngaert, The Effect of Poison Pill Securities on Shareholder Wealth, 20 J. FIN. ECON. 377, 410 (1988) (finding evidence of a stock price decline associated with the announcement of the adoption of the most restrictive forms of poison pills).}\]

announcement of a poison pill adoption) may also provide an informative signal to investors, leaving it difficult to isolate the effect of the relevant corporate governance mechanism itself.\textsuperscript{34} Moreover, those event studies that follow poison pill adoptions fail to recognize that a firm need not have a pill in place to be able to benefit from its potential protection, given the ability to adopt a pill after a hostile bid has commenced.\textsuperscript{35}

Measuring the aggregate effect of Delaware state corporate law on shareholder wealth, Robert Daines finds that firms incorporated in Delaware are associated with higher Tobin's Q values than their non-Delaware counterparts (where Tobin's Q is used as an estimate of firm value).\textsuperscript{36} Daines avoids the pitfalls of event study methodologies by employing cross-sectional regressions on a large sample of exchange-traded corporations.\textsuperscript{37} Daines' results support the view that state corporate law is an important feature of corporate governance.\textsuperscript{38}

A recent line of research has studied the impact on firm valuation of a broad index of governance provisions that proxy for the strength of shareholder rights. Pioneering these studies, Gompers, Ishii, and Metrick construct one such index (the GIM Index) and find that firms with stronger shareholder rights are associated with higher firm value.\textsuperscript{39} However, it is difficult to interpret these results in light of various limitations of an index approach. There is little reason to believe that each component of the index will affect firm value in the same manner. Moreover, the provisions of the index may be endogenous to each other. For instance, a firm that adopts a classified board may find many other provisions unnecessary and consequently drop them. The net effect may be a reduction in the index value, despite an increase in entrenchment potential. A more powerful

\textsuperscript{34}See Coates, supra note 16, at 298-306.
\textsuperscript{35}See id. at 286-88.
\textsuperscript{37}Id. at 531-32.
\textsuperscript{38}Id. at 556.
\textsuperscript{39}Gompers et al., supra note 10, at 126-29. For related work, see Belen Villalonga & Raphael Amit, How Do Family Ownership, Control and Management Affect Firm Value?, 80 J. Fin. Econ. 385, 397-401 (2006) (estimating the effect of family ownership, control, and management on firm value while controlling for the GIM Index); K.J. Martijn Cremers & Vinay B. Nair, Governance Mechanisms and Equity Prices, 60 J. Fin. 2859, 2887-88 (2005) (examining the interaction between internal and external governance mechanisms and finding that firms with high takeover vulnerability, as proxied by two alternative governance indices, but no public pension fund blockholder, have a higher firm value than firms with both high takeover vulnerability and public pension fund blockholders); Mark Klock et al., Does Corporate Governance Matter to Bondholders?, 40 J. Fin. & Quant. Anal. 693, 708-09 (2005) (finding evidence that strong antitakeover provisions, as proxied by high levels of the GIM Index, are associated with a lower cost of debt financing).
approach would be to look within this index and identify the effects of specific structures. Bebchuk, Cohen, and Ferrell move in this direction by constructing a smaller index, which, they believe, accounts for the bulk of the entrenchment potential among the twenty-four provisions identified in the GIM Index. This narrower index focuses on four constitutional limitations on shareholder voting power (including classified boards) and two key takeover readiness measures. Their findings confirm the negative relationship found by Gompers, Ishii, and Metrick.

Several recent studies have focused on the relationship between classified board status and firm performance. Using a dataset that includes all hostile bids between 1996 and 2000, Bebchuk, Coates, and Subramanian find evidence that the possession of an "effective" staggered board in the face of a hostile bid nearly doubles the odds of remaining independent, halves the odds that a first bidder will be successful, and reduces the odds that a target will be forced to sell to a white knight. Moreover, they find that "effective" staggered boards reduce expected returns for shareholders in target firms during the period of time after a hostile bid is launched. Bebchuk and Cohen address the same question motivating this article: what effect do classified boards have on bottom-line firm value? Running pooled OLS regressions, they document a negative and statistically significant relationship between classified board status and firm value, where firm value is proxied by industry-adjusted Tobin's Q. Estimating Fama-MacBeth, pooled OLS and certain other specifications on a sample of 2,021 firms between 1995 and 2002, Faleye also documents a negative association between classified boards and Tobin's Q.

Including the Tobin's Q value in 1990 (i.e., a lagged dependent variable) as an additional control variable in their specification, Bebchuk

41See Bebchuk et al., supra note 40, at 2.
42Compare id. at 3-4, with Gompers et al., supra note 10, at 126-29.
43Lucian Bebchuk et al., The Powerful Antitakeover Force of Staggered Boards: Theory, Evidence and Policy, 54 Stanford L. Rev. 887, 890, 931 (2002). Bebchuk, Coates, and Subramanian classify a staggered board as an "effective staggered board" if (i) the staggered board is established in the firm's charter, (ii) the firm's directors may only be removed for cause and (iii) the firm's shareholders are prohibited from "pack[ing] the board" by increasing the number of directors and filling the resulting vacancies. Id. at 895.
44Id. at 937-39.
45Bebchuk & Cohen, supra note 18, at 410.
46Id. at 420-24.
and Cohen continue to find a negative association between classified boards and firm value. 48 As a result, they argue that the relationship may be causal because the negative correlation between classified boards and post-1995 firm value cannot be explained by the initial selection of classified boards by firms with low Qs in 1990. 49 Supporting this claim, Bebchuk and Cohen note that classified board status was largely determined as of 1990. 50 However, while Bebchuk and Cohen's analysis may alleviate concerns as to the direction of the association, it still does not ensure that the estimated relationship itself is unbiased. To the extent there are unobserved fixed characteristics of firms that are correlated with both firm value and classified board status, one will be unable to disentangle the effects of classified boards from the effects of such unobservable fixed factors, even after controlling for the 1990 Tobin's Q value. 51 More generally, the use of lagged dependent variables will not address correlations with unobservable fixed characteristics. 52 Part IV of this article attempts to account for these fixed factors in a more direct manner.

III. DATA

A. Sources

The basic data source for this analysis is the Investor Responsibility Research Center (IRRC). The IRRC published seven volumes of governance data between 1990 and 2004 in Corporate Takeover Defenses. 53 The IRRC data is available in electronic form at the Wharton Research Data Services (WRDS) website. 54 Each IRRC volume contains

49 Id. Faley also addresses potential self-selection concerns by, among other things, including controls for prior performance. See Faley, supra note 47, at 11-12.
50 Bebchuk & Cohen, supra note 18, at 426.
51 Moreover, the various steps taken by Faley, supra note 47, at 11-13, to address a possible self-selection bias, including the use of controls for prior performance and the use of average historical Q values as instruments, will not address any bias generated from the correlation between classified board status and an unobservable fixed firm effect.
52 Even if one includes lagged dependent variables as additional regressors, "differencing" steps, such as "first differencing" or transforming the observations into deviations from individual means (i.e., fixed effects regression), are likely still required to remove unobservable fixed factors. See, e.g., Manuel Arellano & Stephen Bond, Some Tests of Specification for Panel Data: Monte Carlo Evidence and an Application to Employment Equations, 58 REV. ECON. STUDIES 277, 280 (1991) (transforming specifications into "first differences" in order to remove individual effects).
comprehensive information pertaining to the corporate governance structures of individual firms. The IRRC collects such information from numerous resources: corporate bylaws, corporate charters, annual reports, 10-Ks, 10-Qs, and proxy statements. The IRRC data offers a comprehensive representation of the nation's largest corporations. In the 1990 volume alone, the IRRC firms accounted for over 90% of the total market capitalization of the nation's major stock exchanges. The data covers all firms in the Standard & Poor's 500, along with firms represented on annual lists of the largest corporations in the United States, as published by various national business magazines. The resulting dataset is an unbalanced panel of 3,030 individual firms over the fourteen-year period between 1990 and 2004. Each IRRC volume contains data on 1,400 to 1,900 firms. Many firms are represented continuously throughout the relevant time period; however, the represented list of firms does change somewhat among the IRRC volumes.

The IRRC data documents twenty-four unique governance structures, representing charter provisions, bylaw provisions, exposure to six state takeover laws, and other firm-specific governance rules. Included in this list of twenty-four provisions is the presence of a classified board. While not provided in the electronic data made available by the WRDS, the IRRC volumes do indicate, for each firm with a classified board, whether such firm established the classified board in its bylaws or in its charter. Most of the analysis below uses a general indicator for a classified board. However, given that the effectiveness of classified boards is enhanced when they are established in a firm's charter, I also explore specifications that measure the impact of charter-based classified boards.

Each specification employed below uses classified board status (general or charter-based) as the regressor of interest. Each specification, however, also controls in varying ways for the remainder of the IRRC governance provisions. Failure to account for these measures may bias the estimated coefficients because classified board status is likely correlated with the presence of other governance devices. In each reported regression,

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55 Gompers et al., supra note 10, at 110.
56 Id. at 111.
57 Id. at 110-11.
58 For one of the specifications used in Part IV below, I form a balanced panel consisting of 580 firms.
59 Gompers et al., supra note 10, at 111.
60 Bebchuk et al., supra note 43, at 895.
61 For the charter-based specifications explored below, I only use data for the years 1990-2002.
I control for the remaining IRRC provisions by including the GIM Index, adjusted accordingly to remove the contribution of classified boards (the adjusted GIM Index). To avoid imposition of linearity in the relationship between the adjusted GIM Index and firm value, I take a flexible approach and include the index value and its square. As suggested above, aggregating governance provisions into an index may limit the interpretability and flexibility of the model. To alleviate the reliance on a broad index, I also run unreported regressions in which I follow the approach of Bebchuk, Cohen, and Ferrell and break the adjusted GIM Index into two pieces. The first piece (the "entrenchment" index) is a smaller index of five provisions that, together with classified boards, are arguably most responsible for managerial entrenchment. These five provisions include limits to amend bylaws, limits to amend charters, supermajority requirements for mergers and charter amendments, poison pills, and golden parachutes. The second piece (the "other-provisions" index) consists of an index covering the remainder of the IRRC provisions.

