Capital Market Pressures and Earnings Management: Evidence from U.S. Dual-Class Firms

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ABSTRACT

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In a dual-class share structure, one class of common stock typically has more votes per share than the other, but both classes have equal or similar cash flow rights per share. While this dual-class structure is likely to entrench management, it potentially reduces capital market pressures on managers thereby reducing the need for managers to manipulate earnings. In this study, I compare the earnings management behavior among dual-class firms relative to a matched sample of single-class firms. I find that dual-class firms report a lower level of abnormal accruals, are less likely to produce small positive earnings surprises, and are less likely to release hidden earnings reserves. Further supporting the notion of reduced capital market pressures, when dual-class firms miss analysts' consensus forecasts, their three-day abnormal returns around earnings announcements are significantly less negative compared to their single-class counterparts. Dual-class firms also attract more long-term institutional investors, and have higher future operating and stock return performance. Overall, the results support the hypothesis that reduced capital market pressures lead to less short-term earnings manipulation among dual-class firms.

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"Many companies are under pressure to keep their earnings in line with analysts' forecasts. Therefore, they often accept smaller, but predictable, earnings rather than larger and more unpredictable returns. Sergey and I feel this is harmful, and we intend to steer in the opposite direction. ... We will make decisions on the business fundamentals, not accounting considerations, and always with the long term welfare of the company and shareholders in mind."

– An excerpt from the IPO Prospectus of Google Inc.

1. Introduction

In a firm with a single-class of common stock, cash flow rights (i.e., claims on cash dividend payouts) and voting rights (i.e., the ability to elect the board of directors and make decisions that require shareholder approval) are equal, and a controlling shareholder bears *pro rata* the wealth consequences of the shareholder's decisions. In a dual-class share structure, one class of common stock typically has more votes per share ("the superior class") than the other ("the inferior class"), but both classes have equal or similar cash flow rights per share.[1] Thus dual-class structures allow controlling shareholders to escape the *pro rata* consequences of their decisions by creating a wedge between cash flow and voting rights. In many cases controlling shareholders are able to control dual-class firms while holding only a minority equity stake. For example, utilizing a dual-class share structure, insiders of Google Inc. are able to control 78 percent of the voting power, even though they own only 30 percent of the total shares outstanding.

Because dual-class share structures allow controlling shareholders to insulate themselves from the market for corporate control, insiders (i.e., executive officers and board directors) are able to expropriate private benefits at the expense of outside

[1] According to Gompers, Ishii, and Metrick (2005), about 6 percent of the publicly-traded companies in the U.S. have more than one class of common stock, comprising 8 percent of the market capitalization of all firms. The most common arrangement is a 10:1 voting structure in which the superior class has ten votes per share and the inferior class has one vote per share.

shareholders (e.g., Partch 1987; Grossman and Hart 1988). However, many insiders argue that dual-class share structures allow them to focus on long-term value maximization without the distraction of temporary fluctuations in a firm's share price due to its missing analysts' earnings forecasts. Capital market pressures usually impel managers to use their accounting discretion to affect reported earnings and short-term stock prices, which runs counter to firms' long-term growth and value (Stein 1989; Graham, Harvey, and Rajgopal 2005). Without undue capital market pressures, insiders at dual-class firms may have a reduced incentive to manipulate earnings to meet short-term earnings goals. For example, in their letter to shareholders, Google's founders made clear their desire to continue focusing on long-term value creation even after its IPO. To my knowledge, the latter argument has not been empirically tested in existing research.

This study examines the impact of dual-class structures on earnings management behavior using a sample of dual-class and single-class firms over the period of 1994-2006. My main hypothesis is simple: With less capital market pressure, dual-class structures in the U.S. reduce managers' incentives to aggressively manage earnings when actual earnings fall short of market expectations. I test the capital market pressure hypothesis using three interrelated settings. In the first setting, I investigate whether relative to single class firms, dual-class firms exhibit a similar propensity of meeting or just beating analysts' forecasts. Since insiders can distort firms' economic performance through financial reporting choices and/or through real operating decisions, my second and third settings are related to, respectively, the propensity of accruals management, and the propensity of releasing prior earnings reserves, which involves cutting R&D expenses to boost short-term earnings.

I initially analyze the distributional properties of the earnings surprises of dual-class firms and compare them to the distributions of single-class firms. For single class firms, in the distribution of earnings surprises there is a discontinuity around the threshold of zero, while the earnings surprise distribution for dual-class firms does not exhibit such discontinuities. Thereby I confirm that capital market pressures appear to prompt managers in single-class firms to opportunistically manage earnings in order to minimally meet the analysts' forecasts when they otherwise would have missed the benchmark. In contrast, dual-class firms are less likely to meet or just beat analysts' forecasts, after controlling for managers' incentives and abilities to manage earnings and external monitoring factors. Furthermore, reported earnings for dual-class firms exhibit less extreme accruals relative to single-class firms, and dual-class firms appear less prone to releasing earnings reserves to satisfy short-term earnings objectives. Overall, these results suggest that U.S. dual-class firms exhibit less short-term earnings management, supporting the capital market pressure hypothesis.

