Extreme Governance:
An Analysis of Dual-Class Firms in the United States

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ABSTRACT

We construct and analyze a comprehensive list of dual-class firms in the United States and use this list to investigate the relationship between insider ownership and firm value. Our data has two useful features for this valuation analysis. First, since dual-class stock separates cash-flow rights from voting rights, we can separately identify the impact of each. Second, we address endogeneity concerns by using exogeneous predictors of dual-class status as instruments. While other data sets have provided one of these features, our data set is the first to provide both. We find robust evidence that firm value is increasing in insiders' cash-flow rights.

1. Introduction

In recent years, researchers have demonstrated the powerful role of shareholder rights. The stream of research on this topic finds that the strength of shareholder rights at a company is associated with stock returns, valuations, operating performance, the frequency of value-destroying mergers, and pay-for-performance sensitivity.[1] While debates exist on the proper measurement of shareholder rights, virtually all authors consider anti-takeover provisions to be among the most important restrictions of these rights.[2] Nevertheless, all of these recent papers have ignored the most extreme example of anti-takeover protection: dual-class stock. About six percent of the publicly-traded companies in the United States have more than one class of common stock, and these companies are virtually immune to a hostile takeover. In this paper, we construct the first comprehensive panel of these companies and analyze the determinants and valuation effects of dual-class status and ownership structure.[3]

In the typical dual-class company, there is a publicly traded "inferior" class of stock with one vote per share and a non-publicly traded "superior" class of stock with ten votes per share. The superior class is usually owned mostly by the insiders (managers and directors) of the firm and causes a significant wedge between their voting and cash-flow rights. In many cases, this wedge is sufficient to provide insiders with a majority of the votes despite their claims to only a minority of the economic value. The other forms

[1] Gompers, Ishii, and Metrick (2003), Bebchuk, Cohen, and Farrell (2004), Chi (2005), Core, Guay and Rusticus (2006), Fahlenbrach (2005), and Masulis, Wang, and Xie (2006).
[2] In addition to the papers cited above, see Bebchuk, Coates, and Subramanian (2002), Gillan, Hartzell, and Starks (2003), Bebchuk and Cohen (2005), and Cremers and Nair (2005).
[3] Some firms have more than two classes of common stock. To keep with the traditions of this literature, we refer to all multi-class firms as "dual class."

of anti-takeover protection – poison pills, staggered boards, golden parachutes – are no match for the power of dual-class stock.

Past studies of dual-class stock in the United States have focused on subsets of dual-class firms at the time of the creation of their inferior class,[4] the time immediately following their IPO,[5] the atypical cases where both share classes trade,[6] or as a subset of "family-controlled" firms.[7] There are no papers analyzing a panel of *all* dual-class firms – perhaps because the identification of these firms is highly labor intensive and has only become feasible with the recent availability of electronic documents from the SEC.[8] In this paper, we fill this gap in the literature by identifying and analyzing a comprehensive list of dual-class companies from 1995 to 2002.

The paper has three sets of analyses, discussed respectively in Sections 2, 3, and 4 of the paper. First, from the SEC disclosures, we code the insider holdings for all dual-class firms for both classes of stock. From these data, we can calculate the cash-flow rights and voting rights for all insiders, and summarize the ownership structure at dual-class firms. Second, by using this comprehensive list of firms – both dual-class and single-class – we can analyze the determinants of dual-class status, and provide insight into why firms choose this extreme form of governance. Finally, we exploit the difference in insiders' cash-flow rights and voting rights to shed light on the relationship between insider ownership and firm valuation. The first two sets of analyses are

[4] Partch (1987), Jarrell and Poulsen (1988), Cornett and Vetsuypens (1989), Amoako-Adu and Smith (2001), Hauser and Lauterbach (2004), Dimitrov and Jain (2006)
[5] Smart and Zutter (2003), Smart et al. (2006).
[6] Lease, McConnell, and Mikkelson (1983, 1984), Zingales (1995), Nenova (2003)
[7] Amit and Villalonga (2006).
[8] To our knowledge, the only other paper making an attempt at a comprehensive sample is the unpublished work of Zhang (2003). Zhang's paper has a somewhat different focus than ours. He uses a single cross-section in 1995 (slightly smaller than our 1995 sample) and does not analyze the determinants of dual-class status nor attempt to control for self-selection or endogeneity in the valuation analysis.

essentially descriptive and confirmative of previous papers. The findings of these descriptive studies are crucial inputs for the valuation analysis, which we consider to be the main contribution of the paper. The details of all these analyses are discussed below in more detail.

In Section 2, we find that about six percent of all Compustat firms are dual-class, comprising about eight percent of the market capitalization of all firms. For about 85 percent of dual-class firms, there is at least one untraded class of common stock, and this untraded class almost always has superior voting rights to the traded class. As pointed out by Smart and Zutter (2003) using a sample of dual-class IPOs, the most common structure is for superior shares to have ten votes per share, while inferior shares have one vote per share. We confirm this finding in our comprehensive sample. On average, we find that insiders have approximately 60 percent of the voting rights and 40 percent of the cash-flow rights in dual-class firms. For almost 40 percent of the dual-class firms, insiders have more than half of the voting rights (thus providing effective control) but less than half of the cash-flow rights. To reflect the stark separation of economic ownership and voting control in these firms, we denote them as the "separation sample" and give them special attention in the later analysis.

In Section 3, we investigate the determinants of dual-class status. These tests follow a line of research begun by DeAngelo and DeAngelo (1985), with contributions by Lehn et al. (1990), Taylor and Whittred (1998), Field (1999), Amoako-Adu and Smith (2001), and Smart and Zutter (2003). Building upon the work of these authors, we extend the sample of firms and use the results as the first-stage of the valuation analysis that follows. Most dual-class firms chose their structure prior to their IPO. For the pre-IPO owners, the dual-class structure is likely to have both benefits and costs. These benefits, typically called "private benefits of control" in the literature, can include both pecuniary and non-pecuniary components. The costs would occur if the new owners of inferior voting stock are willing to pay less for the same cash-flow rights, either because they expect direct expropriation or because they expect weaker shareholder rights to be associated with weaker performance. Thus, we hypothesize that dual-class status should be more prevalent when private benefits are high and perceived costs are low.

To test this hypothesis, we identify proxies for these costs and benefits and we find several variables that are predictive of dual-class status. All of these predictive variables are measured at the time of a company's IPO: the level of sales as compared to other IPO firms in the same year, the number of pre-existing firms in the same metropolitan area, the total amount of sales among pre-existing firms in the same metropolitan area, a firm's inclusion in a media industry, and, the most powerful predictor, whether a person's name appears in the firm's name.

In Section 4, we use special features of our sample to analyze the relationship between ownership structure and firm value. This topic has inspired a long empirical literature, beginning with Morck, Shleifer, and Vishny (1988). They find that market value is related to insider ownership in a non-monotonic way: for the largest listed firms in 1980, market value is increasing in insider ownership over the range of zero to five percent and decreasing with insider ownership over the range of five to twenty-five percent. The authors hypothesize that this pattern occurs because that the positive incentive effect (from cash-flow rights) dominates for low levels of insider holdings, and then the negative entrenchment effect (from voting rights) dominates at higher levels.

The same qualitative pattern has been confirmed in various other studies: McConnell and Servaes (1990) use a more comprehensive sample of firms and find a similar non-monotonic relationship between ownership and *Q*; Holderness, Kroszner, and Sheehan (1999) find a similar pattern in firms listed in 1935. The relationship between inside ownership and firm value has also been explored outside the United States. La Porta, Lopez-de-Silanes, Sheifer, and Vishny (2002) look at firms in 27 developed countries and find that higher insider cash-flow rights by insiders is associated with higher firm value. Seifert, Gonenc, and Wright (2002) confirm these results using a larger sample.

One constraint in these studies is that the two separate forces – incentives and entrenchment – must be identified using only one variable – ownership. An analysis of dual-class companies offers a way around this problem. Since these firms have equity structures that break the link between cash-flow rights and voting rights, an analysis of dual-class firms allows one to separate the role of these two effects. Two recent papers exploit this separation with studies in the same spirit as our analysis. Lins (2003) examines the relation between firm value, voting rights, and cash-flow rights, for over 1000 companies in 18 emerging markets. He finds that firm value is lower when voting rights exceed cash-flow rights. Claessens et al. (2002) study 1300 firms from eight East Asian countries and find that firm value increases with the cash-flow rights of the largest shareholder but decreases when the voting rights exceed the cash-flow rights.

Empirical studies of valuation and insider ownership are always subject to an endogeneity critique. Early work by Demsetz and Lehn (1985) pointed out that since ownership structure is one of many governance variables that are endogenously determined with firm value and performance, it will always be difficult to uncover the underlying relationships with reduced-form empirical analysis. This argument has been repeated many times, most forcefully by Himmelberg, Hubbard, and Palia (1998) and Coles, Lemmon, and Meschke (2005).

In our analysis, we use dual-class shares to disentangle the incentive and entrenchment effects while also dealing with this endogeneity critique. In particular, the analysis of the determinants of dual-class status (Section 3) provides a first-stage regression for the key ownership variables; these first-stage results can be used for two-stage regressions that adjust for either sample-selection (Heckman regressions) or endogeneity (IV regressions). We find strong evidence that firm value is positively associated with insiders' cash-flow rights. The point estimates are economically large and statistically significant for both single-stage and two-stage regressions. For voting rights, in single-stage regressions we find a significant negative relationship with firm value, and this significance is robust to sample-selection corrections. In IV regressions, the point estimates on voting rights remain negative, but are no longer statistically significant. In a separate specification, we find that the wedge between cash-flow rights and voting rights is negatively associated with valuation. This result is robust to sample-selection corrections, and partially robust to endogeneity corrections. Section 5 concludes the paper with a summary and discussion of these results.

2. Data

2.1. The Dual-Class Sample

To build a comprehensive set of dual-class firms, we first construct a list of possible members – the "candidate sample" – and then we check the SEC filings for each candidate to determine whether it is indeed a dual-class firm. We build the candidate sample using data from the Securities Data Company (SDC) (as amended by Jay Ritter), S&P's Compustat, the Center for Research in Security Prices (CRSP), and the Investor Responsibility Research Center (IRRC).

The SDC candidates are taken from the Global New Issues Database, which tracks corporate new issues activity since 1970 and flags those issues that have a separate class of common stock. We supplement the SDC list with amendments from Jay Ritter's website.[9] To find companies with multiple *traded* share classes, we search the CRSP database for issues with identical six-digit CUSIPs but different two-digit extensions. A further group of candidates are identified from firms listed as dual-class in the IRRC's *Corporate Takeover Defenses* texts from 1990 to 2002 (Rosenbaum 1990, 1993, 1995, 1998, 2000, and 2002).

Our final source of candidates comes from a comparison of Compustat and CRSP, as suggested by Zhang (2003). For each firm-year in Compustat, we match the monthly CRSP file for the month corresponding to the end of the fiscal year. We then compare the shares outstanding field in CRSP with the common shares outstanding field (DATA25) in Compustat. Since CRSP counts only the shares outstanding for a particular stock issue, and Compustat counts all shares for any class of common stock, a difference between these measures may be due to the existence of multiple classes. Thus, if these two share counts differ by more than one percent, we add that company-year to our candidate universe.