To form additional controls, I obtain data on firm financials from Compustat's Industrial Annual files. I match this financial data to the IRRC sample primarily using six-digit CUSIPs. In many cases, however, I also use ticker symbols, stock exchanges, and company names to assist in the matching process. Most of the financial controls employed are in ratio format (e.g., return-on-assets) and thus do not require transformation into real values. To account for firm size, however, I control for the total real asset level of each firm, using Consumer Price Index (CPI) data obtained from the Bureau of Labor Statistics to transform nominal asset measures into real 2002 dollars. Using CPI measures, I also transform annual sales figures into real values and then calculate the three-year average annual growth rate in real sales for each observation. I use the growth rate in real sales to proxy for the investment and growth opportunities of the firm.

Using the Compustat financials, I calculate Tobin's Q for each firm/year observation and use this measure as a proxy for firm value. In

64Id. at 2.
65Id. at 11-13.
66Compustat data is compiled by Standard and Poor's and contains fundamental company and market information on securities representing over 90% of the world's total market capitalization. For more information, see http://www.compustat.com.
68See La Porta et al., supra note 30, at 1159.
general terms, Tobin's $Q$ is defined as the ratio of the market value of the firm to the replacement cost of its assets.\textsuperscript{69} While alternative formulations of this variable exist,\textsuperscript{70} I follow Bebchuk and Cohen and Gompers, Ishii, and Metrick and calculate Tobin's $Q$ as the market value of assets divided by the book value of assets, where the market value of assets equals the book value of assets plus the market value of common stock minus the sum of the book value of common stock and the level of deferred taxes.\textsuperscript{71} Industry-adjusted Tobin's $Q$ is calculated by subtracting from each observation's Tobin's $Q$ value the median $Q$ value for the firm's corresponding industry/year group.\textsuperscript{72} Many of the regressions reported in this article use this industry-adjusted $Q$ measure as the dependent variable. In other specifications, however, I control more completely for industry effects by using regular Tobin's $Q$ as the dependent variable, along with a full set of industry dummies.

I construct insider ownership measures for each firm using data matched from Compustat's ExecuComp database.\textsuperscript{73} For each firm, I calculate the fraction of total outstanding shares held by directors and officers. Following Bebchuk, Cohen, and Ferrell, I calculate firm age by determining the first time information about the relevant firm's stock became available in the CRSP monthly database.\textsuperscript{74} I match CRSP data to

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\item \textsuperscript{69}E.g., Kee H. Chung & Stephen W. Pruitt, A Simple Approximation of Tobin's $Q$, 23 FIN. MGMT. 70, 70 (1994).
\item \textsuperscript{70}See, e.g., id. at 70-74 (comparing alternative Tobin's $Q$ calculations).
\item \textsuperscript{71}Bebchuk & Cohen, supra note 18, at 420; Gompers et al., supra note 10, at 151.
\item \textsuperscript{72}Bebchuk & Cohen, supra note 18, at 420. Industry measures are generated using the Compustat data. For each of the forty-eight Fama-French industry groups, I calculate annual median values across all of the firms represented in the Compustat database. For a list of the forty-eight Fama-French industry groups, see Eugene Fama & Kenneth French, Industry Costs of Equity, 43 J. FIN. ECON. 153, 179-181 (1997). After assigning Fama-French industry groups to each of the IRRC firms, I match the Compustat-derived industry measures to the IRRC sample and then calculate industry-adjusted $Q$ values. Finally, following Daines, I trim observations with industry-adjusted Tobin's $Q$ in the upper and lower 1% of the sample. See Daines, supra note 36, at 530. However, the results presented below are robust to alternative treatments of these observations, including no modification at all.
\item \textsuperscript{73}Standard & Poor's provides annual executive compensation data in the ExecuComp database. For more information, see http://www.compustat.com.
\item \textsuperscript{74}BEBCHUK ET AL., supra note 40, at 15. CRSP stands for The Center for Research in Securities Prices. CRSP was established by the University of Chicago's Graduate School of Business and collects historical data on the NYSE, AMEX, NASDAQ and other major securities markets. See http://www.crsp.chicagogsb.edu (last visited Jan. 18, 2007). CRSP monthly data is available from 1926. Id. Thus, firm ages for those firms that incorporated prior to 1926 are censored at this end point. Censored right-hand-side variables may lead to estimation bias. See ROBERTO RIGOBON & THOMAS STOKER, CENSORED REGRESSIONS AND EXPANSION BIAS 1-9 (MIT Sloan Sch. of Mgmt., Working Paper No. 4451-03, 2003), available at http://papers.ssrn.com/sol3/papers.cfm?abstract_id=475481. However, only forty-seven firm/year observations are
the IRRC sample using CUSIP numbers. Where this step is unsuccessful, I match using additional firm information (e.g., company name, ticker, etc.). I obtain data on institutional ownership from the Thomson Financial ownership database.\(^75\) For each firm, I generate a dichotomous variable indicating whether or not the firm has a large institutional shareholder, defined by ownership of at least 5% of the total outstanding shares. This variable, however, only accounts for shareholders that are mutual funds, pension funds, insurance companies, corporations, and other 13(f) institutions.\(^76\) Thus, the variable does not account for large individual shareholders.

While the IRRC data does not contain information for every year in the period expanded by the IRRC volumes, I nonetheless construct the panel on an annual basis. Following the prior literature, I fill in the missing governance provisions assuming that the provisions in place at time \(t\) are also in place at time \(t+1\) and all subsequent years until the next IRRC volume is available.\(^77\) Data obtained from Compustat, ExecuComp, Thomson Financial, and CRSP required no filling.

Following Bebchuk and Cohen and Bebchuk, Cohen, and Ferrell, I exclude Real Estate Investment Trusts (REITs) from the sample, given that such firms have distinct governance restrictions that provide managers with far less discretion in their operations than that afforded by other corporations.\(^78\) I also exclude firms with dual class share structures, given that the entrenchment potential of such firms may leave other governance structures superfluous.\(^79\) Also, the simple binary classification of "dual class" may make it difficult to capture the wide variation in voting and ownership structures that exist across dual class firms.\(^80\)

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\(^73\)For more information on the Thomson Financial ownership data, see the WRDS website at http://wrds.wharton.upenn.edu.


\(^75\)BEBCHUK ET AL., supra note 40, at 13; Bebchuk & Cohen, supra note 18, at 418; Gompers et al., supra note 10, at 113.

\(^76\)Bebchuk & Cohen, supra note 18, at 418; BEBCHUK ET AL., supra note 40, at 16. REITs have various legal requirements that limit managerial discretion to divert cash flows. See, e.g., Hartzell et al., The Role of the Underlying Real Asset Market in REIT IPOs, REAL ESTATE ECONOMICS, Jan. 2005, at 29.

\(^77\)See Bebchuk & Cohen, supra note 18, at 418.

\(^78\)Gompers et al., supra note 10, at 111 n.5.
B. Summary Statistics

Table I reports the means and standard deviations of the variables used throughout this analysis. Sixty-one percent of the sample have a classified board. Fifty-four percent have classified boards established in their charters. The incidence of classified boards remains quite stable across all years of the panel, ranging from 59% in 1990 to 63% in 1996. Amongst all firms in the analysis, roughly 95% did not change their classified board status over the years in which they were represented in the data. These numbers confirm the relatively time-invariant nature of this variable and motivate the instrumental variables strategy employed in Part IV.

The mean value for the adjusted GIM Index is 8.6, with a standard deviation of 2.6. For those firms with classified boards, the average adjusted GIM Index value is 9.3. Fifty-four percent of the sample are incorporated in Delaware. Average values for the other controls used in the regressions are as follows: total assets in real 2002 dollars ($10.1 billion), return-on-assets (2%), leverage (23%), fraction of shares held by insiders (3%), incidence of large institutional shareholder (76%), company age in months (296), ratio of capital expenditures to assets (6%), and three-year average growth rate in sales (14.1%).

The data exhibit significant variation in these control variables across the firms in the sample. With the exception of Delaware incorporation, the data also demonstrate notable variation within individual firms. For instance, 18% of the variation of firm assets across the sample is due to variation within firms over time. This intra-firm variation allows for more efficient estimation of the coefficients in the fixed-effects models of Part IV.

The mean value of Tobin's Q across the sample is 1.85, with a standard deviation of 1.5. The mean Q value for firms with classified boards is 1.76. Few conclusions should be drawn from this simple observation, however, given the failure to account for other factors correlated with both firm value and classified board status. Firms with classified boards are generally smaller than those without such structures ($7.4 billion in real assets versus $14.2 billion). They also have a lower average growth rate in real sales (12.7% versus 16.3%) and a lower fraction of shares held by insiders (2.8% versus 3.6%). Correlations between the covariates and Tobin's Q are as follows: real assets (-0.05), return-on-assets (0.11), leverage (-0.20), growth rate in real sales (0.16), insider ownership share (0.04), large institutional shareholder incidence (-0.06), age (-0.14), and capital expenditures-asset ratio (0.07). The correlation between Q and the adjusted GIM Index is -0.11. The average Q value for Delaware firms
is higher than that for firms incorporated elsewhere (1.94 versus 1.73). To identify an unbiased relationship between classified board status and firm value, it becomes important to control for these additional factors.