In additional tests of the capital market pressure effect, I examine the abnormal returns around earnings announcements, as well as ownership by long-term institutional investors. First, I show that for firms whose annual earnings miss analysts' consensus forecasts, abnormal returns around earnings announcements are less negative (larger) for dual-class firms than for matched single-class firms. Dual class firms also attract a higher proportion of long-term institutional investors, which is consistent with these firms facing less intense capital market pressures to achieve short-term performance. Moreover, consistent with the superior earnings quality of dual-class firms, these firms outperform their single-class counterparts in terms of operating and stock return

performance.

I control for endogeneity using multiple methods. Studies about share structure choice face the methodological issue of endogeneity because firms may self-select into a dual-class share structure. Consequently the earnings quality of dual-class firms may be correlated with various firm characteristics. I address the endogeneity issue by constructing a matched sample of single-class firms to better assess the impact of dual-class structure on a firm's financial reporting behavior.[2] In robustness checks, I also conduct a two-stage regression analysis that accounts for potential endogeneity in firms' decisions to select into dual-class structures. In addition, I utilize a sample of firms that either initiated or eliminated dual-class structures during the sample period, on which I test the main hypotheses with fixed-effects panel specifications to control for unobserved heterogeneity. Taken as a whole, the main results are both statistically and economically significant and are robust to concerns of endogeneity, outliers, alternative regression specifications, and the inclusion of other control variables.

My analyses of dual-class structures offer two advantages over other research settings to examine the impact of capital market pressures on managers' financial reporting behaviors. First, unlike most prior studies that use managerial ownership to simultaneously investigate the effects of incentive alignment and managerial entrenchment, an analysis of dual-class firms allows one to separate the role of these two effects (Gompers et al. 2005). Second, as one of control-enhancing mechanisms, dual-

[2] Heckman, Ichimura, and Todd (1998) provide theoretical support for matching as an econometric technique for addressing endogeneity.

class structures arguably provide cleaner measures of control than other settings, such as family ownership, managerial ownership and ESOPs.[3]

My study contributes to the literature in three ways. First, this study uses a U.S. setting to provide evidence on the relationship between earnings management and dual-class firms. In contrast, much of the recent empirical literature (e.g., Fan and Wong 2002; Haw, Hu, Hwang, and Wu 2004) takes a broad view, analyzing the earnings quality of dual-class firms across countries. In particular, Haw et al. present evidence suggesting that earnings management (captured by absolute total abnormal accruals), induced by the divergence of insiders' control and cash flow rights, is significantly limited in countries with high statutory protection of outside shareholders and effective extra-legal institutions. Such multi-country studies run the risk of only imperfectly controlling for country-specific legal and institutional factors (Wysocki 2004). It is therefore important to complement these international studies with country-specific studies, in which the regulatory environment and accounting rules are held constant. In adopting the U.S. setting, my study is one of the few papers that offer evidence regarding the earnings quality of U.S. dual-class firms.

Second, my paper highlights the linkage between reduced capital market pressures and the financial reporting quality of U.S. dual-class firms. To my knowledge, this relationship has not been systematically examined in the literature. When discussing the financial reporting quality of dual-class firms, existing research invariably emphasizes the entrenchment effect. For example, Francis, Schipper, and Vincent (2005) find that dual-class share structures, when accompanied by entrenched managers, result

[3] Villalonga and Amit (2006) show that, in their sample of Fortune 500 firms, the average excess control over cash flow rights held by family blockholders is merely 3.5%.

in lower earnings informativeness in the U.S. However, strong legal protection of outside shareholders in the U.S. mitigates excessive expropriation by controlling shareholders (Anderson and Reeb 2003). This paper is the first study that emphasizes the beneficial effect of dual-class structures on financial reporting behavior in the U.S. setting. That is, dual-class structures remove a substantial amount of capital market pressures, allowing managers to adopt a strategy to build long-term enterprise value. As predicted, the likelihood of having small positive earnings surprises, accruals management, and the release of earnings reserves, all decrease with dual-class structures.

Finally, existing research on U.S. dual-class firms presents mixed evidence regarding whether dual-class share structures create or destroy shareholder value.[4] There are few attempts to examine operating performance of dual-class firms, despite the availability of the data (Adams and Ferreira 2007). With the aid of a larger sample, I show that the presence of dual-class structures in the U.S. alleviates earnings management, and facilitates superior accounting and stock performance. Thus, this study contributes to the debate about the impact of dual-class structures on maximizing long-term shareholder value.

The remainder of this paper proceeds as follows. In section 2, I describe prior research on dual-class structures and formulate three hypotheses regarding earnings management behaviors of dual-class firms as opposed to their single-class counterparts. Section 3 describes the research methodology and sample selection. Sections 4-6 provide empirical results and robustness tests. Section 7 concludes this paper.