To go from the candidate sample to the actual list of dual-class firms – the "dual-class sample" – we first eliminate trusts, closed-end funds, ADRs, units, and REITs. We

[9] http://bear.cba.ufl.edu/ritter/ipodata.htm

then examine the proxy statements and/or 10-Ks for every candidate. Under the Securities and Exchange Act of 1934, all firms are required to disclose the share ownership for each director, and for all officers and directors as a group. Following the past literature on this topic, we refer to this disclosed group of officers and directors as "insiders." These disclosures give separate entries for the holdings in each class – usually as separate columns in a single table – so dual-class status can be definitively determined from these disclosures. After reviewing documents for every firm-year in our candidate sample, we construct the final dual-class sample. These procedures were designed in an attempt to capture all dual-class firms with at least one share class trading on a major U.S. exchange: NYSE, AMEX, or NASDAQ. While we did not check the proxies of all listed firms – only those in our candidate sample – we are confident that the dual-class sample is very close to comprehensive. To test this view, we randomly selected 200 firms from outside the candidate sample and checked their proxies. None of these 200 firms were actually dual class.

Once the dual-class sample is completed, the next step is to determine the insider ownership for each class of stock in every firm year. The SEC disclosures often combine the ownership of stock in the same table with ownership of options, warrants, deferred shares, and other purchase rights. We parsed the tables and footnotes to compute the common-stock ownership, excluding all options and other rights. Since the share classes sometimes have differing cash-flow rights, we also collected dividend data for all firms. First, we coded the dividend information contained in the 10-Ks by class for each firm and year. Second, we used CRSP to identify large, "special" distributions paid out to shareholders.

2.2. Summary Statistics

Table 1 summarizes our universe of companies. We include all Compustat firms except for trusts, closed-end funds, ADRs, units, and REITs. All firms are classified either as single-class or dual-class, with the latter group including all multi-classed firms identified by the procedures in Section 2.1, above. The dual-class sample size varies between a minimum of 362 (in 2002) and a maximum of 504 (in 1998).[10] The single-class sample size varies between a minimum of 6345 (in 2002) and a maximum of 7619 (in 1997). Across all years, about 85 percent of the dual-class sample has at least one non-traded class. In the vast majority of these cases, the superior class is not traded.

Table 1 also gives the voting structure of the dual-class firms. The most common arrangement is a 10:1 structure in which the superior class has ten votes per share and the inferior class has one vote per share. To see how such structures affect the eventual ownership and control of the firms, Table 1 shows the fractions of cash-flow and voting rights held by the insiders. On average, the insiders of dual-class firms own a majority of the voting rights (about 60 percent) and a significant minority of the cash-flow rights (about 40 percent). Nearly all of these voting rights come from the superior voting class stock: less than fifteen percent of the insiders' voting rights come from the inferior voting class.

In about one-third of all dual-class firms, the insiders have a majority of the voting rights ($VOTE > 0.50$) but do not have a majority of the cash-flow rights ($CF < 0.50$). These firms comprise a special case of the dual-class sample, since insiders have

[10] The number of "dual-class" firms with three or more classes varies between 17 and 28 firms over the sample period.

effective control over all corporate decisions, but have claims to less than half of the economic value. We classify these firms as the "separation sample."

A more complete picture of the ownership structure of dual-class firms can be seen in Table 2, which shows a cross-tabulation of cash-flow and voting rights for each firm-year in the dual-class sample. For the majority of firm-years, insiders' voting rights are higher than their cash-flow rights (entries below the 45 degree line), with few firms showing the reverse (entries above the 45 degree line). Entries in the bottom left quadrant are members of the separation sample.

Table 3 gives summary statistics for dual-class and single-class firms for the representative year of 2000. Both groups have about the same average level of assets and equity-market capitalization. Differences in size distributions become more apparent when we compare the medians. The median dual-class company has $482 million in assets versus $138 million for the median single-class company. Similarly, dual-class firms have a median market value of $295 million versus $100 million for single-class firms. Overall, dual-class firms comprise about six percent of the number of public companies and eight percent of the market capitalization.

Table 3 also shows that dual-class firms are significantly more levered than single-class firms, with significant differences in both the means (0.23 vs. 0.17) and medians (0.18 vs. 0.06) of the debt-to-assets ratio. This difference has several possible drivers. One possibility is that dual-class firms are reluctant engage in seasoned equity offerings (SEOs), for fear of diluting some of their control. Cronqvist and Nilsson (2005) find a similar reluctance for SEOs in family-controlled firms. Another possibility, suggested by Moyer, Rao and Sisneros (1992), is that debt is used as an alternative control mechanism in dual-class firms. This finding raises the interesting question of whether dual-class firms possess other countervailing governance mechanisms such as outside directors (as suggested by Moyer, Rao, and Sisneros (1992)), family ties (as suggested by DeAngelo and DeAngelo (1985)), stronger pay-for-performance, or stronger monitoring by outside blockholders. We leave the analysis of this issue for future research.

Dual-class firms are, on average, significantly older than single-class firms, where age is defined as the time (in years) from the firm's CRSP listing date. The average (median) age of dual-class firms in 2001 is 12.87 (7.21) years while the average (median) age for single-class firms is 9.60 (6.67) years. The most likely explanation for this difference is that dual-class firms are less likely to be acquired. The age difference is particularly striking considering the offsetting effect of IPOs, since the fraction of IPOs with dual-class stock has been steadily climbing during our sample period. Since the IPO market is adding proportionally more firms to the dual-class sample, the seasoned dual-class firms must have an even greater age difference over their seasoned single-class counterparts. Finally, there are only small differences for the book-to-market ratio, both in means and medians.[11] This univariate result does not hold up in multivariate analysis, as we will see in Section 4.

To get a sense for the industry concentration of dual-class stocks, we use the industry classification of Fama and French (1997) to place all stocks into one of 48 industry groups as of December 2001. We find that Communications, Business Services, Printing and Publishing, Retail, and Machinery are the five industries with the greatest

[11] For dual-class companies where only one class of stock trades, we compute market value by assuming that the non-traded stock has the same value per share as the traded stock. In the valuation analysis of Section 4, we discuss the sensitivity of our results to this assumption.

number of dual-class firms in 2001. This distribution is different from the rest of the population of firms. Business Services is the largest industry for single-class companies, followed by Electronic Equipment, Trading, Pharmaceutical Products, and Retail. The predominance of communications and printing and publishing is not surprising. DeAngelo and DeAngelo (1985) were the first to suggest that the non-pecuniary private benefits of control may be high in media-related firms and hence may induce founders to establish a dual-class structure.

Our final set of summary statistics focus on the relationship between dual-class status and stock returns. While there is no *a priori* reason to expect any relationship between ownership structure and returns, several recent papers find that returns are related to other governance variables during our sample period [Gompers, Ishii, and Metrick (2003), Bebchuk, Cohen, and Ferrell (2004), Cremers and Nair (2005)]. Thus, it is logical to check whether any correlation exists in our sample. We begin by constructing portfolios: the dual-class portfolio includes all firm-months in the dual-class sample, and the separation portfolio includes all firm-months in the separation sample. The status of each firm is updated once per fiscal year, with all changes to the portfolios made in the following July. For each portfolio, we compute both value-weighted and equal-weighted monthly returns from July 1995 to June 2003. We then estimate the performance-evaluation regression

$$R_t = \alpha + \beta_1 * RMRF_t + \beta_2 * SMB_t + \beta_3 * HML_t + \beta_4 * Momentum_t + \varepsilon_{t,} \quad (1)$$

where R_t is the excess return to the relevant portfolio (dual-class or separation, equal-weighted or value-weighted) in month t, $RMRF_t$ is the month t value-weighted market return minus the risk-free rate, and the terms SMB_t (small minus big), HML_t (high minus low), and $Momentum_t$ are the month t returns to zero-investment factor-mimicking portfolios designed to capture size, book-to-market, and momentum effects, respectively.[12] Although there is an ongoing debate about whether these factors are proxies for risk, we take no position on this issue and simply view the four-factor model as a method of performance attribution. Thus, we interpret the estimated intercept alpha as the abnormal return in excess of what could have been achieved by passive investments in the factors.

Table 4 summarizes the results. For both portfolios, the value-weighted alphas are negative and the equal-weighted alphas are positive. While some of these alphas are economically large, the limited power of these regressions prevents strong inference. Of the four estimated alphas, only the equal-weighted alpha for the dual-class portfolio is significant, and only at the ten-percent level. Given the inconsistent signs and weak significance, we conclude that there is no clear pattern to the abnormal returns.

Our finding of no abnormal returns differs from the findings of other recent papers in the corporate-governance literature. This difference requires some comment, so we offer two interpretations. First, it could be that the results of these earlier papers [Gompers, Ishii, and Metrick (2003), Bebchuk, Cohen and Farrell (2004), Cremers and Nair (2005)] are period-specific or are driven by some omitted factor that is correlated with their respective governance measures but *not* with dual-class status. Alternatively, it

[12] This model extends the Fama-French (1993) three-factor model with the addition of a momentum factor (Carhart (1997)). All factor returns were downloaded from Ken French's website: http://mba.tuck.dartmouth.edu/pages/faculty/ken.french/

is possible that dual-class status was fully incorporated into stock prices by the beginning of our sample period, while the governance measures used in the other papers were not fully incorporated into prices.

While both of these explanations are plausible, we lack the evidence to draw any strong conclusions. On intuitive grounds, we are inclined towards the second explanation. Dual-class stock was a hotly debated governance device during the 1980s and 1990s, with many academic papers and regulatory scrutiny. By the beginning of our sample period in 1995, investors would have had substantial information about dual-class companies and a significant amount of time to digest this information. Given the roughly zero abnormal returns in the subsequent years, we can say that the circa-1995 investors knew what they were buying. In contrast, the shareholder-rights index of Gompers, Ishii and Metrick (2003) is an amalgam of many disparate governance provisions, most of which had limited empirical evidence and regulatory attention. It seems possible that the market would have had far less information about the components of this arbitrary index than about the salient feature of dual-class stock.

3. The Determinants of Dual-Class Status

Using our classification of firms, we can analyze the determinants of dual-class status. This exercise is useful for two reasons. First, we would like to go beyond the univariate comparisons of Table 3 to get a more sophisticated view of the multivariate correlates and determinants of dual-class status. Second, in the valuation analysis of Section 4, we will need to have first-stage regressions for both sample-selection and endogeneity correction, and the analyses here can serve that purpose.

Why would a firm choose to have a dual-class structure? Take the typical case of a firm that chooses a dual-class structure at IPO. Before the IPO, the firm is completely controlled by managers, venture capitalists, and other private investors. Some of these shareholders are expecting to remain active in the firm after the IPO ("insiders"), while others ("investors") are hoping to sell their shares and disengage. The decision to go public often involves negotiations between these two groups, with the insiders trying to preserve some private benefits of control and the investors trying to maximize the post-IPO share price. Assume, for now, that a single-class structure (one share, one vote) would indeed maximize the share price. Then, while a dual-class structure would preserve more of the private benefits of control for insiders, this control would come at some cost to the investors, since the post-IPO price would be lower. If the private benefits of control are larger than the aggregate reduction in share value, then there is scope for the insiders to bargain with investors and obtain a dual-class structure upon the IPO.[13]

For what kinds of firms would we expect to see large private benefits of control? To answer this question, for each year of the sample we estimate a probit regression of

$$DUAL_{it} = BZ_{i,} + u_{it} \qquad (2)$$

where $DUAL_{it}$ is a dummy variable equal to one if firm i is a dual class firm in year t, and 0 otherwise. Z_i is a vector of firm, industry, or market characteristics at the time of firm i's IPO. We include the following variables as components of Z. All of these variables

[13] See Coates (2001) for a detailed exposition about the motivations of dual-class status.

are measured for the year *prior* to firm i's IPO, so each Z_i vector remains constant throughout the sample period.