IV. FIXED EFFECTS: METHODOLOGY AND RESULTS

A. Background

I estimate regression equations of the following form:

\[ Q_{it} = \alpha + \beta_1 CB_i + \beta_2 X_{it} + \beta_3 G_{it} + \lambda_i + \tau_t + \epsilon_{it} \]

where \( i \) indexes individual firms and \( t \) indexes the year. \( Q \) is the Tobin's Q value for each firm (or its industry-adjusted value). \( CB \) is an indicator for whether or not the firm has a classified board. \( X \) is a set of covariates and their squares (e.g., firm age, firm size, etc.). \( G \) is the adjusted GIM Index value and its square. \( \lambda_i \) and \( \tau_t \) are firm-fixed effects and year-fixed effects, respectively.

This model estimates the impact of classified boards on firm value, controlling for fixed differences across firms and across years. This specification also controls for various firm specific characteristics: firm age, firm size, ratio of capital expenditures to assets, leverage ratio, return-on-assets, three-year average annual growth rate in real sales, Delaware incorporation, insider ownership share, incidence of large institutional shareholder, and the adjusted GIM Index score. To avoid the imposition of linearity, I control for levels of the continuous covariates along with their squares.

Industry effects are addressed in two ways. Primarily, I use industry-adjusted Tobin's Q as the dependent variable, in lieu of the firm's actual Tobin's Q value. To control more completely for fixed differences across industries, I also run regressions that include a full set of industry dummies.

While classified board status does change for some firms over the sample period, \( CB \) is modeled in the above equation with an index of "i" only to demonstrate its predominantly time-invariant nature. For each regression in this section, I use Huber-White sandwich estimators to form heteroskedasticity-consistent estimates of the standard errors.\(^81\)

One major limitation in the current literature is the failure to account for omitted variables that may also be correlated with a firm's classified board status. This problem is of particular concern given the endogenous nature of the classified board variable. A plausible argument can be made, however, that CB is uncorrelated with the contemporaneous error term, εt. As noted by Bebchuk and Cohen, the classified board status of most firms was largely fixed as of 1990.\(^{82}\) That is, firms that did not have classified boards in their charters by the early 1990s found it very difficult to adopt such provisions, given that shareholders over this period became reluctant to approve these structures. Thus, according to Bebchuk and Cohen, the fact that a firm does not have a classified board at any given year in the sample does not reflect a decision by management at that time period.\(^{53}\)

Moreover, for those firms with classified boards at the beginning of the 1990s, shareholders were generally powerless to remove them, given the agenda-setting powers of the board.\(^{84}\) While this argument is inapposite to those firms that incorporated during the 1990s, it does suggest that there may be little need generally to worry about correlation between CB and εt.

However, to the extent that each firm has a fixed characteristic over time for which the specification does not control and that is correlated with both classified board status and firm value, the above argument as to the predetermined nature of classified boards does not fully resolve the omitted variables concern. If this firm-fixed factor was present in 1990 and played a part in the classified board decision at that time, then the current classified board status will still be correlated with this unobservable fixed factor, leading to biased results. Bebchuk and Cohen do acknowledge this observation and attempt to control for these fixed factors by including lagged Tobin's Q as an additional right-hand side variable.\(^{85}\) However, while this approach may remove some sources of autocorrelation and address questions as to simultaneity, forming a dynamic panel by including lagged dependent variables does not remove the influence of unobservable firm-fixed effects.\(^{86}\) These omitted factors will still bias the estimated

\(^{82}\)Bebchuk & Cohen, supra note 18, at 426.

\(^{83}\)Id.

\(^{84}\)Id.

\(^{85}\)See id. at 427-28.

\(^{86}\)If correlation between classified boards and the unobservable fixed factor is resulting in biased estimates, then one must remove the unobservable factor from the specification to resolve this bias (or account for it in some other direct manner). Controlling for 1990 Tobin's Q, however, does not accomplish this removal. Essentially, controlling for this lagged Q value does not provide much new information. CB and A are present in the years following 1990 and impact firm value, and both of them were present in 1990 and impacted firm value at that time. Thus, controlling for 1990 Q does not allow one to sort out the differential effects of these two factors,
relationship between classified boards and firm value. A more direct solution is required to address correlation with firm-fixed effects. In linear specifications, such as those employed in this section, economists typically utilize "within" estimators to annihilate the unobservable fixed factors. 87

Academics have identified various unobservable fixed factors of concern in empirical governance studies. 88 One such factor is the "self-serving" inclinations of management. 89 Managerial inclinations may impact firm value in that the most self-serving managers are likely to seek significant private benefits of control. Those inclinations are also likely to be correlated with the decision of whether or not to stagger the board. Managers seeking to entrench themselves for their own gains will welcome the leeway provided by classified boards. Of course, as pointed out by Bebchuk and Cohen, management changes hands enough that one may not expect this particular factor to be fixed over time. 90 Nonetheless, whether one is considering managerial inclinations or some other omitted factor, it is certainly plausible that some element of these firm-specific characteristics do remain fixed over time. Such factors must be addressed to ensure unbiased estimates.

It is possible that the unobservable individual effect is not correlated with the right-hand side variables, in which case a fixed-effects (or "within") estimator is unnecessary and a random effects estimator would

given that (a) the two factors are correlated with each other, and (b) one of them remains unobservable. To see this another way, consider a simplified version of the model: 

\[ Q_{i,1999} = \beta_1 CB_i + \beta_2 X_{i,1990} + \lambda_i + \epsilon_{i,1999} \]  

(for the year 1999) and 

\[ Q_{i,1990} = \beta_1 CB_i + \beta_2 X_{i,1990} + \lambda_i + \epsilon_{i,1990} \]  

(for the year 1990). Now, consider accounting for \( \lambda_i \) in the 1999 specification by using \( Q_{i,1990} \) as a proxy for this measure. Since \( \lambda_i \) and \( Q_{i,1990} \) are not precisely equivalent, we impose a measurement error, \( \nu_i \), in the 1999 specification by using \( Q_{i,1999} \) as a proxy, where 

\[ \nu_i = \lambda_i - Q_{i,1990} \]  

(or \( -\beta_1 CB_i - \beta_2 X_{i,1990} - \epsilon_{i,1990} \)). The 1999 specification now looks like 

\[ Q_{i,1999} = \beta_1 CB_i + \beta_2 X_{i,1990} + Q_{i,1990} + \epsilon_{i,1999} + \nu_i \]  

Including the components of \( \nu_i \), this becomes 

\[ Q_{i,1999} = \beta_2 (X_{i,1999} - X_{i,1990}) + Q_{i,1990} + \epsilon_{i,1999} - \epsilon_{i,1990} \]  

CB is now removed from the specification. Thus, using 1990 Q values to account for the unobservable fixed factor still leaves us unable to estimate the coefficient of the classified board variable. For a general discussion regarding dynamic panel data models (i.e., those that include lagged dependent variables as covariates) and unobservable individual/firm effects, see Arellano & Bond, supra note 52; Stephen Bond, Dynamic Panel Data Models: A Guide to Micro Data Methods and Practice (Ctr. for Microdata Methods and Practice Working Paper No. CWP-09/02, 2002).

87 "Within" estimation or "fixed effects" estimation entails transforming each observation into deviations from individual/firm means. Thus, the resulting estimation makes use of within-firm variation in the data.

88 See, e.g., Himmelberg et al., supra note 31, at 357-58 (noting several sources of unobserved fixed firm heterogeneity, including the quality of the owners' monitoring technology).

89 See, e.g., Bebchuk & Cohen, supra note 18, at 427-28 (noting that an unobservable firm characteristic, such as a self-serving management, may have led both to a firm's decision to establish a classified board prior to 1990 and to a low firm value throughout the 1990s).

90 Id. at 428.
generate more efficient estimates. As a first step, I run a Hausman specification test\textsuperscript{91} to determine whether a fixed-effects model or a random effects model should be employed and to test for the presence of an individual fixed effect correlated with the regressors. Based on the results of this test, I reject the hypothesis that the random effects estimator is equivalent to the fixed-effects estimator, and thus, I reject the assumption that there is no correlation between the classified board indicator (and the other regressors) and an unobservable firm-fixed factor.

Instead of relying on pooled OLS estimates, I embrace the generous tools provided by panel data and consequently run fixed-effects specifications to control for unobserved factors. However, standard fixed-effects estimators utilize variation within individual firms and not between individual firms. That is, fixed-effects or "within" regression requires transforming the data into deviations from individual means, essentially annihilating time-invariant variables from the specification. While classified boards are not perfectly time-invariant in the sample, the vast majority of sample firms (95\%) have fixed classified board status, in which case a fixed-effects estimator significantly limits the amount of information utilized to identify the impact of classified boards.

I hesitate to abandon the fixed-effects specification, however, given the concern over unobservable factors. Thus, in order to generate unbiased estimates, I explore the instrumental variables model developed by Hausman and Taylor.\textsuperscript{92} With a clever approach, the Hausman-Taylor estimator utilizes instruments from within the specification and allows for the use of fixed-effects estimation without sacrificing identification of the coefficients of the time-invariant variables.\textsuperscript{93}

\textbf{B. Hausman-Taylor Model}

The Hausman-Taylor instrumental variables estimator entails using some of the right-hand-side variables \textit{twice}: once as instruments for themselves (i.e., as control variables), and once as instruments for the time-invariant variable (which, in the present specification, is the classified board status).\textsuperscript{94} Hausman and Taylor consider a model very much like the one specified above: $Y_{it} = \beta_1 Z_i + \beta_2 X_{1it} + \beta_3 X_{2it} + \lambda_i + \varepsilon_{it}$ ("equation 1").\textsuperscript{95} That is, Hausman and Taylor separate the right-hand side variables into

\textsuperscript{92}Hausman & Taylor, \textit{ supra} note 1, at 1378-79, 1383-89.
\textsuperscript{93}See id.
\textsuperscript{94}See id.
\textsuperscript{95}See id. at 1379.
three components: \( Z \) is a time-invariant variable that is correlated with the individual effect, \( \lambda_i \); \( X_i \) are time-varying regressors that are not correlated with \( \lambda_i \) (the "good" regressors); and \( X_2 \) are time-varying regressors that are correlated with \( \lambda_2 \). The objective is to derive a consistent estimate of \( \beta_1 \). As can be seen, simply running a fixed-effects or "within" regression on this model will annihilate \( Z \) and render estimation of \( \beta_1 \) impossible. The Hausman-Taylor model does make use of "within" estimation, but it attempts to retain as much information as possible to identify all of the desired coefficients.