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[4] For an extensive review of empirical evidence of U.S. dual-class firms, see Adams and Ferreira (2007). In contrast, dual-class structures of international firms are always detrimental to shareholders (Claessens, Djankov, Fan, and Lang 2002; Lins 2003; Cronqvist and Nilsson 2003).

2. Hypotheses Development

2.1 Dual-class structures and managers' incentive to manipulate earnings

Dual-class share structures are created by insiders, especially founding families, to prevent the dilution of control while gaining access to capital markets, and to provide an effective defense against hostile takeovers (DeAngelo and DeAngelo 1985; Stulz 1988). Insiders at dual-class firms typically have voting control and a relatively small claim on dividends and earnings, so they can pass off the majority of the financial risk to their outside shareholders. Consequently, dual-class firms face severe agency problems that arise from conflicts between controlling and non-controlling shareholders (Fama and Jensen 1983; Shleifer and Vishny 1986). Insiders are able to extract private benefits through self-dealing, extravagant executive pay, perquisite consumption, and even outright theft (Faccio, Lang, and Young 2001). Such extraction of private control benefits, if detected, is likely to invite external intervention by outside shareholders, stock exchanges, and/or regulators. The desire to avoid external monitoring, potential legal costs, and an associated loss of reputation may encourage insiders to mask their private benefits by managing reported accounting numbers (Haw et al. 2004). In addition, according to Healy and Whalen (1999), earnings management becomes more likely when external stakeholders find the practice difficult to detect due to inadequate disclosure by dual-class firms (e.g., Tinaikar 2006). Therefore, dual-class structures can be associated with aggressive earnings management.

However, dual-class structures may actually reduce managers' incentive to manipulate earnings given reduced capital market pressures. Because a dual-class share structure is the most effective form of anti-takeover protection (Gompers et al. 2005), it

can insulate a company from the need to meet short-term financial expectations that can be detrimental to building long-term value. With no takeovers to worry about, controlling shareholders and managers may be more attentive to firms' long-term value creation. Furthermore, since superior-class stocks that provide extra voting rights often cannot be traded, dual-class firms' performance may benefit from controlling shareholders' long-term and sustainable presence. Based on the above discussion, I hypothesize that dual-class share structures lead to a lower incidence of earnings manipulation relative to single-class structures (the capital market pressure hypothesis).

The long-term operational history of dual-class share structures residing in a number of U.S. firms provides an ideal capital market setting to examine earnings management behavior given reduced capital market pressures surrounding dual-class firms. *Ex ante*, there are at least two reasons that I expect to find evidence supporting the capital market pressure hypothesis.

First, the U.S. regulatory environment is highly effective in protecting outside shareholders and as such significantly limits the advantages of controlling shareholders. Consequently, much of the potential for opportunistic actions that can be taken by controlling shareholders under foreign legal regimes is sharply limited in the U.S. (Stulz 1999). A number of empirical studies have found supportive evidence that the U.S. legal system reduces private benefits that controlling shareholders at dual-class firms can extract from the firm (Nenova 2003; Dyck and Zingales 2004; Doidge 2004). In light of this, it is less likely that insiders at U.S. dual-class firms manage earnings out of their desire to cover up expropriation activities.

Second, among U.S. public firms, capital market pressures have intensified over the past twenty years.[5] In response to capital market pressures, managers often behave myopically by manipulating current earnings in order to inflate short-term stock prices at the expense of long-term firm value creation (Stein 1989). Recent studies (reviewed by Healy and Wahlen 1999; Dechow and Skinner 2000) and a survey of financial executives (Graham et al. 2005) show that capital market incentives dominate other incentives for earnings management. Regulators have also expressed similar concerns.[6] Thus, the intensity of capital market pressures existing in U.S. single-class firms provides a sharper contrast with the reduced pressures of dual-class firms, and allows me to triangulate financial reporting behaviors of dual-class firms affected by reduced capital market pressures.

In the following sections, I test the capital market pressure hypothesis using the three interrelated settings on dual-class firms' tendency to manage earnings: the propensity of meeting or beating earnings expectations; the propensity of accruals management; and the propensity of releasing prior earnings reserves to meet or beat market expectations.

2.2 Propensity to meet or just beat analysts' forecasts

Due to pressures and incentives from capital markets, managers are paying increased attention to meeting forecasted earnings benchmarks (Graham et al. 2005;

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[5] Capital market pressures may take the form of equity incentives (Bartov and Mohanram 2004; Cheng and Warfield 2005), takeover threats (Stein 1988), career concerns (Graham et al. 2005), or needs for external financing (Bhojraj and Libby 2005).

[6] In a speech made at the National Association for Business Economics on March 24, 2003, the former Chairman of U.S. Securities and Exchange Commission (SEC), William Donaldson, stated: "Corporate America developed a short-term focus, fueled by an obsession with quarter-to-quarter earnings and the pervasive temptation inherent in stock options. The game of earnings projections ... created an atmosphere in which "hitting the numbers" became the objective, rather than sound, long-term strength and performance."