Industry: Since private benefits of control are likely to vary across industries, we include dummy variables for each of the industries as defined by Fama and French (1997) [FF]. We follow the related literature and remove all financial firms (FF industries #44, #45, and #47) and regulated firms (FF industry #31) from the sample. Thus, we are left with dummy variables for the 44 remaining FF industries, which represents between 92 and 95 percent of the full sample across different years.

Media: A dummy variable equal to one if the firm was a "media" company in its IPO year, and a zero otherwise. We define media companies as those belonging to SIC codes 2710-11, 2720-21, 2730-31, 4830, 4832-33, 4840-41, 7810, 7812, and 7820. Note that this definition overlaps imperfectly with two industries as defined by Fama and French (1997), communications and printing/publishing, so the *Media* variable is not perfectly correlated with the industry dummies. As documented by DeAngelo and DeAngelo (1985), Field (1999), and Smart and Zutter (2003), media firms are more likely to have dual-class status. This seems logical, since control of a media company – newspaper, television network, etc. – provides many opportunities for private benefits.

Name: One possible signal that insiders place a high value on the private benefits of control is if the company is named for one of the insiders. We construct a noisy measure of this naming congruence by examining the company names at the time of IPO for all

13,000+ firms that ever appear in the Compustat sample (both single class and dual class) during our sample period and coding the *Name* dummy variable equal to one if the firm's name at IPO includes a person's name in it.[14] While this *Name* variable is new, the result that family ownership is predictive of dual-class status is shown in other samples by DeAngelo and DeAngelo (1985), Field (1999) and Amoako-Adu and Smith (2001).

StateLaw: Dual-class stock is a powerful anti-takeover protection. Since this protection may be less valuable for companies incorporated in states with anti-takeover laws, we include the StateLaw anti-takeover index from Gompers, Ishii and Metrick (2003).

SalesRank: We hypothesize that private benefits of control are stronger for firms where the founders are still active. It would be prohibitively time-consuming, if not impossible, to determine founder status at IPO for all 13,000 firms in our sample. One plausible proxy for founder status is the age of the firm (time since first incorporation), but this data is also noisy and difficult to collect. As an alternative proxy for firm age (and thus for founder status), we use *SalesRank*, the percentile ranking of the IPO-year sales of the firm relative to other firms with the same IPO year. The reasoning here is that, other

[14] To avoid any biases in the construction of this variable, we employed three research assistants (RAs) who had not worked on any other part of the project. Each of these RAs coded the full sample of firms (both single-class and dual-class, with the full sample given in alphabetical order within industries) on their own, with the instruction to use their judgment as to whether the firm name included a person's name. This name does not have to be a founder's name, since founder's name is unknown; any name will do. Thus, "Bob's Sporting Goods" is just as good as "Trump Casinos." Then, for any firm where the RAs' judgments were not unanimous, they discussed their choices and tried to reach unanimity. The hard cases mostly occur when a proper name is included that could either represent a place name or a person's name. If there was still disagreement after the discussion, then the team did additional research using library databases, internet search engines, company websites, and SEC filings to attempt to learn the origin of the company's name at IPO. If this search failed to provide conclusive evidence, then the RAs voted their beliefs and the majority judgment was used to code the variable. This majority rule was required for less than 2 percent of the sample. While the final variable is certainly a noisy measure, it is untainted by knowledge about dual-class status.

things equal, sales at the time of the IPO is likely to be positively correlated with firm age, which would then be correlated with founder status. Thus, we expect that *SalesRank* (0 = lowest, 100 = highest) would have a negative coefficient in the estimation of (2).

ProfitRank: We measure *ProfitRank* analogously to *SalesRank*, by computing a percentile ranking (0 = lowest, 100 = highest) in the IPO year relative to other firms in the same IPO year. Profits affect both the private benefits of control and the private costs of control. Private benefits of control are likely to be positively correlated with cash flow and profitability, as free cash flows can be diverted towards pet projects and excess compensation. Since investors recognize this relationship, they are likely to demand control discounts that are positively correlated with profitability, thus increasing the private costs of control. Ultimately, the net effect of these benefits and costs is an empirical question, which we examine through the coefficient on *ProfitRank* in (2).

%Firms and *%Sales*: In recent years, some high-profile cases of corporate fraud have occurred at large firms that were major employers in their geographic region. Firms such as Adelphia and HealthSouth have highlighted the phenomenon of CEOs playing local benefactor with corporate funds. One possibility is that private benefits of control are bigger when insiders have the opportunity to be the major employer in their region: that is, when the firm is the "only game in town." In this case, we would be less likely to observe dual-class status the more other firms are located in the same region. Furthermore, firms with an important local presence may use dual-class status as a promise to local authorities that the firm will resist unsolicited takeovers in order to honor

implicit contracts with local governments and other stakeholders.[15] An alternative possibility is that regions crowded with large firms allow more scope for local M&A activity, and dual-class status could serve as an anti-takeover protection. In this case, we would be more likely to observe dual-class status the more "large" firms are located in the same region. To measure the net effect of these forces, we use two variables, *%Firms* and *%Sales*. *%Firms* is the percentage of all Compustat firms located in the same Metropolitan or Micropolitan Statistical Area (MSA) as firm *i* in the year before firm *i*'s IPO.[16] Under this construction, if firm *i* is the first Compustat company in its MSA, then *%Firms* will be equal to zero. *%Sales* is defined analogously using sales: the percentage of all Compustat sales by firms located in the same Metropolitan or Micropolitan Statistical Area (MSA) as firm *i* in the year before firm *i*'s IPO. Under this construction, if firm *i* is the first Compustat company in its MSA, then *%Sales* will be equal to zero.

Sales/RegionSales: As an alternative measure of the "only game in town" effect, we use the ratio of a firm's sales to the sales of all firms in the same region. This measure is a proxy for the firm's "share of the local pie," while the *Sales%* measure (discussed above) is a proxy for the "size of the local pie."

Of course, there are many other variables that might be correlated with dual-class status, several of which have been used by previous authors: for example, CEO salary (Field 1997) and quality of the investment bank Smart and Zutter (2003). In each of

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[15] The same mechanism – a promise to local authorities to resist takeovers – could also explain a positive correlation between dual-class status and the *Name* variable, since the value of the promise would be greater when attached to a person's name and reputation.

[16] The MSA location of the firm is its headquarters location, as given by Compustat. For firms in counties not included in an MSA, we use the county as the geographic unit.

these cases, however, it is not possible for us to obtain a comprehensive sample of these data for all firms at the time of their IPO. Thus, we restrict ourselves to the variables for which we can find either comprehensive data or good proxies.

The results of our probit estimations of (2) are given in Table 5. Since most of the firms remain the same from year to year and most variables are slow-moving, these annual cross-sectional estimates are highly correlated. For a summary measure (last column of Table 5), we compute pooled estimates, where each firm is included in the regression only for the first year it appears in the sample, and then again in any year where its dual-class status has changed.

There are only 159 changes of dual-class status across the seven years of the sample period, so this pooled regression is dominated by firms with only one appearance in the sample. Some readers may wonder why we did not focus more attention on these 159 firms that switched status. The main problem is that status changes are not exogenous, and so it is difficult to draw general conclusions that are free of endogeneity concerns. In the valuation analysis of Section 4, we will handle endogeneity problems by using a two-stage regression with regression (2) as the first stage. It would not be possible to use this regression to adjust for endogeneity of changes, because (2) is designed to predict dual-class status at the time of the IPO, not *changes* in this status following the IPO. While one could try to construct a different regression to predict these changes, the 159 examples in our data are insufficient for a powerful test.

In the pooled regression, we find significant results for five coefficients. Consistent with the intuition sketched above, the coefficient on *Name* is positive and significant at the one-percent level, the coefficient on *Media* is positive and significant at the five-percent level, and the coefficient on *SalesRank* is negative and significant at the one-percent level. For the *%Firms* and *%Sales* variables, the signs of the coefficients are significant in opposite directions: negative for *%Firms* and positive for *%Sales*. One interpretation of this result is consistent with the notion that the "local benefactor" effect dominates when there are few firms (so it is more likely to choose dual-class status when there are few other local firms), while the anti-takeover effect dominates when there are more potential acquirors (so it is more likely to choose dual-class status when there are large local firms).

4. The Valuation of Dual-Class Firms

4.1. Estimation Methods

In this section, we analyze the valuation of dual-class firms, with a focus on the relationship of firm value with dual-class status, cash-flow rights, voting rights, and the wedge between them. The theoretical work on this topic finds no clear relationship between dual-class structure and firm value [Grossman and Hart (1988), Harris and Raviv (1988)], so researchers must turn to the data. Our analysis in this section is the first attempt to estimate these relationships using a comprehensive panel of dual-class firms.

Our valuation measure is industry-adjusted Tobin's *Q*. Despite its limitations, average *Q* has been the workhorse of large-sample valuation studies since Morck, Shleifer, and Vishny (1988). We follow Kaplan and Zingales' (1997) method for the computation of *Q*, as described below in Section 4.1.1. Industry-adjustments are done from the median Q in the 48 industries classified by Fama and French (1997). Then, for each month in our sample, we estimate

$$Q'_{it} = \mathrm{a} + \mathrm{b}X_{it} + \mathrm{c}W_{it} + \mathrm{e}_{\mathrm{it}}, \tag{3}$$

where Q'_{it} is some transformation of industry-adjusted Q (=Q_i minus value-weighted industry Q) for firm i in month t, X_{it} is a vector of insider variables (dual-class status in some specifications, cash-flow rights and voting rights in other specifications) and W_{it} is a vector of firm characteristics. As elements of W, we follow Shin and Stulz (2000) and include the log of the book value of assets and the log of firm age as of December of year t. Morck and Yang (2001) show that S&P 500 inclusion has a positive impact on Q, and that this impact increased during the 1990s; thus, we also include a dummy variable for S&P 500 inclusion in W. While we are interpreting Q as a valuation measure for assets in place, the numerator of Q will also reflect the capitalized value of growth options. As an attempt to control for these options, we include the ratio of R&D to sales, the ratio of capital expenditures to assets, and the ratio of advertising to sales as additional elements of W. Finally, since our measure of Q may be sensitive to capital structure, we also include the book ratio of debt to assets in W. Note that many of these variables are likely to be endogeneously determined with Q, and thus their coefficient estimates in (3) would be biased. Since we are not interested in the interpretation of the coefficients on these controls, such bias is not a problem for our analysis.

In a perfect world, we could estimate an OLS pooled time-series cross-sectional regression of (3) with all firms and all years. There are five problems with this simple pooled regression. The first problem is that empirical estimates of Q are noisy measures of the true Q, and this noise is not symmetric around the mean. This problem can be alleviated by using transformations of Q or robust estimation methods, as is discussed in Section 4.1.1. The second problem is that the error terms in (3) are correlated in each cross-section. This problem can be solved by clustering the cross-sectional observations or by taking the time-series average of cross-sectional regression coefficients (Fama and MacBeth (1973)). These methods are discussed in Section 4.1.2. The third problem is that these cross-sectional estimates are not independent across time. We can adjust for this dependence using either the "double-clustering" approach of Thompson (2006) or by combining techniques developed by Andrews (1993) and Chordia et al. (2004). These methods are discussed in Section 4.1.3. The fourth problem is that we may have sample-selection bias when we estimate (3) only for the subsample of dual-class firms. We handle this problem using the methods of Heckman (1979), as is discussed in Section 4.1.4. Finally, the endogeneity of ownership structures is a perennial problem in empirical studies of ownership and value. In Section 4.1.5, we discuss how we handle endogeneity concerns by using instrumental variables from regression (2).