The Hausman-Taylor estimator is equivalent to a two-stage least squares regression of equation 1, using \((Q, X_i, Q_vX_2, P_vX_1)\) as the set of instruments, where \( Q_v \) is the matrix \( M \) converted into deviations from individual means and \( P_vM \) is the matrix \( M \) converted into individual means. The estimator is quite intuitive. Essentially, in order to make the most out of the available information within the panel, the estimator breaks the "good" time-varying variables into two components: individual means and deviation from individual means. The deviation from individual means for the "good" variables can be used to instrument (i.e., control) for these "good" variables. The individual means are not necessary for these purposes. Thus, the estimator sets aside this "extra" information and uses it to instrument for the time-invariant variable. Assuming that certain assumptions have been met, this approach allows for consistent estimation of the coefficient of the time-invariant variable, while also controlling for the presence of the unobservable individual effect.

The Hausman-Taylor approach is more easily explained by considering a simplified, two-step variation of the above estimator. In the first step, one runs a fixed-effects or "within" regression. Eliminating \( \lambda_i \) from the specification, this allows for consistent estimation of both \( \beta_2 \) and \( \beta_3 \). With consistent estimation of these coefficients, one then transforms the model into the following: \( D_{it} = \beta_1 Z_i + \lambda_i + \epsilon_{it} \) ("equation 2"), where (i) \( D_{it} = Y_{it} - \beta_2 X_{1it} - \beta_3 X_{2it} \) and (ii) \( \beta_2 \) and \( \beta_3 \) are the estimated coefficients from this first step. Transforming the observations into individual means, equation 2 becomes: \( D_{i*} = \beta_1 Z_i + \lambda_i + \epsilon_{i*} \) ("equation 3"), where \( D_{i*} \) equals the average of the \( D_{it} \) values over all time periods (for

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96 See Hausman & Taylor, supra note 1, at 1379. Hausman and Taylor actually consider four sets of covariates, where the fourth is a set of time-invariant variables that are not correlated with the individual effects. See id. I ignore this fourth category for the purposes of this illustration.

97 An estimate of a particular parameter is said to be consistent if it converges in probability to the true parameter value as the number of observations tends towards infinity.

each individual firm i). In the second step, under the additional assumption that \( X_i \) and \( Z \) are correlated, one can now instrument \( Z_i \) with \( X_{iit} \) in equation 3 to address the correlation between \( Z \) and \( \lambda_i \) and derive consistent estimates of the coefficient of the classified board indicator.99

The Hausman-Taylor methodology enhances the estimation of the classified board coefficient by allowing one to make use of the variation in classified board status across firms, as opposed to relying on the limited variation within firms.100 All that is needed is the existence of "good" time-varying regressors—that is, regressors that are uncorrelated with the unobservable individual effects. While this condition presents less of an obstacle than that posed by ordinary instrumental variables, it is not altogether easy to find variables capable of satisfying this requirement. Fortunately, a benefit of the Hausman-Taylor model is that one is able to test the assumption that the set of variables in \( X_i \) is uncorrelated with the individual effect, \( \lambda_i \). Additionally, one may do this collectively for all of the \( X_i \) variables, as opposed to traditional specification tests, which only test over-identifying restrictions.101

Required for these purposes are factors correlated with the classified board status but uncorrelated with the unobservable firm characteristic. Many of the covariates will probably not fit this bill. Consider this unobservable fixed factor as representing the degree to which management is self-serving. Now, consider the fraction of shares held by insiders. It is reasonable to believe that insider ownership may be connected to the inclinations of management. Perhaps well-behaving managers wish to signal their good nature to the market by holding more of the firm's shares.

In general, variables that represent firm-specific outcomes will likely fail to satisfy the condition that they be uncorrelated with this firm-specific fixed effect. Ideally, one would make use of variables that capture factors out of the hands of the firm participants, such that there can be little feedback between the firm-fixed effect and these other factors. Consequently, I explore the age of the firm as the \( X_i \) variable (i.e., the "good" time-varying variable) for this Hausman-Taylor analysis.

99See Hausman & Taylor, supra note 1, at 1383. As indicated previously, \( X_{iit} \) is uncorrelated with \( \lambda_i \). See supra Part IV.B. Thus, \( X_{iit} \) is uncorrelated with \( \lambda_i \), as required to perform instrumental variables regression. A necessary condition for identification of \( \beta_i \) is that the number of columns in \( X_i \) must be at least as great as the number of columns in \( Z \). See id. In other words, the number of time-varying regressors not correlated with the individual effect must be at least as great as the number of time-invariant variables.

100Consistent estimation of the classified board coefficient under this approach will not be impacted by the small amount of time variation in classified board status within firms. The identification assumptions of the Hausman-Taylor instruments still hold.

101See Hausman & Taylor, supra note 1, at 1388-89.
It is reasonable to assume that a firm's age or experience is correlated with the classified board decision, particularly in light of the arguments made by Bebchuk and Cohen concerning the difficulties in modifying classified board status following 1990.\textsuperscript{102} The development of shareholder uneasiness concerning classified boards, combined with the agenda-setting powers provided to boards of directors, creates a specific link between classified board status and firm age.

Furthermore, the age of a firm may be uncorrelated with the unobservable fixed characteristic, allowing firm age to satisfy the necessary identification conditions under the Hausman-Taylor model. First of all, this intrinsic characteristic should not be impacted by increases in firm age. As the firm ages, this fixed characteristic, by nature, does not change. Of course, there may be reason to be concerned with cohort effects.\textsuperscript{103} Consider the firm-fixed effect as representing the self-serving inclinations of management. Different generations of firms may have different ranges of self-serving tendencies. To alleviate concerns over generational bias, I limit the regressions to those firms that entered the CRSP monthly database after 1973 (i.e., those firms under thirty years of age).\textsuperscript{104} Focusing on this time period will still capture the transformation from little resistance to significant resistance in enacting classified boards, and, thus, should still pick up meaningful correlation between age and classified board status.

Additionally, unlike the case with other covariates, such as the share of insider ownership, there is not likely to be much feedback in the other direction—from the firm-fixed factor to firm age—considering that such fixed characteristics cannot alter the course of time. Of course, feedback may work in this direction under a survivorship story.\textsuperscript{105} However, limiting the sample to those firms under thirty years of age will also alleviate any survivorship bias. Moreover, any correlation between the firm-fixed effect

\textsuperscript{102}See Bebchuk & Cohen, supra note 18, at 426 (noting shareholders' reluctance after 1990 to approve charter-based provisions creating staggered boards).

\textsuperscript{103}See Hausman & Taylor, supra note 1, at 1393.

\textsuperscript{104}The results presented below are relatively robust to this precise age cut-off (particularly with respect to the sign of the estimated coefficients). Even within this thirty-year period, the possibility remains that firms incorporated at the beginning of the period differ fundamentally in their self-serving inclinations. Nonetheless, in taking this approach, I do alleviate concerns over cohort bias driven by those firms that have been around for a significant period of time (well over the thirty-year mark). In any case, these fixed factors are meant to represent the kinds of structures or impulses that are likely to be present in all generations. Moreover, while different cohorts may face different levels of resistance to these inherent impulses, the immediate focus here is on capturing the impulse itself and not on capturing the means by which the firms can act on that impulse (which are addressed by the use of other control variables).

\textsuperscript{105}That is, if self-serving managers cause firms to have a higher probability of failure, then those firms that survive may have a distinct set of managerial inclinations.
and firm age attributable to a survivorship story would likely generate a positive bias in the estimated classified board coefficient. Thus, the fact that negative coefficients are estimated in the face of a potential positive bias only lends support to the conclusion that classified boards suppress firm value.

Table II presents the results of Hausman-Taylor estimation using firm age and its square as $X_1$ (i.e., the set of time-varying variables that are uncorrelated with the firm-fixed effect). The results largely confirm the findings of Bebchuk and Cohen. That is, I document a negative relationship between classified board status and firm value. This relationship is statistically significant in most of the specifications explored in this section. Column 1 presents results using industry-adjusted Q as the dependent variable. The estimated coefficient on the classified board variable is $-0.9$; this estimate is statistically significant at the 5% level. Running a generalized Hausman specification test, I fail to reject the assumption that firm age and its square are uncorrelated with the firm-fixed effect. Column 2 of Table II presents results using charter-based classified boards in place of the general classified board indicator. The estimated coefficient on this more limited classified board variable is $-1.0$; this estimate is significant at 5%. In column 3, I replace industry-adjusted Q with regular Tobin's Q values and control for industry effects using industry dummies. The estimated coefficient of the classified board indicator remains negative in value (-0.8) and is significant at 10%.

In various unreported regressions, I run additional specification checks on the above results. First, when I limit the sample to those firms in the post-1995 period, the estimated coefficient of the classified board indicator becomes $-1.0$ and remains significant at 5%. Second, the

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106 Attributing positive values of $\lambda$ to "good" managerial inclinations, one might expect a positive correlation between $\lambda_i$ and firm age under this survivorship story. With this positive correlation, the coefficient generated by the Hausman-Taylor model would be biased in a positive direction.