Brown and Caylor 2005). Prior research has documented the reward for meeting analysts' forecasts (Bartov, Givoly, and Hayn 2002; Kasznik and McNichols 2002). There are also substantial penalties for firms that miss analysts' consensus earnings forecasts (Skinner and Sloan 2002). Consistent with these findings, Matsunaga and Park (2001) document a significant adverse effect on a manager's annual cash bonuses when the firm's quarterly earnings fall short of certain earnings benchmarks, while Farrell and Whidbee (2003) document higher rates of job dismissal for managers who fail to meet analysts' earnings targets.

Recent studies of earnings management suggest that the disproportionate likelihood of meeting or just beating analysts' consensus earnings forecasts is a manifestation of earnings management (e.g., Degeorge, Patel, and Zeckhauser 1999; Matsumoto 2002; Burgstahler and Eames 2003).[7] In the face of reduced capital market pressures, I expect that managers at dual-class firms have less incentive to manage earnings so as to meet or beat market expectations. Specifically, the first hypothesis is given as follows:

> **H1**: *Ceteris paribus*, dual-class firms have a lower incidence of meeting or just beating analysts' earnings forecasts than single-class firms.

2.3 Propensity for accruals management

In order to meet or beat market expectations, or in response to other capital market incentives, managers have two major ways to manage earnings: by manipulating

[7] Meeting or beating certain earnings benchmarks may not indicate earnings management since proxies of abnormal accruals may be correlated with firm performance (Dechow, Richardson, and Tuna 2003; Durtschi and Easton 2005; Jacob and Jorgensen 2007). Specifically, merely showing positive associations between a specific discretionary accrual measure and beating the profit and earnings increase benchmarks is not sufficient to conclude that the discretionary accrual measure detects earnings management around these benchmarks. However, Ayers, Jiang, and Yeung (2006) demonstrate that the underlying associations between firm performance and discretionary accrual measures are less problematic for tests of earnings management related to analysts-based unexpected earnings. Therefore, I focus on analysts' earnings expectations as the only earnings benchmark in this study.

accruals and/or by strategically manipulating real activities, such as R&D expenses. Following the discussions from the prior sections, I expect that dual-class share structures, due to reduced capital market pressures, induce less accruals management. The hypothesis is formulated as follows:

> **H2**: *Ceteris paribus*, dual-class firms report lower abnormal accruals than single-class firms.

2.4 Propensity for releasing earnings reserves

Conservative accounting – such as expensing R&D expenditures rather than capitalizing and amortizing them – creates unrecorded earnings reserves.[8] Earnings reserves can be released, creating short-term earnings, by subsequently reducing investment or reducing growth in assets that are subject to conservative accounting (Penman and Zhang 2002). Existing literature provides evidence that firms opportunistically reduce R&D spending to meet or beat earnings benchmarks (Baber, Fairfield, and Haggard 1991; Perry and Grinaker 1994; Roychowdhury 2006). In a survey conducted by Graham et al. (2005), about 80 percent of financial executives indicated that they would decrease discretionary spending on items like R&D and advertising expenses to meet earnings targets. Without undue capital market pressures, managers at dual-class firms might not manage investment activities and create short-term earnings by releasing earnings reserves from prior periods. Thus, I hypothesize the following:

> **H3**: *Ceteris paribus*, dual-class firms are less likely to release earnings reserves compared to single-class firms.

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[8] More examples of conservative accounting include the following: choosing last in, first out (LIFO) over first in, first out (FIFO) when accounting for inventories with rising prices; using short estimated asset lives for depreciation; and consistently overestimating allowances for doubtful accounts, sales returns, or warranty liabilities.

3. Sample Construction and Research Design

3.1 Data sources and sample construction

The initial sample of dual-class firms comes from Gompers, Ishii, and Metrick (2005). Since Gompers et al. only provide dual-class data from 1994 to 2002, I extended the dual-class sample to 2006. Specifically, I first constructed a list of possible dual-class firms over the 2003-2006 time period using data from the Securities Data Company, Compustat, CRSP, and the Investor Responsibility Research Center (IRRC).[9] Then I manually checked the proxy statements and/or 10-Ks for each candidate to determine whether it was indeed a dual-class firm. Subsequently, my initial sample covers dual-class firms over the period of 1994-2006. The rights and obligations of each class of common shares in a dual-class U.S. firm are spelled out in SEC filings, so investors can readily obtain information regarding the divergence of cash flow and voting rights. I collected insiders' ownership data for each class of common stock over the sample period. The yearly distribution of the resulting sample ("the initial dual-class sample") is given in Table 1.