4.1.1 – Problem #1: Measurement Error in *Q*

As a valuation measure, average Q is an imperfect but frequently used choice. Ideally, we would measure Q as Tobin intended – the ratio of the market value of assets to their replacement cost – but data limitations render estimates of replacement cost to be very imprecise for a sample of this size. Instead, we follow the now-common procedure introduced by Kaplan and Zingales (1997). For each firm i and month t we compute

$$Q_{it} = [BV_{it}\text{ Assets} + MV_{it}\text{ of Common Stock} - BV_{it}\text{ of Common Stock} - \text{Deferred Taxes}_{it}] / BV_{it}\text{ Assets}, \qquad (4)$$

where BV is "book value" and MV is "market value." The MV of equity is measured at the end of each month, and the accounting variables are measured in the current fiscal year for December and the previous fiscal year for January through November. The MV of equity for dual-class firms with non-trading classes is calculated using shares outstanding from proxy statements and assuming equal prices across classes.[17]

Like all estimates of *Q*, ours is subject to measurement error. The accounting data in the denominator of (4) is of particular concern, since book values of some intangible assets can often be quite different from their "true" replacement cost. This is less of a problem for the numerator, since we can capture the market value of intangible assets through the MV of common stock term. Measurement error in the denominator of (4) causes errors in *Q* to be right-skewed, with some very extreme outliers. While measurement error in the dependent variable does not cause bias, it does inflate the residuals and standard errors, making inference more difficult. To the extent that such measurement error is correlated within industries, our use of industry-adjusted Q can provide some mitigation. Nevertheless, the most extreme outliers tend to be driven by firm-specific measurement error, so further steps are necessary.

[17] On average, non-traded stock makes up a small part of capital structure, so this assumption does not have a significant impact on our results. As a robustness check, we made an alternative assumption that the non-traded superior shares were worth 10 percent more than the traded inferior shares, with no qualitative effect on any of the results. Given the results of Nenova (2000), ten percent is certainly an upper bound on this premium. The premium for the superior non-traded class is very unlikely to be negative, since the superior class can be freely converted to the (traded) inferior class in most companies.

To reduce the measurement error, we estimate three different variations of (3) for every variation of X. The first variation uses robust regression to deal with the measurement error by estimating a median regression with industry-adjusted Q ($Q' = Q$ – industry Q) as the dependent variable in (3). The second variation uses $\ln Q'$ (= $\ln Q - \ln$ industry Q) as the dependent variable and estimates (3) using OLS. This log transformation is one natural way to reduce the influence of outliers, but any concave function could serve a similar purpose. The third variation is also estimated by OLS and uses $-1/Q'$ (= $-[1/Q - 1/(\text{industry } Q)]$) as the dependent variable. In this case, the measurement error is in the numerator, which induces much less noise then its inverse. As shown below, all three of these robust methods give quantitatively similar results. In contrast, in untabulated results we also replicated all of our tests using the baseline Q as given in equation (4). As expected, the standard errors were much larger than in the three other variations, and none of the results were statistically significant.

4.1.2 – Problem #2 – Cross-Sectional Dependence

In a cross-sectional regression of firm-level returns on exogenous characteristics, the error terms are correlated, with an unknown correlation structure. Since the numerator of Q is market value, and past returns are a component of market value, our regression in (3) will also have some unknown cross-sectional dependence. With a long-enough time series, this problem can be handled by clustering observations within years. Asymptotically, such clustering is equivalent to the method of Fama and MacBeth (1973) (FM): estimate a separate regression for each cross-section, and then perform inference on the time-series mean and standard deviation of the cross-section coefficients. In one set of regressions, we follow the FM procedure, with some necessary adjustments, as explained below in Section 4.1.3.

It is important to note that all inference in these regressions requires the assumption that each set of cross-sectional regression coefficients is drawn from the same distribution. In other words, the relationship between valuation and ownership structure must be stable over time. It is not possible to explicitly test this assumption: the main reason we need FM in the first place is that we cannot estimate the standard errors of any given set of cross-sectional coefficients. Nevertheless, the approximately "zero" abnormal returns of Table 4 give us some confidence that the relative valuation of dual-class vs. single-class stocks is relatively stable over our sample period.

4.1.3 – Problem #3: Time-Series Dependence of Coefficients

The FM procedure was designed to handle cross-sectional regressions of stock returns on characteristics. For return regressions, it is reasonable to assume that the monthly coefficients are independent. For *Q* regressions, this assumption would be too strong; at the very least, any measurement error in *Q* is likely to be persistent. Thus, to apply FM correctly, we must adjust for time-series dependence in the estimated cross-sectional coefficients. The typical solution to this problem is to estimate a first-order autoregressive model for each coefficient and then use the estimated autoregression coefficients to adjust the FM standard errors. Petersen (2005) demonstrates that this procedure is flawed when there is small-sample bias in the autoregression coefficients. In many applications, this small-sample bias is unavoidable, since researchers must use annual data over limited time periods. In our application, *Q* can be estimated monthly (albeit with high levels of persistence), resulting in a substantial number of observations. Specifically, we begin by estimating monthly cross-sections of (3). Then, to account for time-series correlation in our monthly regression coefficients, we estimate an AR(1) process for each regression coefficient, β_k, as

$$\beta_{kt} = \mu + p\beta_{kt-1} + \varepsilon_{kt}, \qquad (5)$$

where $|p| < 1$ and ε_{kt} is a white noise process, for $t = 1, \ldots, T$. This least-squares estimator of p is known to be downward-biased. If T is "too small," then it is not possible to correct this bias. For large enough T – as we have here – Andrews (1993) provides a method for using least-squares estimates of p to produce median-unbiased estimates. We use Table 2 of Andrews (1993), which is based on 100 time-series observations. Linear interpolation between table entries is reported to be quite accurate, so we use this method.

Given our median unbiased estimate of p, we calculate our corrected standard error of the mean coefficient as:

$$\text{CSE} = \sigma\sqrt{\frac{1}{T} + \frac{2\rho}{(1-\rho)T} - \frac{2\rho(1-\rho^T)}{(1-\rho)^2 T^2}}\,, \qquad (6)$$

where σ^2 is the time-series variance of the coefficients and $T = 96$. This formula is derived in Chordia et al. (2004). Under an AR(1) assumption, this method, which we call "AR-corrected" standard errors, will provide the most efficient estimates. If we drop the AR(1) assumption, then we can still compute robust standard errors that allow for arbitrary correlations of error terms across firms (within each cross-section) and over time (for each firm). Thompson (2006) provides properties for this "double-clustering" technique and shows how it can be easily computed in standard statistical packages. We refer to these standard errors as "Thompson-corrected".

4.1.4 – Problem #4: Sample-Selection Bias

In Section 4.2.1, we use all firms in the sample, with dual-class status serving as a right-hand-side variable. These regressions do not have any sample-selection bias. In Section 4.2.2, however, we use only the dual-class firms, and in Section 4.2.3, we use only the firms in the separation sample. While these estimations can provide useful results for their respective samples, our ability to draw inference for all firms is clouded by the possibility that dual-class firms are different from single-class firms, with these differences inducing different relationships between ownership and firm value. We correct for the possibility of such sample selection by using the methods of Heckman (1979), with regression (2) serving as the selection equation. We cannot use quantile regression in this setting, so we rely on OLS. Furthermore, we cannot compute Thompson-corrected standard errors in this case, so we rely on the AR(1) assumption (discussed in 4.1.3, above) and compute only the AR-corrected standard errors.

4.1.5 – Problem #5: Endogeneity

The endogeneity of ownership structure and firm value is a serious concern for regressions like (3). If firms choose ownership structures based in part on any other input into valuation, then the residuals in (3) would be correlated with the ownership variables, and the coefficient estimates will be biased. In practice, the problem is difficult to solve, since it is hard to find good instruments for ownership structure. In this case, our use of dual-class firms provides a good opportunity to find instruments, because the exogenous variables used in (2) have some explanatory power for the choice of dual-class structure, and may also be useful for ownership variables based on this structure. Note that endogeneity problems are conceptually distinct from the sample-selection problems discussed above in Section 4.1.4. For example, even if the residuals in (3) are uncorrelated with the regressors, it could be the case that dual-class companies are not representative of all companies for the relationship between ownership structure and firm value. In this case, we would have a sample-selection problem without an endogeneity problem. Conversely, dual-class companies could be representative of all companies while ownership structure is endogenous for all companies. In this case, we would have an endogeneity problem without a sample-selection problem. Unfortunately, it is not possible to correct for both problems at the same time. If both problems do exist, then the IV results can still allow for correct inferences, but only on the samples used for the estimation.

4.2. Results

4.2.1 – Full-Sample Results

We begin with the full-sample results. Table 6 gives the results of estimating (3) using a variety of specifications. The coefficients on the *W* vector are omitted from this and all other tables and are available by request from the authors. In Panel A, we give the single-stage regression results. The first five columns of Panel A show the results when a dual-class status dummy variable is used as *X*: column (1) gives median-estimation results with *Q'* as the dependent variable; the next two columns give OLS results with *lnQ'* as the dependent variable with AR-corrected standard errors (column 2) and Thompson-corrected standard errors (column 3); the next two columns gives OLS results with *-1/Q'* as the dependent variable with AR-corrected standard errors (column 4) and Thompson-corrected standard errors (column 5). The coefficients on *DUAL* are negative in all five of these regressions.

In interpreting these results, it is important to realize that dual-class status does not have uniform implications for incentives and entrenchment. Some dual-class firms have only a small difference between insider cash-flow rights and voting rights, while others have a large difference. In the last five columns of Panel A, we use this difference, defined as *WEDGE*, as our *X* variable. We calculate *WEDGE* as insider voting rights minus insider cash-flow rights. For single-class firms, *WEDGE* is zero by definition. The estimation techniques are analogous to those in the *DUAL* regressions. These results show some pattern, with all five point estimates negative, three of which are significant at the ten-percent or five-percent levels.

As discussed in Section 4.1.5, the endogeneity of dual-class status can cloud inference in the single-stage regressions. We estimate two-stage regressions by combining all the regressors from (2) and (3) into a first-stage probit estimate for dual-class status (or for *WEDGE*), and then substituting the fitted value into the right-hand side of (3). Panel B gives the results of these regressions. Each column in Panel B is the two-stage analogue of the respective column in Panel A. Only in column (5) is there a significant coefficient at the ten-percent level; with ten regressions, we should expect that to occur just by chance. Furthermore, in several of the IV regressions the sign of the key coefficient is reversed from their single-stage analogues in Panel A. Overall, we do not find a consistent relationship between *Q* and *DUAL* or between *Q* and *WEDGE*.

4.2.2 – Dual-Sample Results

While the full-sample results are interesting, the real payoff of this data set is the ability to simultaneously estimate coefficients for different ownership variables. In the absence of a structural model, we should not rely on any one specification. Instead, we use several different reduced-form regressions to reflect a variety of plausible pathways from ownership structure to valuation. In our first set of regressions (Table 7), we use cash-flow rights (*CF*) and voting rights (*VOTE*) as the key independent variables. The idea behind these specifications is that *CF* and *VOTE* act differently on the incentives of insiders, and each type of ownership can exert an independent effect on valuation. This motivation was first suggested by Morck, Shleifer and Vishny (1988), and has been repeated in many of the papers that followed. To handle nonlinearities in these effects, many of these previous papers estimated piecewise regressions. To avoid having to choose specific piecewise thresholds, we just include squared terms in some of the specifications.