108 For the reported results, I estimate the coefficients under the instrumental variables interpretation of the Hausman-Taylor estimator. In unreported regressions, I generate efficient or "optimal" instrumental-variables estimates by estimating the variance-covariance matrix and then transforming the specification, such that the transformed error term has a scalar variance-covariance matrix. See Breusch et al., supra note 98, at 696. The reported results are robust to this exercise.

109 Hausman & Taylor, supra note 1, at 1388-89.

110 The impact of classified boards may have changed significantly following 1995 as a result of the developments in Delaware case law, which solidified the combined use of classified boards and poison pills to fight proxy contests. See, e.g., Moore Corp. Ltd. v. Wallace Computer Servs., Inc., 907 F. Supp. 1545 (D. Del. 1995). Thus, as a specification check, I run the model using a sample that is confined to these years. However, there may be little benefit in isolating
results from column 1 are virtually unchanged (1) when the "entrenchment" index and the "other-provisions" index\textsuperscript{111} are used instead of the adjusted GIM Index, and (2) when this "entrenchment" index is replaced with a set of indicator variables for each of its separate provisions. This latter specification avoids some of the limitations that inhere in an index approach and controls more flexibly for the precise components of the firm's governance portfolio. Third, using the ratio of research and development expenses to sales as a proxy for growth opportunities in place of the growth rate in sales, the estimated negative effect of classified boards on industry-adjusted Q increases to \(-1.2\); this estimate is significant at 5%. Finally, using the log of Tobin's Q to alleviate the influence of outlying observations, the estimated association between classified boards and firm value remains negative but is no longer statistically significant.

Column 5 of Table II presents results of Amemiya-MaCurdy estimation\textsuperscript{112} of the effect of classified boards on industry-adjusted Q. The Amemiya-MaCurdy model makes use of a greater range of the \(X_1\) values to identify the coefficient of the time-invariant variable, as opposed to just the individual means of the \(X_1\) variables, as is the case with the Hausman-Taylor estimator.\textsuperscript{113} This approach, however, requires a slightly stronger exogeneity assumption. Hausman and Taylor only required that the means of \(X_1\) be uncorrelated with the individual effects. In contrast, Amemiya and MaCurdy require that \(X_1\) be uncorrelated with the individual effect at every point in time.\textsuperscript{114} Of course, as Amemiya and MaCurdy discuss, the conditions of their model will hold generally in the same set of instances that the conditions of the Hausman-Taylor model will hold.\textsuperscript{115} Under the Amemiya-MaCurdy approach, the estimated coefficient of the classified board variable remains negative at \(-1.0\); however, the estimate is not significantly different from 0.\textsuperscript{116}

\textsuperscript{111}See BEBCHUK ET AL., supra note 40, at 6-11.
\textsuperscript{112}Takeki Amemiya & Thomas E. MaCurdy, Instrumental-Variable Estimation of an Error-Components Model, 54 ECONOMETRICA 869, 872-77 (1986).
\textsuperscript{113}Hausman and Taylor use \(X_1\) as two instruments: first as deviation from individual means \((Q, X_1)\), and then as individual time means \((P, X_1)\). Breusch et al., supra note 98, at 697-98. Amemiya and MaCurdy use \(X_1\) as \(T+1\) instruments: first as deviations from individual means, and then as one instrument for each of the \(T\) time periods. Id.
\textsuperscript{114}Id. at 698.
\textsuperscript{115}See Amemiya & MaCurdy, supra note 112, at 877.
\textsuperscript{116}The Amemiya-MaCurdy estimator also requires a balanced panel data set. Thus, for the purposes of this estimator, I limit the sample to those firms that are represented in all IRRC volumes.
Despite some variation in the level of significance, the negative relationship between Tobin's Q and classified boards is relatively consistent across the numerous specifications described above. To the extent that the experience of those firms under thirty years of age may generalize to the full sample, this evidence supports the results found by Bebchuk and Cohen that classified boards may lead to a reduction in firm value.\textsuperscript{117}

Column 1 of Table II also presents the estimated coefficients of the control variables used in the primary regression. The adjusted GIM Index has a negative and statistically significant effect (at 5%) on industry-adjusted Q with a coefficient of -0.2. Firm age has a negative and significant relationship with firm value, as does the leverage ratio and Delaware incorporation. Return-on-assets and the capital expenditure-asset ratio have a positive and significant effect on industry-adjusted Q. These results are relatively consistent across the range of specifications. The sales growth rate, the incidence of large institutional shareholders, and the share of insider ownership have insignificant associations with firm value.

To illustrate the impact of the Hausman-Taylor approach, in column 4, I present the results of a pooled OLS regression of industry-adjusted Q on classified board status, the set of covariates used above, and a set of year dummies, without controlling for firm-fixed effects. Similar to the findings of Bebchuk and Cohen, the estimated coefficient is negative and significant at 5%.\textsuperscript{118} The magnitude of the effect on firm value (-0.05), however, is much smaller than the Hausman-Taylor estimates above. Indeed, the Hausman-Taylor estimates are perhaps implausibly large, suggesting that the underlying model itself (see equation *** above) may be misspecified. Encouragingly, based on a generalized Hausman test, I fail to reject the identification assumptions of the Hausman-Taylor model. However, such a test only evaluates the correlation between the instruments and the omitted fixed effect. The possibility remains that other misspecifications persist in the underlying model. In any case, given the potential for omitted firm effects, one should be just as suspicious of the pooled OLS regression results. Moreover, the consistent findings throughout this article of a negative relationship between firm value and classified board status, even if implausibly large, still provide some probative evidence against the claim that classified boards increase, rather than decrease, the value of firms.

\textsuperscript{117}Bebchuk & Cohen, \textit{supra} note 18, at 420-24.
\textsuperscript{118}I\textit{d.}
V. DIFFERENCE-IN-DIFFERENCE-IN-DIFFERENCE MODEL

Motivating the above section was the desire to account for unobservable factors that could lead to bias in estimating the effect of classified boards on firm value. Panel data techniques are powerful in this regard. In this section, I take an alternative approach towards the same goal by making use of natural experiment methodologies common to the labor and public economics literature. Exploiting several sources of variation in the data, I attempt to remove the influence of certain unobservable factors by constructing a variation on the difference-in-difference-in-difference (DDD) estimator recently employed by Rauh.119

The DDD specification explored in this section utilizes one source of variation that is arguably exogenous to the model: the mid-1990s transformation in Delaware case law. During this time period, the outcome of several takeover contests involving Delaware firms (i.e., Younkers, Wallace Computer, and Circon) legitimized the ability of managers in firms with staggered boards to continue to assert poison pills after having lost the first round of a proxy contest.120 This stamp of approval strengthened the classified board/poison pill combination and opened the door for classified boards to affect firm behavior in a meaningful fashion. Following Rauh, I treat this transformation in the legal environment as a natural experiment by which to test the impact of classified boards.121 I mark 1995 as the point of transformation given that the Delaware District Court decided Wallace122 in that year, representing the only case law generated by the trilogy of takeover contests.123 More generally, the Delaware Supreme Court's 1995 Unitrin decision124 solidified the ability of managers to maintain poison pills over indefinite periods of time.125 This decision also strengthened the potential impact of classified boards, given that the power of such structures derives largely from their combined use with poison pills.126

Studies implementing natural experiment methodologies often rely on changes in legislative or regulatory positions over time. These approaches are vulnerable to legislative endogeneity concerns. The first

119 See Rauh, supra note 2, at 392-93, 399-401 (using a DDD model to estimate the impact of classified boards on employee ownership of own-company stock).
120 See Subramanian, supra note 17, at 52-53.
121 See Rauh, supra note 2, at 392-93, 399-401.
123 See Subramanian, supra note 17, at 54.
125 Bebchuk & Cohen, supra note 18, at 412.
126 Id.
source of concern is the potential for reincorporation by firms. Since reincorporations were relatively rare over the applicable time period, and since such events require stockholder approval, commentators have argued that the possibility of reincorporations pose few endogeneity concerns.\textsuperscript{127} The second source of concern is perhaps more troubling: in passing legislation or regulations, policymakers are typically acting in response to the interests of those parties whose behavior is being studied. This creates a link between firm/individual characteristics and the legislative environment, leading to potential biases. Changes in legal environments attributable to the judiciary, such as those utilized in this article, are less prone to this source of bias. In comparison with legislation, judicial decisions are arguably less influenced by the specific interests of the managers and firms represented in our sample.

I estimate the following DDD model:

$$Q_{it} = \alpha + \beta_1 X_{it} + \beta_2 CB_{it} + \beta_3 \text{DEL}_{it} + \beta_4 \text{POST}_{it} + \beta_5 (\text{POST}_{it} \ast \text{DEL}_{it}) + \beta_6 (\text{POST}_{it} \ast \text{CB}_{it}) + \beta_7 (\text{CB}_{it} \ast \text{DEL}_{it}) + \beta_8 (\text{POST}_{it} \ast \text{DEL}_{it} \ast \text{CB}_{it}) + \tau_i + \alpha_{state} + \alpha_{industry} + \epsilon_{it}$$

where $X$ is a set of covariates (adjusted GIM Index, firm age, firm size, capital expenditures-asset ratio, leverage, return-on-assets, sales growth rate, insider ownership share, and the incidence of large institutional shareholders). CB is the predicted likelihood of having a classified board. DEL is an indicator for Delaware incorporation. POST is an indicator for years after 1995.