The sample of dual-class firms is matched with single-class firms from the Compustat universe by year, industry, and size (proxied by total assets). I also obtained data from the following sources: the Compustat annual industrial and research files for accounting information, the CRSP files for stock prices and returns, and the IBES data

[9] A more detailed description of constructing the initial sample is provided in Gompers et al. (2005). In the process, I eliminated trusts, closed-end funds, ADRs, units, and REITs. I removed firms with one class of shares but extra voting rights through keyword search. When the votes per share of the two classes are equal (23 firms, 68 firm-year observations), I define the superior class as the class with the right to elect a disproportionate share of the board of directors. Consistent with Gompers et al. (2005), I also include a small number of firms that have more than two classes of common stock.

for raw earnings forecasts unadjusted for stock splits and actual earnings.[10] The main analyses further require data from S&P's ExecuComp database for executives' stock-based compensation and ownership data, and Thomson Financial's CDA/Spectrum database for institutional ownership data. While a majority of sample firms have ExecuComp compensation data, I hand collected compensation data from proxy statements for the remaining 151 firms in the sample.

Table 1, Panel A gives the distribution of the restricted sample by year. Specifically, the restricted dual-class sample consists of 496 unique dual-class firms along with 2,502 firm-year observations, and the same number of single-class firm-year observations. To mitigate the effects of influential observations, all test variables are winsorized at the 1st and 99th percentile values of their sample distributions. The final sample used for specific analyses varies due to additional data requirements and is discussed in the subsequent sections.

Panel B presents the sample distribution by industry groups. Dual-class firms are relatively concentrated in industries such as durable manufacturing, transportation, textiles and publishing, computers, and services. Thus, dual-class firms tend to operate in industries where a considerable amount of value is created by making investments in the pursuit of long-term goals.

3.2 Testing H1: Propensity of meeting or just beating analysts' forecasts

I use logistic regressions to test my first hypothesis, i.e., the propensity of meeting or just beating analysts' earnings forecasts for dual-class firms. Prior literature indicates

[10] Baber and Kang (2002), and Payne and Thomas (2003) find that analysts' forecast data in IBES are subject to split adjustment bias, in that these databases adjust stock splits retrospectively and split-adjusted data are rounded to the nearest cent. This measurement problem leads to the appearance of a greater proportion of zero forecast errors in the earlier years.

that managers have various incentives (e.g., Matsumoto 2002) and different abilities (Barton and Simko 2002; Rees 2005) to meet or beat analysts' forecasts. After controlling for these determinants, as well as for factors related to external monitoring, the marginal probability of meeting or beating forecasts should not differ significantly across single-class vs. dual-class firms, absent earnings management. The pooled cross-sectional and time-series regression is specified as follows:

$$\begin{aligned}\text{Prob}(MBE_{i,t} \in [\$0, \$0.01]) = F(&\alpha_0 + \alpha_1 DualProxy_{i,t} + \alpha_2 LnMV_{i,t} + \alpha_3 LnMB_{i,t} + \alpha_4 Lev_{i,t} \\ &+\alpha_5 Option_{i,t} + \alpha_6 Bonus_{i,t} + \alpha_7 MgmtOwn_{i,t-1} + \alpha_8 NOA_{i,t-1} + \alpha_9 RNOA_{i,t-1} \\ &+\alpha_{10} Loss_{i,t-1} + \alpha_{11} RDAdv_{i,t-1} + \alpha_{12} StdCFO_{i,t} + \alpha_{13} LnAnalysts_{i,t} \\ &+\alpha_{14} InstOwn_{i,t} + \alpha_{15} ExpMgmt_{i,t} + \sum \delta_T Year + \sum \delta_I Industry + \varepsilon_{i,t})\end{aligned} \quad (1)$$

where, for firm *i* in year *t* (unless otherwise indicated):

MBE = A dummy variable that equals 1 if the difference between actual earnings per share and consensus analysts' forecasted earnings per share is within [$0, $0.01], 0 otherwise.
DualProxy = Proxies for dual-class status, including *Dual* and *Wedge*.
LnMV = The natural logarithm of market value of equity at the fiscal year end.
LnMB = The natural logarithm of market value of equity divided by book value of equity.
Lev = Financial leverage.
Option = CEO's stock option incentives.
Bonus = Bonuses paid to the CEO scaled by firm-specific CEO wealth.
MgmtOwn = The percentage of stock holdings (including restricted stock) held by the top managers at the end of year *t*-1.
NOA = Net operating assets at year *t*-1 divided by lagged sales.
RNOA = Return on net operating assets.
Loss = A dummy variable of loss incidence at year *t*-1.
RDAdv = R&D and advertising expense at year *t*-1 scaled by sales.
StdCFO = Standard deviation of cash flow from operations, scaled by lagged total assets, over the rolling prior 5 years.
LnAnalysts = The natural logarithm of the number of analysts issuing annual earnings forecasts.
InstOwn = The percentage of total outstanding shares held by institutional investors.
ExpMgmt = A dummy variable that equals 1 if a firm misses an initial consensus forecast but meets or beats the most recent consensus forecast, and 0 otherwise.
Year = Year dummies.
Industry = 2-digit SIC industry dummies.