As in the full-sample case, we first discuss the single-stage results. The first five columns of Panel A provide estimates when only the levels of *CF* and *VOTE* are included. As in all tables in this section, we use five different specifications: column (1) gives median-estimation results with *Q'* as the dependent variable, column (2) gives FM OLS results with *lnQ'* as the dependent variable and AR-corrected standard errors, column (3) gives pooled OLS results with Thompson-corrected standard errors, and columns (4) and (5) replicate columns (2) and (3) with *-1/Q'* as the dependent variable. In each case, the coefficient on *CF* is positive and the coefficient on *VOTE* is negative. These coefficients are statistically significant in seven out of ten cases, with somewhat stronger results for the *VOTE* coefficients.

Columns (6) through (10) estimate the same five specifications using a quadratic form, with *X* containing *CF*, CF^2, *VOTE,* and $VOTE^2$. The pattern of results is similar in all specifications: positive coefficients on *CF* and negative coefficients on CF^2 for the cash-flow variables, and the reverse pattern of negative coefficients on *VOTE* and positive coefficients on $VOTE^2$ for the voting variables. Out of 20 total coefficients, 14 are significant at the ten-percent level, and seven are significant at the one-percent level.

With different signs on the level and squared terms, a diagram can help us to visualize the total effect for these variables. Figure 1 plots the total effect of *CF* and *VOTE* – with 95 percent confidence bands – based on the coefficients in Column 6 of Table 7. The total effect of the *CF* variables peaks when cash-flow rights are about 60 percent, where this level is associated with 25 percentage points higher (median) *Q*. The total effect for the *VOTE* variables falls for the entire range, reaching about –32 percent at *VOTE* = 100 percent. Figure 2 plots the net effect. Like Morck, Shelifer, and Vishny (1988) and McConnell and Servaes (1990), we find that this relationship first increases and then decreases. In our sample, the net effect of insider ownership peaks at about 5 percentage points when ownership is approximately 30 percent.

As discussed above in Section 4.1, these single-stage results may suffer from sample-selection bias: even if the results are correct for all dual-class firms, it does not necessarily mean that we can extend the results to all firms. To make this extension, we adjust for sample selection using the technique of Heckman (1979), with a first-stage probit regression with all the elements of *Z* from regression (2) and *W* from regression (3) as regressors.

The Heckman procedure cannot be used when median regression is the second stage, nor can we compute Thompson-corrected standard errors, so we can only perform these tests for the OLS estimations of *lnQ'* and *–1/Q'* with AR-corrected standard errors. Panel B of Table 7 gives the results. Columns (1) and (2) provide the results for the level regressions, and columns (3) and (4) provide the results for the quadratic regressions. For the level regressions, all the coefficients have the same signs as their analogues in Panel A, and with higher t-statistics in all cases. The quadratic regressions summarized in columns (3) and (4) of Panel B also yield nearly identical results as in Panel A. Overall, the sample-selection regressions are very similar to the single-stage regressions.

In Panel C, we summarize the IV estimations for each specification. The point estimates in Panel C are similar to their single-stage analogues in Panel A, but the standard errors are much larger. Nevertheless, the coefficient on the level term for cash flow is positive significant in four out of the ten IV specifications, with relatively stronger results in the quadratic specifications. This evidence on the importance of cash flow incentives will be explored further in Section 4.2.3.

In our second set of dual-sample regressions (Table 8), we use *CF* and *WEDGE* as the two ownership variables. Since *WEDGE* = *VOTE* – *CF*, it is not possible to include both *CF* and *VOTE* in this regression, so we arbitrarily chose to omit the latter. Readers should think of Table 7 as showing results for *VOTE* and *CF* (since neither of these variable depends directly on the other), and Table 8 as showing results for *WEDGE* (since the *CF* variable is contained in *WEDGE*, we cannot interpret the *CF* coefficients as marginal effects.) Overall, we have three variables of interest, but only two dimensions. In Appendix B, we derive the relationships among the coefficients in Table 7 and 8. For this discussion, we ignore these relationships and just focus on the *WEDGE* coefficients.

The first five columns of Panel A provide estimates when *CF* and *WEDGE* are the *X* variables. We use the same dependent variables and estimation techniques as in Table 7. In all cases, the coefficients on *WEDGE* are negative, and are significant in four out of five of the regressions. When we include squared terms in columns (4) – (6), the coefficients on $WEDGE^2$ are negative in all cases and significant in two of them. The coefficients on the level terms in these quadratic regressions are always insignificant.

The Heckman regressions (Panel B) have qualitatively similar results to the one-stage regressions of Panel A. In columns (1) and (2), the coefficient on *WEDGE* is negative in both cases, significant in one, and very similar to the analogues in Panel A. In columns (3) and (4), the coefficient on $WEDGE^2$ is negative and significant in both cases, and the coefficient on *WEDGE* is close to zero and insignificant in both cases.

Panel C summarizes the results of the IV regressions. In the left half of Panel C, the coefficients on *WEDGE* are negative in all cases, but are only half the magnitude of their analogues from Panels A and B, and are only significant in one of the regressions. With squared terms, there is more evidence of a significant *WEDGE* influence. The coefficients on $WEDGE^2$ are negative in all five specifications, and significant in two of them.[18] Note, however, that the coefficient on *WEDGE* is *positive* and significant in column (7). In all five of these quadratic specifications, the point estimate for the

[18] Note that the standard error on the CF^2 coefficient is missing for column (7). This occurs because there is “too much” autocollinearity in the time-series of coefficient estimates, so the standard error of the mean is effectively equal to infinity.

marginal impact – the derivative of *Q* with respect to *WEDGE* -- turns negative for values of *WEDGE* close to 20 percent, which is approximately the mean value of *WEDGE* in the dual-class sample.

4.2.3 – Separation-Sample Results

When analyzing the entire dual-class universe, we are mixing several different ownership structures. This mixing can cloud inference, especially if the underlying relationship between value and ownership has complex nonlinearities that are not captured by our simple quadratic specification. Previous studies have attempted to handle these nonlinearities with piecewise regressions, but it is difficult to specify the "correct" piecewise segments *a priori*, and concerns about data mining make us wary of conducting an empirical search. Instead, we attempt to adjust for the most obvious possibility of nonlinearity by analyzing firms in our separation sample. Recall that this sample includes all dual class firms where insiders have voting control (*VOTE* > 0.50) but do not have a majority of the cash-flow rights (*CF* < 0.50), and comprises about 40 percent of the dual-class universe. Since the marginal impact of additional voting rights above 50 percent does not affect insider control, we can obtain sharper results by focusing this analysis on the relationship between value and cash-flow rights.

Table 9 summarizes the results of estimating (3) on the separation sample. The panels and columns of Table 9 are analogous to those in Table 7, except that here the *X* vector omits the *VOTE* variables and includes only the *CF* variables. The results for the level regressions (columns (1) to (5) in all panels) are robust across all specifications, with the coefficients on *CF* positive and significant in all cases. For the Heckman specifications (Panel B), the selection equation uses separation-sample status as the dependent variable, and provides nearly identical results to their single-stage analogues (Panel A). It also appears that the first-stage regression fits much better here than in previous regressions, as the standard errors in IV regressions of Panel C are not much higher than their single-stage analogues in Panel A, and the point estimates are significant in three of the five cases for the level regressions. The weakest results are found in the regressions with Thompson-corrected standard errors, where the coefficients are positive but are not significant.

4.2.4 – Summary of Findings

Overall, the results support a positive relationship between firm value and insiders' cash-flow rights, and a negative relationship between firm value and the wedge between voting rights and cash-flow rights. For voting rights by themselves, none of the IV regressions (Table 7) provide significant results. The evidence on cash flow is a bit stronger than the evidence on the wedge, mostly due to the clean results on the separation sample, where the IV regressions (Table 9) provide relatively precise estimates and yield significant coefficients in the majority of the linear specifications.

5. Conclusions

In this paper, we construct and analyze a comprehensive sample of dual-class companies in the United States during the 1995 to 2002 period. We use these data to perform three sets of analyses. First, we show the prevalence of the dual-class structure and summarize the most salient characteristics of dual-class firms. We find that approximately six percent of all Compustat firms have more than one class of common stock, with this group comprising eight percent of the market capitalization of all the firms. The typical dual-class firm has a superior (nontraded) class of shares with ten votes per share and an inferior (traded) class of shares with one vote per share, with insiders controlling about 60 percent of the voting rights and 40 percent of the cash-flow rights.

Second, we use the classification of all Compustat firms to analyze the determinants of dual-class status. We find several variables that predict dual-class status. The most powerful predictor is whether a person's name appears in the company's name at the time of the IPO. Other significant predictors – all measured at the time of IPO – are whether the firm is in a media industry, the number of firms in the same metropolitan area, the size of firms in the same metropolitan area, and the sales of the firm relative to others going public in the same year.

Our third set of analyses constitutes the main contribution of the paper. We use the separation of voting rights and cash-flow rights to disentangle the incentive and entrenchment effects in the relationship of insider ownership and firm value. In a series of single-stage regressions, we find that firm value is positively associated with insiders' cash-flow rights, negatively associated with insiders' voting rights, and negatively associated with the wedge between the two. One advantage of the dual-class sample is that we can use predictive variables from the determinants analysis to control for self-selection of dual-class status or endogeneity of ownership structure. We find that all of the results are robust to a self-selection correction, and the cash-flow and wedge results are partially robust to an endogeneity correction.

A significant relationship between firm value and ownership structure does not imply that any actor is behaving irrationally. A majority owner of a private company can rationally choose to sacrifice some firm value in order to maintain private benefits of control. The ability to control editorial policy at a newspaper, corporate strategy at a software company, or brand identity at a consumer company can all bring utility to individual manager-owners. Such utility can outweigh financial losses, particularly if the insiders are already very wealthy. The analysis of this paper allows us to estimate the average size of these losses in mature companies. With these measurements, we can begin to quantify the impact of this most extreme form of corporate governance.

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Appendix A: Variable Definitions for Q Regressions

We define year *t* as beginning in December of calendar year *t* and ending in November of calendar year *t+1*. Then the additional variables for regressions in each month of year *t* are as follows:

Log assets: The log of assets (Compustat item 6) measured in millions of dollars during fiscal year t.

SP500: An indicator variable for inclusion in the S&P 500 as of the end of calendar year t.

Log age: The log of age in months as of the end of calendar year t.

Debt/Market Equity: The ratio of debt (Compustat item 9) to the market value of equity, where debt is measured in fiscal year *t-1*, the market value of equity is measured at the end of calendar year *t-1*, and the market value of equity for dual-class firms with non-trading classes is calculated using shares outstanding from the year-*t-1* proxy statements and assuming equal prices across classes.

Dividends/Book Equity: The ratio of dividends (Compustat item 21) to book equity (Compustat item 60) in fiscal year *t-1*.

R&D/Sales: The ratio of research and development (Compustat item 46) to sales (Compustat item 12) in fiscal year *t-1*. This variable is set equal to zero when R&D is missing.

Capex/PPE: The ratio of capital expenditures (Compustat item 128) to gross property, plant, and equipment (Compustat item 7) in fiscal year *t-1*.

Advertising/Sales: The ratio of advertising (Compustat item 45) to sales in fiscal year *t-1*. This variable is set equal to zero when advertising is missing.