The coefficient of interest is $\beta_8$. It captures the isolated effect of classified boards on firm value.\textsuperscript{128} It differs from the DDD coefficient estimated by Rauh\textsuperscript{129} in that, instead of taking the difference between firms

\textsuperscript{127}See, e.g., Subramanian, supra note 8, at 668.

\textsuperscript{128}The magnitude of the triple-differences coefficient, $\beta_8$, may not capture the precise magnitude of the effect of classified boards on firm value (in an absolute sense), given that (a) classified boards may still have had some effect on firm value, albeit a lesser effect, prior to the transformation in Delaware law, and (b) firms incorporated in states other than Delaware may still have responded to the transformation in Delaware law, albeit to a lesser extent. Nonetheless, given that the effect of classified boards should be more pronounced in the post-1995 period and given that the effect of the mid-1990s transformation in Delaware law should be more pronounced for firms incorporated in Delaware, the triple-differences coefficient should still capture the sign of the relationship between classified boards and firm value and, thus, should still provide insight on the question of whether classified boards enhance or suppress firm value. In regard to magnitude, the coefficient should indicate the differential effect on firm value of the mid-1990s validation of the classified board/poison pill combination for Delaware firms relative to non-Delaware firms.

\textsuperscript{129}See Rauh, supra note 2, at 399-400.
with and without classified boards, I form the predicted likelihood of having a classified board and then interact this variable with the POST * DEL interaction. The following discussion offers a more thorough illustration of the mechanics behind this approach.

I begin with the consideration of a simple difference-in-difference (DD) model which is estimated as follows: (1) calculate the difference in Tobin's Q values between firms with classified boards and firms without classified boards, and then (2) calculate the difference in this difference between observations in the post-1995 period and observations in the pre-1995 period. Essentially, whatever the initial difference was between the Q values of firms with and without classified boards, one should expect this difference to magnify (or contract) in the post-1995 period, given that the transformation in Delaware law only affected the "treatment" group (i.e., those with classified boards). Given Bebchuk and Cohen's arguments as to the difficulty in modifying classified board status throughout the 1990s, those firms with classified boards can be viewed as the "treatment" group for this natural experiment with little concern as to firms modifying classified board status in response to the change in law. This simple approach eliminates the effects of certain unobservable factors. The first layer of differences (pre- and post-1995) removes any fixed differences in firm value between firms with and without classified boards. The second layer of differences (classified board/lack of classified board) removes the effects of any aggregate economic shocks that hit the system in the post-1995 period. However, this DD model does not identify the true effect of classified boards to the extent that there are any shocks specific to firms with classified boards in the post-1995 period. In this case, the DD calculation may be picking up the effect of these unobservable shocks instead of the effect of classified boards.

To eliminate the effects of these CB/POST-specific shocks, I construct the DDD estimator specified above and consider an additional level of variation: Delaware incorporation versus non-Delaware incorporation. For two reasons, one should expect classified boards to have more of an impact, in the post-1995 period relative to the pre-1995 period, on firms incorporated in Delaware than on firms incorporated elsewhere. First and foremost, Delaware's 1995 transformation directly

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130 See Bebchuk & Cohen, supra note 18, at 426.
131 See Rauh, supra note 2, at 381.
132 The general DDD methodology is motivated by Gruber. Jonathan H. Gruber, The Incidence of Mandated Maternity Benefits, 84 AM. ECON. REV. 622, 627 (1994). The specific DDD model employed in this paper is motivated by Rauh. See Rauh, supra note 2, at 399.
affected firms covered under Delaware corporate law. Second, considering that the entrenchment power of classified boards derives primarily from their combined use with poison pills (i.e., the two provisions together significantly inhibit any attempt at engaging in a proxy contest to remove management), Delaware firms may experience a relatively greater impact from the validation of the classified board/poison pill combination since Delaware authorized relatively potent poison pills (though not the most potent pills) throughout the sample period. Consequently, taking this third layer of difference (Delaware incorporation versus non-Delaware incorporation) should isolate the impact of classified boards, under the assumption that any shocks specific to the CB/POST firms do not differ across Delaware and non-Delaware incorporation. That is, the model is identified, unless there are factors that are specific to that set of observations that (1) have a classified board, (2) are incorporated in Delaware, and (3) are in the post-1995 period. In addition to accounting for any effects specific to classified board firms in the post-1995 period (see \( \beta_6 \)), the DDD specification also accounts for factors specific to (1) Delaware firms in the post-1995 period (\( \beta_3 \)), and (2) Delaware firms with classified boards (\( \beta_7 \)).

The only concern at this point is the endogenous relationship between classified boards and the state of incorporation. If classified board status is endogenous to the state of incorporation, then there are likely to be factors specific to Delaware firms with classified boards for which the specification does not control and that may also be driving firm value. Of course, when one takes the difference before and after 1995, the effect of such factors may be removed. Any such endogenous relationship, however, may increase the chance that other shocks hit Delaware firms with classified boards contemporaneously with the transformation in Delaware

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133Of course, firms incorporated in other states may still feel an impact from Delaware's decision, given the assumption that their courts may follow Delaware's lead. However, the uncertainty inherent in this assumption should lead to a weaker impact in comparison with Delaware-incorporated firms.

134Certain states, however, authorize poison pills that are more potent than those in Delaware, including Pennsylvania, Maryland, Virginia, and Georgia. Subramanian, supra note 8, at 625-29. Nonetheless, considering that the transformation in Delaware law directly impacted Delaware firms and that the Delaware pills are stronger than those for many states, the crude difference between Delaware firms and non-Delaware firms should still pick up a difference in the likely impact of Delaware's validation of the classified board/poison pill combination. Moreover, in alternative specifications, I run the same DDD model, but instead of looking across Delaware and non-Delaware firms, I take the difference across two groups: one consisting of firms incorporated in Delaware, Virginia, Georgia, Maryland, and Pennsylvania and the other group consisting of firms incorporated in the remaining states. As discussed later in Part V, the results are robust to this alternative specification.
law, thereby creating a bias in the DDD estimate. To address this concern, instead of using a firm's actual classified board status, I follow Gruber and Mullainathan and use each firm's predicted likelihood of having a classified board.\textsuperscript{135} I generate these predictions using the coefficients derived from annual regressions of classified board status on a set of observable predictors. I form the applicable set of predictor variables using the covariates employed in the primary DDD specification. I then interact the predicted likelihood with an indicator for Delaware incorporation and an indicator for the post-1995 period. The coefficient of this interaction, $\beta_8$, is our measure of interest.

In Table III, I present the results of the DDD model specified above. In each regression, I use regular Tobin's Q values as the dependent variable and control for industry effects by including a full set of industry dummies. In each regression, I also run year-fixed effects and state-fixed effects and include controls for firm age, firm size, return-on-assets, insider ownership share, capital expenditures-assets ratio, leverage, sales growth rate, and the incidence of large institutional shareholders. As reported in column 1, I estimate a triple-differences coefficient, $\beta_8$, of -0.47; this estimate is statistically significant at 1%. Thus, I find that the mid-1990s change in Delaware law had a negative effect on firm value for those Delaware firms likely to have classified boards. This coefficient can be interpreted as a negative effect of classified board status on firm value. These results remain unchanged when I replace the adjusted GIM Index with the two sub-indices: the "entrenchment" index and the "other-provisions" index. The same is true when I replace this "entrenchment" index score with a set of dummy variables capturing each of the separate provisions of this subindex.

For the regressions reported in columns 1 and 3-5 of Table III, I use Huber-White sandwich estimators to form heteroskedasticity-consistent estimates of the standard errors.\textsuperscript{136} In column 2 of Table III, I run the same regression from column 1, but I make a standard-error correction to allow for an arbitrary variance-covariance matrix within each state of incorporation.\textsuperscript{137} With this correction, the estimated triple-differences

\textsuperscript{135}See Jonathan H. Gruber & Sendhil Mullainathan, \textit{Do Cigarette Taxes Make Smokers Happier?}, 5 ADVANCES IN ECON. ANALYSIS & POLY 1, 2 (2005) (using predicted smoking likelihoods, as opposed to actual smoking status, in order to alleviate endogeneity between smoking status and excise tax rates).

\textsuperscript{136}See Bebchuk & Cohen, \textit{supra} note 18, at 418.

\textsuperscript{137}Estimating grouped error terms by clustering the standard errors in this manner may be advised, given that the unit of observation in the data (i.e., firm/year) is more detailed than one of the levels of variation exploited in the specification (i.e., incorporation in Delaware). See Marianne Bertrand et al., \textit{How Much Should We Trust Differences-in-Differences Estimates?}, 119 Q.J. ECON. 249, 254, 270-72 (2004). The results reported in column 1 of Table III are also robust
Coefficient of -0.47 continues to be statistically significant but only at the 10% level. Column 3 modifies the specification from column 1 to replace the general classified board variable with an indicator variable for charter-based classified boards (that is, the specification uses the predicted likelihood of having a charter-based classified board). The estimated $\beta_8$ coefficient is -0.37 (significant at 5%). The above findings are also robust to the use of actual classified board status, as opposed to the predicted likelihood of having a classified board. Making this substitution, I estimate $\beta_8$ at -0.28 (significant at 1%), as reported in column 4.

In column 5 of Table III, the sample is limited to those firms that went public prior to the mid-1990s transformation in Delaware law. As discussed above, those firms with classified boards can be viewed as the "treatment" group with little concern as to modification of classified board status in response to the change in Delaware law. However, some concern still remains with respect to those firms that had initial public offerings (IPOs) in the years following this legal transformation. To alleviate any bias caused by such firms, I remove from the model any firm that went public in the years following this transformation. The estimated triple-differences coefficient under this approach remains similar in magnitude at -0.39 and is statistically significant at the 5% level. Thus, the above DDD findings appear to be robust to concerns over the invalidation of the "treatment" group due to IPOs in the sample period.