Earnings surprises are calculated as the difference between actual earnings and analysts' consensus forecasts, both of which are measured on a per-share basis. To better capture the market's expectation, I use consensus forecasts at the fiscal year end.[11] Consistent with Cheng and Warfield (2005), the dummy variable *MBE* equals 1 for a small earnings surprise falling within the range of [$0, $0.01], and 0 otherwise.

Key variables of interest. In the regression of *MBE* probability, the treatment variables are *Dual* and *Wedge*. *Dual* is a dummy variable that equals 1 if a firm has dual-class shares, 0 otherwise. *Wedge* equals the difference between voting and cash flow ownership by insiders (i.e., *VoteRights* - *CFRights*), where *VoteRights* is the total percentage of votes owned by officers and directors across classes, as reported in proxy statements; and *CFRights* is the total percentage of cash flow ownership by officers and directors.[12] The first hypothesis predicts negative coefficients after controlling for other determinants of *MBE* probability.

Managers' incentives for MBE. The first three variables, including *LnMV*, *LnMB*, and *Lev*, relate to managers' incentives to meet or beat analysts' forecasts.[13] *LnMV* controls for firm size since large firms have more media exposure and thus managers may experience greater pressure to meet or beat earnings expectations. *LnMB* proxies for growth potential (e.g., Smith and Watts 1992), and growth firms are penalized more

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[11] In additional analyses, I use the most recent consensus analyst forecasts within the three-month period prior to earnings announcements as the alternative proxy for earnings expectations and get qualitatively similar results.

[12] Separate share classes sometimes have different cash flow rights. In this study I assume cash flow rights per share to be equal across the two classes of shares. In the sensitivity tests, where the dividend data for each class is available from 1994 to 2002, I alternatively assume cash flow rights to be proportional to the ordinary dividends of that class, and the results remain unchanged.

[13] Given the skewed distribution of several control variables, such as market capitalization, and market-to-book ratio, log transformations are used when appropriate.

severely for negative earnings surprises (Skinner and Sloan 2002). Leverage positively affects a firm's likelihood of meeting or beating expectations due to managerial incentives to avoid covenant violations (e.g., Press and Weintrop 1990; DeFond and Jiambalvo 1994).

I use three measures to capture the effect of incentive compensation on managers' financial reporting behavior: CEO stock option compensation (*Option*); CEO bonus compensation (*Bonus*); and managerial ownership (*MgmtOwn*). Recent research suggests a positive association between accounting discretion and stock options granted to executives (e.g., Cheng and Warfield 2005). Moreover, several studies (e.g., Healy 1985; Matsunaga and Park 2001) find that compensation plans that pay bonuses based on accounting performance are positively correlated with income-increasing accounting choices in periods when a firm's accounting income falls below earnings targets. Finally, managerial ownership is expected to be negatively related to accounting discretion (Warfield, Wild, and Wild 1995). Inclusion of *MgmtOwn* is intended to disentangle the possible incentive alignment effect of concentrated managerial ownership from the entrenchment effect of the wedge.

Managers' abilities for MBE. Managers' abilities to meet or beat earnings targets are proxied by *NOA*, *RNOA*, *Loss*, *RDAdv*, and *StdCFO*. Barton and Simko (2002) find that the likelihood of reporting positive earnings surprises decreases with overstated net asset values. They reason that the level of net operating assets from the prior year (NOA_{t-1}) partly reflects accumulations of previous accruals management, which curtails managers' ability to manage earnings upward in the current period. I include *RNOA* to control for the possibility that firms with better performance are more likely to meet or

beat analysts' forecasts. Prior research also suggests that meeting or beating market expectations is less important for firms that incur losses (Degeorge et al. 1999), so I include the incidence of loss from prior year (*Loss*) in the *MBE* probability model. In addition, due to the interaction of conservative accounting practices and investment activity, the intensity of R&D expense from prior years is correlated with hidden earnings reserves. I expect that firms with high R&D intensity are more capable of meeting or beating market expectations. Finally, I add cash flow volatility (*StdCFO*) and expect it to be negatively related to *MBE*.

External monitoring. External monitoring mechanisms provide independent monitoring of management, and promote more transparent financial reporting. This category consists of two variables, *LnAnalysts* and *InstOwn*. Stronger analyst following (*LnAnalysts*) and higher institutional ownership (*InstOwn*) reflect shareholder interest in a firm, which may exert pressure on managers to meet or beat analysts' earnings forecasts (Matsumoto 2002). However, higher institutional ownership may provide a higher degree of monitoring since institutional investors are cognizant that an increase in earnings achieved by manipulating accruals (Mitra and Cready 2005) or by cutting R&D (Bushee 1998) would not be sustainable. Thus, I expect the coefficients on *LnAnalysts* in Equation (1) to be positive, and the coefficient on *InstOwn* to be negative.