Sales Growth: One-year sales growth in fiscal year *t-1*.

Diversification: An indicator variable that equals one when the firm has more than one Compustat business segment during fiscal year *t*.

Appendix B: The Relationship Between the Regressions in Tables 7 and 8

The linear version of the regressions in Tables (7) and (8) yield the same inferences: i.e., after omitting the time and firm subscripts, the regressions without squared terms in Table 7 are

$$Q' = a + b_1 CF + b_2 VOTE + cW + e. \qquad (7)$$

The analogues in Table 8 are

$$Q' = a + b_3 CF + b_4 WEDGE + cW + e = a + b_3 CF + b_4 (VOTE - CF) + cW + e. \qquad (8)$$

Thus, $b_1 = b_3 - b_4$, and $b_2 = b_4$. These are indeed the results for all regressions shown in the left halves of Table 7 and Table 8.

However, for the regressions with squared terms, the inferences can be different. In Table 7, the right half of the table uses the specification

$$Q' = a + b_1 CF + b_2 CF^2 + b_3 VOTE + b_4 VOTE^2 + cW + e \qquad (9)$$

The analogues in Table 8 are

$$Q' = a + b_5 CF + b_6 CF^2 + b_7 WEDGE + b_8 WEDGE^2 + cW + e$$

$$= a + b_5 CF + b_6 CF^2 + b_7 (VOTE - CF) + b_8 (VOTE - CF)^2 + cW + e$$

$$= a + (b_5 - b_7)\, CF + b_7\, VOTE + (b_6 + b_8)\, CF^2 + b_8\, VOTE^2 - 2b_8\, (VOTE * CF) \qquad (10)$$

Thus, (10) is the same as (9), except for the inclusion of the interaction term, with a coefficient on that interaction term restricted to be twice the coefficient on $VOTE^2$.

Table 1: Voting and Ownership Structure

Panel A of this table describes dividend and voting arrangements in the dual-class sample between 1995 and 2002. It summarizes the relationship between the superior class and the inferior class with the most votes per share of any inferior class. Panel B summarizes cashflow and voting ownership in the dual-class firms. *VOTE* is the total percentage of votes owned by officers and directors across classes, as reported in proxy statements. *CF* is the total percentage of cashflow ownership by officers and directors. Rights to the firm's cashflows are assumed to be proportional to the ordinary dividends of that class if dividend data exists. If dividend data does not exist or if the dividend distribution is not ordinary, cashflow rights are assumed to be equal across classes. The separation sample refers to dual-class firms that have *VOTE* greater than 50 percent and *CF* less than 50 percent. Proportional directors refers to an arrangement in which different classes elect different directors.

Panel A: Voting and Dividend Structure

	1995	1996	1997	1998	1999	2000	2001	2002
Number of single-class firms	6785	6996	7619	7609	7230	7015	6678	6345
Number of dual-class firms	400	444	485	504	489	482	434	362
In separation sample	167	180	167	175	181	167	171	149
All classes trade publicly	77	76	72	74	77	68	64	52
Only inferior classes trade	281	312	346	360	344	352	322	278
Dual-class voting arrangements								
Voting ratio > 1:10	61	72	84	85	80	81	71	56
Voting ratio = 1:10	105	120	143	155	155	161	150	129
Voting ratio < 1:10	96	107	118	118	111	101	79	65
With proportional directors	138	145	140	146	143	139	134	112
Dual-class dividend arrangements								
Superior < Inferior	73	71	69	68	69	64	52	46
Superior = Inferior	317	360	407	426	412	409	374	311
Superior > Inferior	10	13	9	10	8	9	8	5

Panel B: Ownership Structure

	1995	1996	1997	1998	1999	2000	2001	2002
Average *CF*	39.8	39.7	39.8	39.3	39.7	40.0	40.2	37.6
In Superior Class	22.8	22.6	24.2	24.0	24.0	24.1	24.2	21.5
In Inferior Class(es)	17.0	17.1	15.6	15.3	15.7	15.9	16.0	16.1
Average *VOTE*	61.7	61.1	60.9	60.4	61.0	61.5	62.9	61.8
In Superior Class	54.5	53.9	54.1	53.6	53.7	54.0	55.6	54.5
In Inferior Class(es)	7.2	7.2	6.8	6.8	7.4	7.5	7.3	7.3

Table 2: Relationship Between Managerial Voting Ownership and Cash Flow Ownership

This table summarizes cashflow and voting ownership in the dual-class firms between 1995 and 2002. As defined in Table 1, *VOTE* and *CF* are the total percentages of votes and cashflow, respectively, owned by officers and directors across classes as reported in proxy statements.

		Cashflow (*CF*)																			
		0-5	5 - 10	10 - 15	15-20	20-25	25-30	30-35	35-40	40-45	45-50	50-55	55-60	60-65	65-70	70-75	75-80	80-85	85-90	90-95	>95
	0-5	69	10	0	1	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0
	5 - 10	10	38	5	6	2	0	1	0	0	0	0	0	0	0	0	0	0	0	0	0
	10 - 15	5	17	41	11	7	0	0	0	1	0	1	0	0	0	0	0	0	0	0	0
	15-20	3	10	20	39	17	10	1	3	0	0	0	0	0	0	0	0	0	0	0	0
	20-25	9	11	10	24	24	8	3	4	3	2	0	0	0	0	0	0	0	0	0	0
	25-30	1	5	9	10	22	27	6	2	1	2	0	3	0	0	0	0	0	0	0	0
	30-35	2	13	22	30	18	16	17	7	6	3	0	2	3	0	0	0	0	0	0	0
	35-40	2	22	14	18	27	23	10	11	6	4	0	1	1	1	2	0	0	0	0	0
Voting	40-45	2	17	12	21	21	10	21	11	9	6	1	0	0	0	0	1	0	0	0	0
(*VOTE*)	45-50	4	4	12	19	22	35	19	19	11	21	8	0	1	1	0	1	0	0	0	0
	50-55	0	8	7	9	23	26	24	18	16	16	17	4	2	0	0	1	2	0	0	0
	55-60	1	4	11	19	20	18	17	24	20	24	18	31	14	2	1	0	0	1	2	0
	60-65	0	7	12	22	25	26	25	28	20	31	16	20	21	9	4	1	2	0	0	0
	65-70	0	7	18	6	35	30	38	52	30	33	36	22	11	15	7	0	1	0	1	0
	70-75	0	5	8	9	16	19	24	27	35	29	18	17	9	14	16	2	2	0	0	0
	75-80	0	5	8	5	12	28	16	31	21	23	36	17	25	15	11	10	0	0	0	0
	80-85	2	15	3	6	5	10	22	25	13	17	20	36	24	19	7	12	3	4	0	0
	85-90	6	5	3	2	4	3	2	5	18	23	19	26	17	35	12	22	7	11	0	0
	90-95	1	3	2	1	2	0	0	0	8	7	27	16	42	30	21	15	34	12	3	0
	>95	3	1	6	8	8	8	13	8	11	7	8	5	15	20	26	22	65	30	8	13

Table 3: Summary Statistics

This table gives means and medians (in brackets beneath the mean) of several variables for dual- and single-class firms in 2000. The single-class sample consists of all firms in the CRSP-Compustat merged database, excluding the firms identified as dual-class. *Assets* is the book value of assets in millions of dollars (Compustat item 6); *Debt/Assets* is the ratio of long-term debt (item 9) to assets; *SP500* is a dummy variable for inclusion in the S&P 500 as of the end of calendar year 2000; *Age* is firm age in years as of December 2000; *Market Cap* is market value in millions at the end of 2000, where the market value for dual-class firms with non-trading classes is calculated using shares outstanding from proxy statements and assuming equal prices across classes; and *BM* is the ratio of book value [the sum of book common equity (item 60) and deferred taxes (item 74)] to size at the end of 2000. Significant differences for the means are indicated at the ten-, five-, and one-percent levels by *, **, and ***, respectively. The Wilcoxon rank-sum test p-values for the medians are given in brackets in the third column.

	Dual Class	Single Class	Difference
Assets	3143.38	2143.08	1000.30
	[481.59]	[138.13]	[0.0000]
Debt/Assets	0.23	0.17	0.06***
	[0.18]	[0.06]	[0.0000]
S&P500	0.05	0.05	0.01
	[0.00]	[0.00]	
Age	12.87	9.60	3.27***
	[7.21]	[6.67]	[0.0011]
Market Cap	2052.11	1538.15	513.96
	[294.75]	[99.57]	[0.0000]
BM	1.21	1.09	0.12
	[0.70]	[0.64]	[0.0995]
N	380	4479	

Table 4: Four-Factor Return Regressions

This table presents the results of four-factor regressions of equal- and value-weighted monthly returns for portfolios of firms in the dual-class sample or the separation sample. The separation sample refers to dual-class firms that have *VOTE* greater than 50 percent and *CF* less than 50 percent. The explanatory variables are *RMRF*, *SMB*, *HML*, and *UMD*. These variables, which were downloaded from Kenneth French's website, are the returns to zero-investment portfolios designed to capture market, size, book-to-market, and momentum effects, respectively. The sample period is from July 1995 through June 2003. Standard errors are presented in parentheses, and significance at the ten-, five- and one-percent levels is indicated by *, **, and ***, respectively.

	α	*RMRF*	*SMB*	*HML*	*UMD*
Dual Class Sample					
Value Weighted	-0.134	1.002***	0.132**	0.165**	0.024
	(0.231)	(0.059)	(0.056)	(0.074)	(0.038)
Equal Weighted	0.364*	0.878***	0.646***	0.250***	-0.272***
	(0.217)	(0.055)	(0.052)	(0.069)	(0.035)
Separation Sample					
Value Weighted	-0.245	1.076***	0.308***	0.091	0.015
	(0.340)	(0.087)	(0.082)	(0.109)	(0.055)
Equal Weighted	0.357	0.906***	0.657***	0.406***	-0.228***
	(0.235)	(0.060)	(0.057)	(0.075)	(0.038)

Table 5: Selection Equation

This table reports the results of annual probit regressions from December of each year, where the dependent variable is an indicator variable equal to one when the firm is in the dual-class sample and zero otherwise. The entire sample of single- and dual-class firms is used. The last column shows the results from a single regression that includes each firm during the first year it appears in the sample and again for each change in dual-class status. *Name* is a dummy variable for firms with family names at the time of CRSP listing. Its construction is discussed in the text. *Media* is a dummy variable for firms in media industries at the time of CRSP listing. Media industries are defined as SIC codes 2710-11, 2720-21, 2730-31, 4830, 4832-4833, 4840-41, 7810, 7812, and 7820. *StateLaw* is an index of state takeover laws defined in Gompers, Ishii, and Metrick (2003) from the firm's state of incorporation in the previous year. *SalesRank* is the firm's percentile of sales in its first year appearing in Compustat in the distribution of all other firms new to Compustat. *ProfitRank* is the analogous percentile for income before extraordinary items available for common. *%Firms* is the percentage of firms located in this firm's region in the year prior to its first appearance in Compustat relative to all firms in Compustat during that year. Region is defined as a metropolitan or micropolitan statistical area or a metropolitan division when one exists, and a county otherwise. *%Sales* is the analogous figure for revenue. *Sales/RegionSales* is sales divided by total sales in this firm's region in its first year appearing in Compustat. Dummy variables for CRSP listing year and the 48 industries designated by Fama and French (1997) are also included in the regressions, but these coefficients are omitted from the table. In this set of regressions and in all regressions in the following tables, we drop firms in utilities or financial industries by omitting firms in the Utilities, Banking, Insurance, and Trading industries as designated by Fama and French (1997). Significance at the ten-, five- and one-percent levels is indicated by *, **, and ***, respectively.