In an unreported regression, I also explore an alternative DDD model where, instead of looking at differences between firms incorporated in Delaware and firms incorporated outside of Delaware, I take the difference between firms incorporated in the group of states consisting of Delaware, Pennsylvania, Virginia, Maryland, and Georgia and firms incorporated elsewhere. This particular group of states authorizes the most potent poison pills and should thus feel the strongest impact of classified boards on firm value. This alternative model assumes that these additional states will still be affected, to some extent, by the transformation in Delaware case law. The estimated triple-differences coefficient under this approach is also negative (-0.42) and significant at the 5% level.

to the clustering of the standard errors on the incidence of Delaware incorporation (as opposed to clustering on each state of incorporation). However, these latter results should be interpreted with caution because clustering techniques perform poorly when the number of groups used is small. See id. at 269.

138 See Rauh, supra note 2, at 381.
139 See Subramanian, supra note 8, at 628.
Finally, in additional unreported regressions, I test the robustness of the above findings to certain functional form assumptions. First, using industry-adjusted Tobin's Q in place of regular Q values and a set of industry dummies, the triple-differences coefficient is estimated at -0.31; this estimate is significant at 10%. In any case, regular Q values and a set of industry dummies, as used above, allow for a more complete control of fixed differences across industries. Second, to alleviate the influence of outliers, I also run the above model using the log of Tobin's Q as the dependent variable and find a triple differences coefficient of -0.14 (significant at 5%). Third, when I modify the set of control variables to include the squares of all continuous variables (as opposed to just the levels of such variables), I estimate the triple-differences coefficient at -0.39; this estimate is significant at 5%.

The results derived from DDD estimation present relatively consistent evidence of a negative and economically meaningful association between classified boards and firm value. While, as in Part IV.B., it could be argued that the estimated results are implausibly large,\textsuperscript{140} the results, nonetheless, lend further support against any claim that classified boards enhance firm value. One question that remains, however, is whether this negative association is homogenous across all levels of firm value or whether this relationship differs along the distribution of value.

VI. QUANTILE REGRESSIONS

As a final robustness exercise, I run quantile regressions of industry-adjusted Tobin's Q on classified board status. Originally proposed by Koenker and Basset, quantile regressions provide estimates that are robust to some of the functional form assumptions typically taken in other estimation models.\textsuperscript{141} Particularly, quantile regressions allow relaxation of the assumption of a normally distributed error process and are less sensitive to outlying observations.\textsuperscript{142} Consequently, such regressions provide a welcome robustness check to the estimation procedures explored in Parts IV and V above.

An additional advantage of the quantile regression model for the purposes of this study is that it provides a method of modeling heterogeneous effects of a variable along different points of an outcome distribution. That is, the approach allows for estimation of the association

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\textsuperscript{140} The estimated triple-differences coefficient of -0.47 is substantial in relation to the average Tobin's Q value of 1.85.

\textsuperscript{141} Koenker & Basset, supra note 24, at 33-34, 38.

\textsuperscript{142} See id.
between classified boards and firm value, not only in the middle of the distribution of firm value, but also at the tails. Conventional regression models generate mean results that assume the regressors affect the dependent variable to the same extent across all points of the outcome distribution. Estimating mean effects may often generate useful information; however, in many cases, it is far more interesting to tell a more complete story by estimating the full distributional impact. In the present governance study, estimating the effects of classified boards throughout the distribution of firm values may offer some additional insight into the question of whether classified boards are wholly negative for firms, or whether such devices generate some positive contributions to firm value.

Based on the above findings, it may be safe to assume that classified boards depress firm value on average. It remains possible, however, that this negative association does not arise in those firms with low levels of Tobin's Q. One element to the relationship between classified boards and firm value which is likely to vary across the distribution of Tobin's Q is the probability of takeover. That is, firms with low Q values may be more likely takeover targets. This possibility, combined with the notion that classified boards provide directors with the power to reject and defend value-enhancing acquisition attempts, may support the general finding of a negative relationship between classified boards and firm value. However, managers of low-value firms may also face a stronger short-term bias as a result of the greater likelihood of takeover. In such firms, classified boards may be beneficial to firm value to the extent the insulation afforded by such devices allows managers and directors to focus on long-term strategy. These considerations may not arise to the same degree in firms on the upper tail of the value distribution where takeover odds are not as high.

Therefore, to the extent the quantile regression results illustrate a stronger negative effect at the upper tail of the value distribution in relation

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144Chernozhukov & Hansen, supra note 143, at 2.

145See id. For instance, in an evaluation of the effect of health outcomes on income, policymakers may be far more interested in looking at the effect at the low end of the income distribution.

146See, e.g., Randall Morck et al., Alternative Mechanisms for Corporate Control, 79 Am. Econ. Rev., 842, 850 (1989) (finding evidence that low industry-adjusted Tobin's Q values are associated with a higher probability of hostile takeovers).

147See Bebchuk et al., supra note 40, at 890.

148See, e.g., Bebchuk, supra note 3, at 1011.
to the lower tail of such distribution, it could be argued that, even though classified boards reduce firm value on average, these devices, nonetheless, generate offsetting benefits to shareholders in certain circumstances.

To explore these considerations, I estimate the following linear conditional quantile function:

\[ M_{QX, CB}(m) = \beta_1(m)X + \beta_2(m)CB \]

where \( M_{QX, CB}(m) \) is the \( m \)-th quantile of Tobin's \( Q \), conditional on the set of covariates. The scalar \( m \) takes on values between 0 and 1. Essentially, the \( m \)-th quantile of a variable \( Z \) is the level of \( Z \) representing the \( m \)-th percentile (i.e., the probability that \( Z \) is less than the \( m \)-th quantile equals \( m \)).

Using quantile regressions, I estimate \( \beta_2(m) \), which captures the effect of classified boards on the \( m \)-th quantile of firm value. In essence, this coefficient can be interpreted as the effect of having a classified board on firm value for those firms at the \( m \)-th percentile of the value distribution. Estimating these coefficients for a range of \( m \) values allows for a more complete understanding of the impact of classified boards.

Quantile regressions estimate \( \beta(m) = (\beta_1(m), \beta_2(m)) \) by solving the following minimization problem:

\[
\text{Min} \quad \sum m |Q_i - X_i \beta(m)| + \sum (1 - m) |Q_i - X_i \beta(m)| \\
\beta^*(m) \quad i : Q_i \geq X_i \beta \quad i : Q_i < X_i \beta
\]

where \( \beta^*(m) \) is the estimate of \( \beta(m) \), and \( X \) is the full set of regressors (\( X_1 \) and \( CB \)). In taking this approach, some of the virtues of the other specifications explored in this article must be sacrificed. Mainly, I cannot run quantile regressions while also employing the instrumental variables strategies of Part IV in an effort to account for firm-fixed effects. While econometricians have developed instrumental variables quantile regression models, the instrumental variables methods employed in Part IV of this article require the transformation of certain variables into deviations from their mean values. "Differencing" approaches of this nature cannot be

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149 More formally, the \( m \)-th quantile of a variable \( Z \) is defined as follows: \( M_Z(m) = \inf \{ z : F_Z(z) \geq m \} \), where \( F_Z(z) \) is the distribution function for \( Z \). See, e.g., Chaudhuri et al., supra note 24, at 715.


151 See Koenker & Bassett, supra note 24, at 38.

152 See, e.g., CHERNOZHUKOV & HANSEN, supra note 143, at 3-8.
employed in connection with nonlinear estimation procedures such as quantile regressions.\(^{153}\)

Despite limitations in the ability to control for unobservable factors, this quantile regression exercise should be of some analytical value given the ability to explore distributional implications. Row 1 of Table IV presents the results of quantile regressions of classified board status on industry-adjusted Tobin's Q. In each regression, I control for the same set of covariates employed in Part IV. I also control for year-fixed effects. The coefficients reported in columns 1 through 9 indicate the estimated effect of classified boards on the 0.1 through 0.9 conditional quantiles of industry-adjusted Tobin's Q, respectively. The results suggest that classified boards have no effect, or a slightly positive effect, on firm value at the lower tail of the distribution of firm value. The effect becomes more negative as one moves up this value distribution. At the upper quantiles of firm value, classified boards have a strong negative effect on firm value.

Using the general classified board variable, the estimated effect of classified boards on the 0.1 and 0.2 quantiles is 0.04 and 0.03, respectively (significant at 5% and 1%, respectively). The estimated effect falls and turns negative at the 0.5 quantile. The estimate does not reach a statistically significant negative value until the 0.6 quantile. At the 0.9 quantile, I estimate the coefficient on classified board status at -0.18; this estimate is significant at 1%. For purposes of comparison, column 4 of Table II presents the results of a pooled OLS regression of classified board status on industry-adjusted Q, using an identical set of controls. As indicated above, I estimate a mean effect of classified boards of -0.05.

This pattern does not change substantially when the general classified board variable is replaced with an indicator for the presence of a charter-based classified board, as indicated by row 2 of Table V. The estimated coefficient on classified board status decreases monotonically as one ascends the distribution of firm value. However, for the 0.1 to 0.4 quantiles, the estimate is not significantly different from 0. For the 0.5 to 0.9 quantiles, the estimated effect of classified boards is negative and statistically significant (at the 5% level for the 0.5-0.6 quantiles and at the 1% level for the 0.7-0.9 quantiles).