Expectation management. Prior research (Matsumoto 2002; Richardson, Teoh, and Wysocki 2004; Cotter, Tuma, and Wysocki 2006) provides evidence that firms can also meet or beat market expectations through downward guidance of analysts' forecasts. To control for that possibility, I create a dummy variable, *ExpMgmt*, which is based on the direction of the overall revision of analysts' forecasts (e.g., Bartov et al. 2002).

Lastly, I introduce two-digit industry dummies (*Industry*) and time dummies (*Year*) to account for industry- and time-specific factors that affect the overall likelihood of meeting or beating analysts' forecasts.

3.3 Testing H2: Propensity of managing accruals

To investigate how share structures affect accruals quality, I include factors related to the costs and incentives of accruals management, as well as to firms' governance attributes, in pooled cross-sectional and time-series regressions:

$$\begin{aligned}|AbnAccr_{i,t}| = {} & \beta_0 + \beta_1 DualProxy_{i,t} + \beta_2 LnMV_{i,t} + \beta_3 LnMB_{i,t} + \beta_4 Lev_{i,t} + \beta_5 StdCFO_{i,t} \\ & + \beta_6 Option_{i,t} + \beta_7 Bonus_{i,t} + \beta_8 MgmtOwn_{i,t-1} + \beta_9 ExtFin_{i,t} + \beta_{10} MBE_{i,t} \\ & + \beta_{11} NOA_{i,t-1} + \beta_{12} Loss_{i,t-1} + \beta_{13} RDExp_{i,t-1} + \beta_{14} Herf_{i,t} + \beta_{15} Segment_{i,t} \\ & + \beta_{16} LnOpCyc_{i,t} + \beta_{17} InstOwn_{i,t} + \beta_{18} Auditor_{i,t} \\ & + \sum \delta_T Year + \sum \delta_I Industry + \varepsilon_{i,t}\end{aligned} \quad (2)$$

where, for firm *i* in year *t* (unless otherwise indicated):

|AbnAccr|	=	Absolute value of abnormal accruals. The details are given in the footnotes to Table 2.
ExtFin	=	External financing dependence.
Herf	=	Industry concentration (i.e., the Herfindahl index).
Segment	=	Operating segment diversification.
LnOpCyc	=	The natural logarithm of the length of the operating cycle.
Auditor	=	A dummy variable that is set to 1 if the firm is audited by a Big-5 audit firm or successor, 0 otherwise.

All other variables are defined as before.

To gauge the level of accruals management, I use a modified Jones' model (Dechow, Sloan, and Sweeney 1995), but adjust accruals by performance as proposed by Kothari, Leone and Wasley (2005). My interest is in determining whether dual-class firms exhibit lower abnormal accruals relative to single-class firms after controlling for other determinants of accruals quality noted in prior literature. Given these controls, I assume that the only remaining influence on the difference in accruals quality derives from differences in firms' share structures. A negative coefficient estimate on *Dual* (or

Wedge) would be consistent with Hypothesis 2.

Managers' incentives for accruals management. Following the same logic in the *MBE* probability model, I control for firm size (*LnMV*), market-to-book ratio (*LnMB*), leverage (*Lev*), and variables about incentive compensation (*Option*, *Bonus*, and *MgmtOwn*).

Hribar and Nichols (2007) emphasized the need to control for operating volatility metrics in regressions when absolute abnormal accruals are the dependent variable. I proxy for operating volatility using the standard deviation of operating cash flows (*StdCFO*). Prior research suggests that demand for external financing induces managers to manipulate reported earnings numbers so as to boost stock prices when they issue debt or equity (e.g., Frankel, McNichols, and Wilson 1995; Teoh, Welch, and Wong 1998). Thus, I include firms' dependence on external financing (*ExtFin*) as used by Kaplan and Zingales (1997). Finally, to control for firms' incentive to meet or beat an earnings benchmark, I follow Roychowdhury (2006) and include the dummy *MBE*, which equals one when a firm's earnings surprise falls in the range of [\$0, \$0.01].

Managers' abilities for accruals management. I control for the following variables related to management abilities to manipulate accruals: *NOA*, *Loss*, *RdAdv* (the three variables used in the *MBE* probability model), *Herf*, *Segment*, and *LnOpCyc*.

Managers in industries with less product market competition have more leeway to manage earnings. So I expect abnormal accruals to be positively related to the industry Herfindahl index (*Herf*).[14] Firms with more complex operations are posited to have larger abnormal accruals due to both intentional and unintentional measurement problems

[14] The Herfindahl index ranges from one in the case of pure monopoly to zero in the case of perfect competition. A higher index suggests higher barriers to entry.

associated with cost allocation, transfer pricing, and the elimination of inter-segment sales. Following Givoly, Hayn, and D'Souza (1999), I include number of operating segments (*Segment*) to capture the effect of a firm's operating complexity. Finally, longer operating cycles (*LnOpCyc*) induce more uncertainty (Dechow and Dichev 2002), and more accruals may be utilized to reduce the timing and matching problems of cash flows (Dechow 1994).