	1995	1996	1997	1998	1999	2000	2001	2002	Combined
Name	0.246***	0.242***	0.286***	0.278***	0.277***	0.358***	0.306***	0.232**	0.227***
	(0.078)	(0.081)	(0.080)	(0.080)	(0.083)	(0.087)	(0.093)	(0.100)	(0.078)
Media	0.476**	0.633***	0.672***	0.694***	0.373*	0.304	0.671***	0.514**	0.437**
	(0.197)	(0.203)	(0.192)	(0.201)	(0.208)	(0.211)	(0.203)	(0.210)	(0.192)
StateLaw	0.002	-0.004	0.004	0.004	0.008	-0.006	-0.002	0.005	0.001
	(0.026)	(0.027)	(0.027)	(0.027)	(0.028)	(0.030)	(0.032)	(0.034)	(0.026)
SalesRank	-0.007***	-0.006***	-0.006***	-0.007***	-0.008***	-0.008***	-0.010***	-0.010***	-0.007***
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
ProfitRank	-0.001	-0.001	-0.000	-0.001	0.000	-0.001	0.002	0.002	-0.001
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
%Firms	-0.079**	-0.095**	-0.052	-0.063	-0.085**	-0.085*	-0.063	-0.095**	-0.078**
	(0.038)	(0.039)	(0.038)	(0.039)	(0.042)	(0.045)	(0.045)	(0.048)	(0.038)
%Sales	0.045*	0.061***	0.042*	0.047**	0.056**	0.064**	0.059**	0.084***	0.045**
	(0.023)	(0.023)	(0.022)	(0.023)	(0.024)	(0.026)	(0.026)	(0.027)	(0.023)
Sales/RegionSales	-0.172	-0.293	-0.222	-0.087	-0.089	-0.231	-0.024	-0.188	-0.169
	(0.174)	(0.191)	(0.188)	(0.177)	(0.185)	(0.205)	(0.202)	(0.226)	(0.173)
N	5031	5388	5410	5070	4832	4716	4203	3871	7790

Table 6: Full Sample Valuation Regressions

This table presents results from valuation regressions using the entire sample of dual- and single-class firms. The dependent variable is either Tobin's *Q*, ln(*Q*), or -1/*Q*. *Q* is the ratio of the market value of assets to the book value of assets: the market value is calculated as the sum of the book value of assets and the market value of common stock less the book value of common stock and deferred taxes. The market value of equity for dual-class firms with non-trading classes is calculated using shares outstanding from proxy statements and assuming equal prices across classes. The market value of equity is measured at the end of the current calendar month, and the accounting variables are measured in the current fiscal year for December and the previous fiscal year for January through November. The primary explanatory variables are *DUAL*, an indicator for dual-class status, and *WEDGE*, which equals the difference between voting and cashflow ownership by officers and directors (*VOTE-CF*). Regressions using Thompson standard errors are pooled regressions with standard errors below in parentheses based on Thompson (2006). In the other columns, a separate regression is estimated for each of the 96 months from December 1995 through November 2003. For each explanatory variable, we present the time-series average of the monthly coefficients with the standard error below in parentheses. The calculation of these standard errors corrected for time-series correlation is described in Section 4.1.3.

As additional explanatory variables, we include the log of assets; an indicator for inclusion in the S&P 500; the log of age; the ratio of debt to the market value of equity; the ratio of dividends to book equity; the ratio of research and development to sales; the ratio of capital expenditures to property, plant, and equipment; the ratio of advertising to sales; sales growth; and a diversification dummy. See Appendix A for details on the construction of these variables. Their coefficients are omitted from the table for brevity, but full results are available on request.

In the IV specifications in panel B, we instrument for *DUAL* and *WEDGE* using the other explanatory variables as well as the variables from the selection equation described in Table 5. In the IV-Median specifications, the same instruments are used in a procedure analogous to two-stage-least-squares in which the second stage is a median regression. Significance at the ten-, five- and one-percent levels is indicated by *, **, and ***, respectively.

Panel A

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Regression Type	Median	OLS	OLS	OLS	OLS	Median	OLS	OLS	OLS	OLS
Dependent Variable	*Q*	ln(*Q*)	ln(*Q*)	-1/*Q*	-1/*Q*	*Q*	ln(*Q*)	ln(*Q*)	-1/*Q*	-1/*Q*
DUAL	-0.043	-0.012	-0.016	-0.012	-0.013					
	(0.026)	(0.024)	(0.024)	(0.016)	(0.016)					
WEDGE						-0.187**	-0.116	-0.124*	-0.098	-0.099**
						(0.078)	(0.115)	(0.072)	(0.060)	(0.051)
Thompson std error	No	No	Yes	No	Yes	No	No	Yes	No	Yes

Panel B

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Regression Type	IV-Median	IV	IV	IV	IV	IV-Median	IV	IV	IV	IV
Dependent Variable	*Q*	ln(*Q*)	ln(*Q*)	-1/*Q*	-1/*Q*	*Q*	ln(*Q*)	ln(*Q*)	-1/*Q*	-1/*Q*
DUAL	0.022	0.004	-0.018	0.105	0.099*					
	(0.142)	(0.281)	(0.084)	(0.136)	(0.052)					
WEDGE						-0.136	-0.155	-0.194	0.136	0.147
						(0.443)	(0.838)	(0.283)	(0.403)	(0.183)
Thompson std error	No	No	Yes	No	Yes	No	No	Yes	No	Yes

Table 7: Dual Sample Valuation Regressions: Cash Flow and Voting

This table presents results from valuation regressions using the sample of dual-class firms. The dependent variable is either Tobin's *Q*, ln(*Q*), or (-1/*Q*), where *Q* is as defined in Table 6. The primary explanatory variables are *CF*, CF^2, *VOTE*, and $VOTE^2$. *CF* and *VOTE* are cashflow and voting ownership by officers and directors, as defined in Table 1. Regressions using Thompson standard errors are pooled regressions with standard errors below in parentheses based on Thompson (2006). In other columns, a separate regression is estimated for each of the 96 months from December 1995 through November 2003. For each explanatory variable, we present the time-series average of the monthly coefficients with the standard error below in parentheses. The calculation of these standard errors corrected for time-series correlation is described in Section 4.1.3.

As additional explanatory variables, we include the log of assets; an indicator for inclusion in the S&P 500; the log of age; the ratio of debt to the market value of equity; the ratio of dividends to book equity; the ratio of research and development to sales; the ratio of capital expenditures to property, plant, and equipment; the ratio of advertising to sales; sales growth; and a diversification dummy. See Appendix A for details on the construction of these variables. Their coefficients are omitted from the table for brevity, but full results are available on request.

The Heckman specifications in Panel B use the two-step Heckman procedure, where for each monthly regression, a first-stage probit regression is estimated as in Table 5. In the IV specifications in panel C, we instrument for *CF*, CF^2, *VOTE*, and $VOTE^2$ using the other explanatory variables as well as the variables from the selection equation. In the IV-Median specifications, the same instruments are used in a procedure analogous to two-stage-least-squares in which the second stage is a median regression. Significance at the ten-, five- and one-percent levels is indicated by *, **, and ***, respectively.

Panel A

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Regression Type	Median	OLS	OLS	OLS	OLS	Median	OLS	OLS	OLS	OLS
Dependent Variable	*Q*	ln(*Q*)	ln(*Q*)	-1/*Q*	-1/*Q*	*Q*	ln(*Q*)	ln(*Q*)	-1/*Q*	-1/*Q*
CF	0.269***	0.225	0.204	0.187*	0.168*	0.865***	0.723*	0.603	0.556**	0.481*
	(0.060)	(0.206)	(0.134)	(0.104)	(0.094)	(0.107)	(0.369)	(0.421)	(0.249)	(0.284)
CF^2						-0.715***	-0.587*	-0.464	-0.432**	-0.364
						(0.133)	(0.341)	(0.458)	(0.218)	(0.297)
VOTE	-0.271***	-0.226	-0.225**	-0.176**	-0.172**	-0.578***	-0.646***	-0.512	-0.485***	-0.413*
	(0.053)	(0.150)	(0.107)	(0.072)	(0.076)	(0.149)	(0.185)	(0.313)	(0.144)	(0.222)
$VOTE^2$						0.258*	0.359***	0.241	0.263***	0.203
						(0.140)	(0.117)	(0.254)	(0.081)	(0.178)
Thompson std error	No	No	Yes	No	Yes	No	No	Yes	No	Yes

Table 7, continued

Panel B

	(1)	(2)	(3)	(4)
Regression Type	Heckman	Heckman	Heckman	Heckman
Dependent Variable	ln(*Q*)	-1/*Q*	ln(*Q*)	-1/*Q*
CF	0.214	0.176	0.843**	0.618***
	(0.209)	(0.109)	(0.365)	(0.230)
CF^2			-0.738**	-0.517***
			(0.290)	(0.162)
VOTE	-0.226	-0.179**	-0.635***	-0.484***
	(0.148)	(0.082)	(0.209)	(0.137)
$VOTE^2$			0.346**	0.258***
			(0.134)	(0.073)
Thompson std error	No	No	No	No

Panel C

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Regression Type	IV-Median	IV	IV	IV	IV	IV-Median	IV	IV	IV	IV
Dependent Variable	*Q*	ln(*Q*)	ln(*Q*)	-1/*Q*	-1/*Q*	*Q*	ln(*Q*)	ln(*Q*)	-1/*Q*	-1/*Q*
CF	0.273	0.275	0.432	0.245	0.315*	1.529***	1.169	1.607*	0.998	1.233**
	(0.229)	(0.254)	(0.272)	(0.265)	(0.172)	(0.492)	(0.950)	(0.941)	(0.687)	(0.619)
CF^2						-1.398**	-1.059	-1.393	-0.893	-1.085
						(0.471)	(0.968)	(1.049)	(0.587)	(0.683)
VOTE	-0.198	-0.113	-0.197	-0.115	-0.144	-0.118	-0.488	-0.051	-0.359	0.018
	(0.160)	(0.176)	(0.216)	(0.145)	(0.148)	(0.887)	(0.429)	(0.790)	(0.400)	(0.587)
$VOTE^2$						-0.093	0.305	-0.206	0.198	-0.207
						(0.821)	(0.395)	(0.703)	(0.288)	(0.509)
Thompson std error	No	No	Yes	No	Yes	No	No	Yes	No	Yes

Table 8: Dual Sample Valuation Regressions: Cash Flow and Wedge

This table presents results from valuation regressions using the sample of dual-class firms. The dependent variable is either Tobin's *Q*, ln(*Q*), or (-1/*Q*), where *Q* is as defined in Table 6. The primary explanatory variables are *WEDGE*, $WEDGE^2$, *CF*, and CF^2. *CF* is cashflow ownership by officers and directors, as defined in Table 1, and *WEDGE* equals the difference between voting and cashflow ownership by officers and directors, as defined in Table 6. Regressions using Thompson standard errors are pooled regressions with standard errors below in parentheses based on Thompson (2006). In other columns, a separate regression is estimated for each of the 96 months from December 1995 through November 2003. For each explanatory variable, we present the time-series average of the monthly coefficients with the standard error below in parentheses. The calculation of these standard errors corrected for time-series correlation is described in Section 4.1.3.