These results are robust to the following alternative treatments of the remaining governance provisions: (1) controlling for the "entrenchment"

\(^{153}\)Furthermore, estimation techniques in nonlinear models that attempt to account for individual/firm heterogeneity by directly estimating the individual/firm effects (as opposed to "differencing" out such effects) may result in inconsistent estimates, due to the "incidental parameters problem." See, e.g., Tony Lancaster, The Incidental Parameter Problem Since 1948, 95 J. ECONOMETRICS 391, 394 (2000).
and "other-provisions" indices (and their squares) and (2) controlling for the "other-provisions" index (and its square), while controlling separately for the components of the "entrenchment" index. Moreover, these distributional findings persist when the sample is limited to those observations in the post-1995 period.\footnote{Moreover, I also explore a specification that replaces industry-adjusted Tobin's $Q$ with regular $Q$ values and a full set of industry dummies. The estimator, however, does not converge for all quantiles under this approach. Nonetheless, for those estimators that do converge, the findings correspond with the results observed above.}

Offering the benefits of a robust estimation technique that allows for the relaxation of certain functional form assumptions, these results generally support the findings presented in Parts IV and V that classified boards suppress firm value.\footnote{See supra Parts IV & V.} The results derived from quantile regressions, however, suggest that these negative effects are concentrated along the upper portion of the distribution of firm value. At low levels of Tobin's $Q$, the association between classified boards and firm value is either negligible or positive in sign, supporting the possibility that classified boards, in certain circumstances, actually generate countervailing benefits to shareholder wealth.

However, the estimated effects of classified boards on the 0.1-0.2 quantiles may not be due to an offsetting of certain positive and negative forces as hypothesized above. Rather, the fact that little or no association between classified boards and firm value is estimated at low Tobin's $Q$ levels could simply arise from the possibility that firms with low Tobin's $Q$ face greater pressure to behave properly. With weak investment opportunities, managers and directors may have less leeway to extract private benefits and engage in other destructive behaviors. Consequently, the entrenchment afforded by classified boards may be immaterial at the low tail of the value distribution. However, considering the various ways in which classified boards may affect firms likely to be taken over, such as firms with low Tobin's $Q$ values, it is more likely that the lack of an estimated association between classified boards and firm value for low-$Q$ firms results from an aggregation of both positive and negative forces. This supposition is supported by the fact that positive coefficients are estimated in some of the low-quantile specifications explored above.

Accordingly, while the positive effects of classified boards are, on average, overwhelmed by the negative, the quantile regression results, nonetheless, suggest that some positive forces do exist and that such forces may even outweigh the negative in certain circumstances. That is, the
general findings do not necessarily support a complete rejection of the notion that classified boards benefit shareholders.

VI. CONCLUSION

Sound empirical analysis is required to shed light on the ambiguities that pervade corporate governance theory. The motivation of this article has been to draw on numerous econometric tools to address some of the empirical obstacles facing the corporate governance literature. One of the major limitations of recent studies has been the inability to account for unobservable factors that are correlated with both firm valuation and corporate governance structures. Omissions of this nature leave empirical specifications vulnerable to potentially large estimation biases.

In this study, I have embraced the panel nature of the IRRC dataset in an attempt to explore the relationship between classified boards and firm value, while accounting for unobservable fixed characteristics of firms. To address the dilemma posed by running a fixed effects model on a virtually time-invariant variable (i.e., classified board status), I have employed the Hausman-Taylor instrumental variables estimator. The results of this exercise generally support the findings of Bebchuk and Cohen that classified boards are associated with a reduction in Tobin's Q.

While these approaches may account for potential correlation between classified boards and unobservable fixed characteristics of firms, the estimates may still be biased to the extent that classified boards or any of the covariates are correlated with contemporaneous shocks to the system. The generous tools provided by panel data may be beneficial in this regard as well. In future work, I hope to structure a dynamic panel data model that synthesizes the Hausman-Taylor model and the Generalized Methods of Moments estimator employed by Arellano and Bond to control more comprehensively for correlations between the regressors and unobservable fixed factors, and between the regressors and the contemporaneous error term.

Taking an alternative "differencing" approach to the removal of unobservable factors, I have treated Delaware's mid-1990s validation of the classified board/poison pill combination as a natural experiment by which to construct a difference-in-difference-in-difference (DDD) estimator. The results derived from this DDD specification again suggest that classified boards may be associated with a reduction in firm value. Moreover,

156Hausman & Taylor, supra note 1, at 1384-89.
158Arellano & Bond, supra note 52, at 278-81.
employing quantile regressions to explore the possibility of heterogeneous impacts of classified boards along the full distribution of firm value, I have found that this negative relationship is likely concentrated along the upper tail of the value distribution. For the least valuable firms, I have documented no association, or perhaps a slight positive association, between classified boards and firm value. Such results suggest that classified boards may generate countervailing benefits in certain circumstances, despite leading to a reduction in value on average.

While the negative findings reported in Parts IV and V are robust to a range of specification checks and lend support against any argument that classified boards, on average, enhance firm value, the magnitude of the estimated coefficients are at a level that may be difficult for many to believe. Any doubts as to the plausibility of such magnitudes, however, should not lead one to place unbridled confidence in the findings reported in the literature to date. For the reasons set forth in this article, the estimates generated by such studies remain vulnerable to numerous omitted variables concerns. To provide policymakers, boards of directors, and investors with proper, unbiased guidance, the corporate governance literature must continue to address sources of misspecification in its underlying models, as I have attempted to do in this article. Only with convincing empirical methodologies can we generate the tools with which decisionmakers can structure effective, value-maximizing corporate governance regimes.
### TABLE I: SUMMARY STATISTICS

This table presents the means and standard deviations (in parentheses) of the variables used throughout this article. Tobin's Q equals the ratio of the market value of assets to the book value of assets (where the market value of assets is determined according to the formula specified in Part III). GIM Index equals the Gompers, Ishii, and Metrick governance index, adjusted to exclude classified boards. Firm age is determined according to the first time that the firm appeared in the CRSP monthly data. ROA equals the ratio of income to total assets. CapEx-Assets Ratio equals the ratio of capital expenditures to total assets. Growth Rate equals the average annual growth rate in real sales over the most recent three fiscal years. Leverage equals the ratio of long-term debt plus debt due in one year to total assets. Insider % equals the percentage of the firm's common stock held by directors and officers. Large Institutional SH is an indicator for the presence of an institutional shareholder holding at least five percent of the firm's outstanding common stock.

<table>
<thead>
<tr>
<th>Variable</th>
<th>Mean</th>
<th>Standard Deviation</th>
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</thead>
<tbody>
<tr>
<td>Tobin's Q</td>
<td>1.85</td>
<td>(1.51)</td>
</tr>
<tr>
<td>Classified Board Incidence</td>
<td>0.61</td>
<td>(0.49)</td>
</tr>
<tr>
<td>Charter-Based Classified Board Incidence</td>
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<td>(0.50)</td>
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<td>GIM Index</td>
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<tr>
<td>Firm Age (months)</td>
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<td>ROA</td>
<td>0.02</td>
<td>(0.19)</td>
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<tr>
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### TABLE II: HAUSMAN-TAYLOR ESTIMATES

Column 1 presents the results of Hausman-Taylor regressions of industry-adjusted Tobin's Q on classified board status, as well as on firm financial and governance controls, using the following set of instruments: \( Q_i X_1, Q_i X_2, P_i X_4 \), where \( X_1 = (\text{firm age, firm age squared}) \), \( X_2 \) is the set of remaining control variables, \( Q_i Z \) is the matrix \( Z \) converted into deviations from individual means, and \( P_i Z \) is the matrix \( Z \) converted into individual means. Column 2 uses charter-based classified boards in place of the general classified board indicator. Column 3 replaces industry-adjusted \( Q \) with regular Tobin's \( Q \) and a full set of industry dummies. Column 4 presents the results of pooled OLS estimation. Column 5 presents the results of the Amemiya-MacCurdy estimator. Columns 1-3 and 5 limit the sample to those firms incorporated after 1973. Each regression controls for year-fixed effects. Robust Huber-White standard errors are reported in parentheses. Significance levels are indicated by *, **, and *** for 10%, 5%, and 1%, respectively.

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TABLE III: DIFFERENCE-IN-DIFFERENCE-IN-DIFFERENCE ESTIMATES

This table presents the results of difference-in-difference-in-difference (DDD) regressions. The coefficient of the interaction term \( CB * DE * Post-1995 \) captures the isolated relationship between classified boards and Tobin's Q. \( DE \) refers to Delaware incorporation. Post-1995 is an indicator for the post-1995 period. \( CB \) refers to the predicted likelihood of having a classified board (formed using the coefficients derived from annual regressions of classified board status on a set of observable predictors). Column 4 replaces the predicted classified board likelihood with actual classified board status. Column 3 uses charter-based classified boards in place of the general classified board indicator. Column 5 limits the sample to those firms that went public prior to 1995. Each regression accounts for state, year and industry fixed effects and includes a set of financial and other controls. Standard errors are reported in parentheses. In columns 1 and 3-5, Robust Huber-White standard errors are reported. In column 2, standard errors are clustered by state of incorporation. Statistical significance is indicated by *, **, and *** for 10%, 5%, and 1%, respectively.

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TABLE IV: QUANTILE REGRESSIONS
This table presents the results of quantile regressions of industry-adjusted Tobin's Q on classified board status, as well as on various firm financial and governance controls. Each regression controls for year effects, Delaware incorporation, and the incidence of a large institutional shareholder, along with firm age, firm assets, leverage, capital expenditures-asset ratio, sales growth rate, return-on-assets, share of insider ownership, adjusted GIM Index, and their squares. Each column reports coefficients representing the effect of classified boards on the 0.1-0.9 quantiles of industry-adjusted Tobin's Q. Row 1 presents results using the general classified board indicator. Row 2 presents results using charter-based classified boards. Standard errors are reported in parentheses. Statistical significance is indicated by *, **, and *** for 10%, 5%, and 1%, respectively.

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