As additional explanatory variables, we include the log of assets; an indicator for inclusion in the S&P 500; the log of age; the ratio of debt to the market value of equity; the ratio of dividends to book equity; the ratio of research and development to sales; the ratio of capital expenditures to property, plant, and equipment; the ratio of advertising to sales; sales growth; and a diversification dummy. See Appendix A for details on the construction of these variables. Their coefficients are omitted from the table for brevity, but full results are available on request.

The Heckman specifications in Panel B use the two-step Heckman procedure, where for each monthly regression, a first-stage probit regression is estimated as in Table 5. In the IV specifications in panel C, we instrument for *WEDGE*, $WEDGE^2$, *CF*, and CF^2 using the other explanatory variables as well as the variables from the selection equation. In the IV-Median specifications, the same instruments are used in a procedure analogous to two-stage-least-squares in which the second stage is a median regression. Significance at the ten-, five- and one-percent levels is indicated by *, **, and ***, respectively.

Panel A

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Regression Type	Median	OLS	OLS	OLS	OLS	Median	OLS	OLS	OLS	OLS
Dependent Variable	*Q*	ln(*Q*)	ln(*Q*)	-1/*Q*	-1/*Q*	*Q*	ln(*Q*)	ln(*Q*)	-1/*Q*	-1/*Q*
CF	-0.002	-0.000	-0.021	0.010	-0.005	0.280***	0.158	0.115	0.157	0.126
	(0.087)	(0.076)	(0.109)	(0.040)	(0.071)	(0.083)	(0.215)	(0.384)	(0.134)	(0.253)
CF^2						-0.355***	-0.219	-0.181	-0.190	-0.166
						(0.100)	(0.256)	(0.431)	(0.175)	(0.278)
WEDGE	-0.271***	-0.226	-0.225**	-0.176**	-0.172**	-0.091	0.052	0.021	-0.040	-0.051
	(0.053)	(0.150)	(0.107)	(0.072)	(0.076)	(0.176)	(0.333)	(0.228)	(0.060)	(0.173)
$WEDGE^2$						-0.304	-0.496**	-0.429	-0.247***	-0.216
						(0.194)	(0.222)	(0.317)	(0.069)	(0.254)
Thompson std error	No	No	Yes	No	Yes	No	No	Yes	No	Yes

Table 8, continued

Panel B

	(1)	(2)	(3)	(4)
Regression Type	Heckman	Heckman	Heckman	Heckman
Dependent Variable	ln(*Q*)	-1/*Q*	ln(*Q*)	-1/*Q*
CF	-0.013	-0.003	0.285	0.218
	(0.074)	(0.039)	(0.251)	(0.136)
CF^2			-0.382	-0.277*
			(0.268)	(0.160)
WEDGE	-0.226	-0.179**	0.023	-0.047
	(0.148)	(0.082)	(0.302)	(0.079)
$WEDGE^2$			-0.446**	-0.237***
			(0.215)	(0.074)
Thompson std error	No	No	No	No

Panel C

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Regression Type	IV-Median	IV	IV	IV	IV	IV-Median	IV	IV	IV	IV
Dependent Variable	*Q*	ln(*Q*)	ln(*Q*)	-1/*Q*	-1/*Q*	*Q*	ln(*Q*)	ln(*Q*)	-1/*Q*	-1/*Q*
CF	0.075	0.162*	0.234	0.130	0.171	0.989**	0.716	0.864	0.702	0.872
	(0.068)	(0.096)	(0.209)	(0.079)	(0.129)	(0.470)	(0.975)	(0.852)	(0.477)	(0.547)
CF^2						-1.086**	-0.694	-0.888	-0.711	-0.929
						(0.546)		(1.021)	(0.641)	(0.650)
WEDGE	-0.198	-0.113	-0.197	-0.115	-0.144	0.269	0.319*	0.690	0.046	0.227
	(0.160)	(0.176)	(0.216)	(0.145)	(0.148)	(0.171)	(0.171)	(0.576)	(0.100)	(0.417)
$WEDGE^2$						-0.877**	-0.804*	-1.657	-0.324	-0.741
						(0.396)	(0.477)	(0.950)	(0.199)	(0.671)
Thompson std error	No	No	Yes	No	Yes	No	No	Yes	No	Yes

Table 9: Separation Sample Valuation Regressions

This table presents results from monthly valuation regressions using the sample of dual-class firms that have *VOTE* greater than 50 percent and *CF* less than 50 percent. The dependent variable is either Tobin's *Q*, ln(*Q*), or -1/*Q*, where *Q* is as defined in Table 6. The primary explanatory variables are *CF* and CF^2. *CF* is cashflow ownership by officers and directors, as defined in Table 1. Regressions using Thompson standard errors are pooled regressions with standard errors below in parentheses based on Thompson (2006). In other columns, a separate regression is estimated for each of the 96 months from December 1995 through November 2003. For each explanatory variable, we present the time-series average of the monthly coefficients with the standard error below in parentheses. The calculation of these standard errors corrected for time-series correlation is described in Section 4.1.3.

As additional explanatory variables, we include the log of assets; an indicator for inclusion in the S&P 500; the log of age; the ratio of debt to the market value of equity; the ratio of dividends to book equity; the ratio of research and development to sales; the ratio of capital expenditures to property, plant, and equipment; the ratio of advertising to sales; sales growth; and a diversification dummy. See Appendix A for details on the construction of these variables. Their coefficients are omitted from the table for brevity, but full results are available on request.

The Heckman specifications in panel B use the two-step Heckman procedure where for each monthly regression, a first-stage probit regression is estimated where the dependent variable is a dummy variable equal to one when the firm is in the separation sample and the explanatory variables described in Table 5 are used. In the IV specifications in panel C, we instrument for *CF* and CF^2, using the other explanatory variables as well as the explanatory variables from the selection equation described in Table 5. In the IV-Median specifications, the same instruments are used in a procedure analogous to two-stage-least-squares in which the second stage is a median regression. Significance at the ten-, five- and one-percent levels is indicated by *, **, and ***, respectively.

Panel A

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Regression Type	Median	OLS	OLS	OLS	OLS	Median	OLS	OLS	OLS	OLS
Dependent Variable	*Q*	ln(*Q*)	ln(*Q*)	-1/*Q*	-1/*Q*	*Q*	ln(*Q*)	ln(*Q*)	-1/*Q*	-1/*Q*
CF	0.718***	0.688**	0.705***	0.508***	0.521***	0.592***	0.998	0.548	0.603	0.444
	(0.120)	(0.277)	(0.238)	(0.121)	(0.174)	(0.495)	(0.818)	(0.973)	(0.460)	(0.720)
CF^2						-1.511*	-0.508	0.263	-0.149	0.129
						(0.900)	(1.790)	(1.620)	(0.941)	(1.162)
Thompson std error	No	No	Yes	No	Yes	No	No	Yes	No	Yes

Table 9, continued

Panel B

	(1)	(2)	(3)	(4)
Regression Type	Heckman	Heckman	Heckman	Heckman
Dependent Variable	ln(*Q*)	-1/*Q*	ln(*Q*)	-1/*Q*
CF	0.656***	0.480***	0.825	0.493
	(0.244)	(0.115)	(0.611)	(0.372)
CF^2			-0.275	-0.013
			(1.410)	(0.796)
Thompson std error	No	No	No	No

Panel C

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Regression Type	IV-Median	IV	IV	IV	IV	IV-Median	IV	IV	IV	IV
Dependent Variable	*Q*	ln(*Q*)	ln(*Q*)	-1/*Q*	-1/*Q*	*Q*	ln(*Q*)	ln(*Q*)	-1/*Q*	-1/*Q*
CF	0.540***	0.534*	0.455	0.409**	0.360	1.831**	1.274*	1.028	0.772*	0.362
	(0.192)	(0.299)	(0.408)	(0.163)	(0.296)	(0.806)	(0.756)	(1.848)	(0.394)	(1.345)
CF^2						-2.170	-1.225	-0.976	-0.591	-0.003
						(1.308)	(1.695)	(3.064)	(0.830)	(2.208)
Thompson std error	No	No	Yes	No	Yes	No	No	Yes	No	Yes

Figure 1

This figure plots the estimated relationship between Tobin's *Q* and the ownership variables, *CF* and *VOTE*, with their 95 percent confidence bands. It uses the parameters in the sixth column of Panel A in Table 7 and holds the control variables fixed at zero. To compute the confidence bands, we assume that the regression coefficients follow AR(1) processes as described in Appendix B, and we allow for covariance between the contemporaneous random elements of these processes.

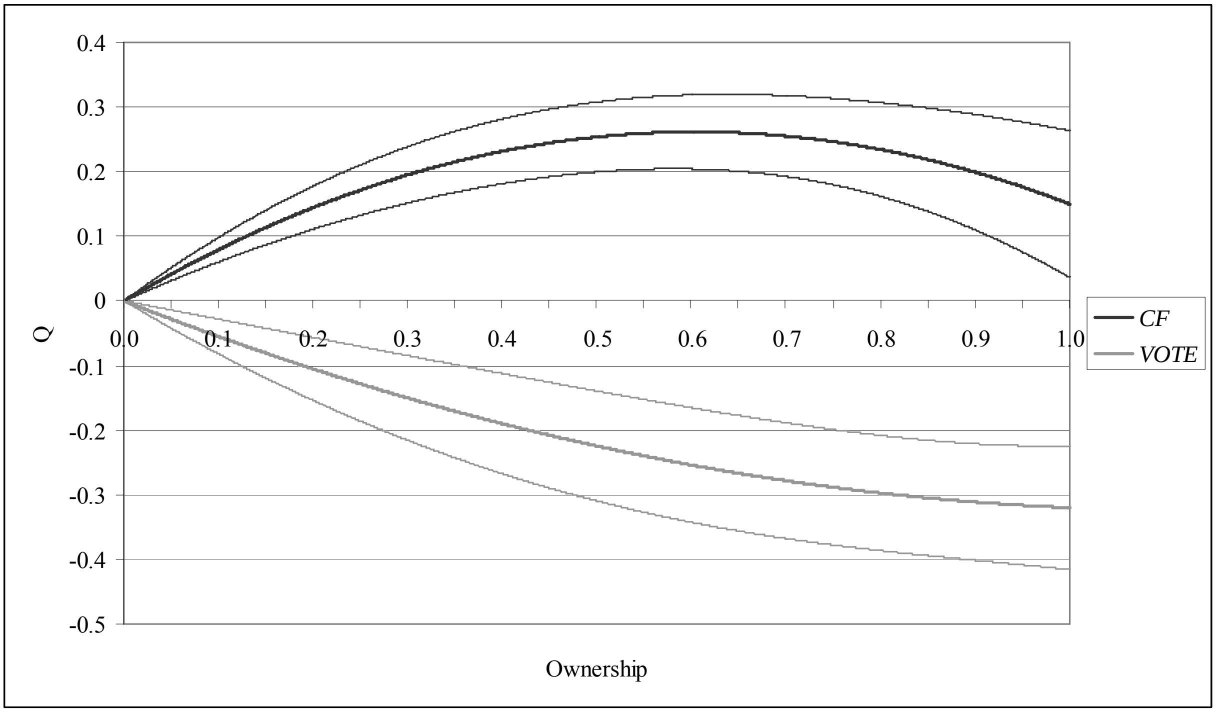

Figure 2

This figure plots the estimated relationship between Tobin's *Q* and ownership when *CF* equals *VOTE*. It uses the parameters in the sixth column of Panel A in Table 7 and holds the control variables fixed at zero.