Key governance attributes. Similar to the *MBE* probability model, I control for *InstOwn*. In addition, I control for auditing reputation (*Auditor*). Prior research (e.g., Becker, DeFond, Jiambalvo, and Subramanyam 1998) finds that Big-Five audit firms help to ensure the reliability of financial information as they are the most experienced firms in their field, and invest more resources in auditing.

3.4 Testing H3: Propensity of releasing earnings reserves

Consistent with prior studies such as Penman and Zhang (2002), actual earnings reserves (*ER*) consist of three components: inventory reserve, R&D reserve, and advertising reserve. To examine managers' opportunistic tendency to release earnings reserves, I first come up with a measure of expected earnings reserves by accounting for their innate determinants. The following regression is estimated for all firms with available data on the Compustat universe for every industry with at least 10 observations in year *t*:

$$ER_{j,t} = \gamma_0 + \gamma_1 ER_{j,t-1} + \gamma_2 \Delta ER_{j,t-1} + \gamma_3 GSales_{j,t} + \gamma_4 RNOA_{j,t} + \gamma_5 FCF_{j,t-1} + \gamma_6 TobinQ_{j,t} + \gamma_7 Age_{j,t} + \varepsilon_{j,t} \quad (3)$$

where, for firm *j* in year *t* (unless otherwise indicated):

ER = Measured as (inventory reserve + R&D reserve + advertising reserve) / net operating assets, and also known as "C-score" in Penman and Zhang (2002). ΔER is the annual change in earnings reserves.

GSales = Annual growth in sales.

FCF = Free cash flow.
TobinQ = Tobin's Q.
Age = Firm's age.

In the model of expected level of earnings reserves, I control for lagged reserves and annual change of reserves. Firms with better performance are assumed to be less capital-constrained and have more funds available for R&D and other investment projects. So I include sales growth (*GSales*), return on net operating assets (*RNOA*), and free cash flow from the prior year (FCF_{t-1}). Further, firm-specific growth potential, as measured by Tobin's Q (*TobinQ*), determines the current level of earnings reserves. Finally, a firm's life cycle (proxied by *Age*) also affects earnings reserve since young firms tend to invest more in R&D. Abnormal earnings reserves (*AbnER*) are the difference between actual reserves and expected reserves estimated from Equation (3).[15]

In the second stage, I test the likelihood of releasing abnormal earnings reserves (i.e., $AbnER_{i,t} < 0$) related to dual-class structures. The logistic regression is specified as follows:

$$\begin{aligned}\text{Prob}(AbnER_{i,t} < 0) = F(&\kappa_0 + \kappa_1 DualProxy_{i,t} + \kappa_2 Option_{i,t} + \kappa_3 Bonus_{i,t} \\ &+\kappa_4 MgmtOwn_{i,t-1} + \kappa_5 ExtFin_{i,t} + \kappa_6 MBE_{i,t} + \kappa_7 Lev_{i,t} \\ &+\kappa_8 StdCFO_{i,t} + \kappa_9 NOA_{i,t-1} + \kappa_{10} InstOwn_{i,t} \\ &+\sum \delta_T Year + \sum \delta_I Industry + \varepsilon_{i,t})\end{aligned} \quad (4)$$

In Equation (4), the capital market pressure effect will manifest itself in a negative coefficient on *Dual* (*Wedge*), after controlling for managerial incentives, costs of manipulating earnings reserves, and external monitoring. Managers may release earnings reserves to boost current earnings driven by the following incentives: (1) incentive

[15] Abnormal earnings reserves, as derived in this study, are essentially a modified version of Q-score in Penman and Zhang (2002). This is to ensure that any decrease of earnings reserves derives from managers' opportunistic behavior, and is not due to firm-specific and industry-specific factors.

compensation (measured by *Option*, *Bonus*, and *MgmtOwn*); (2) demand for external financing (measured by *ExtFin*); (3) meeting or beating analysts' forecasts (measured by *MBE*); and (4) avoiding excess volatility and creating a smooth earnings stream (measured by *Lev* and *StdCFO*). Moreover, managers are more likely to release earnings reserves when the accruals management constraint (proxied by NOA_{t-1}) is binding and firms are compelled to look for other means to boost earnings. Finally, I expect that the propensity to release earnings reserves will be positively related to external monitoring, proxied by *InstOwn*.

4. Main Results

4.1 Descriptive statistics

Table 2 presents descriptive statistics for dual-class firms, and the matching single-class group separately. Panel A reports the statistics on cash flow rights and voting rights of dual-class stocks. On average, insiders hold 40.4 percent of cash flow rights, and 61.7 percent of voting rights. Both the mean and median insider cash flow rights are significantly lower than those of voting rights. The resulting mean *Wedge* is 0.213 and the median *Wedge* is 0.205. Moreover, dual-class firms have a slightly higher dividend yield, yet utilize more leverage than matching single-class firms (in Panel B). Dual-class firms appear to rely more on internal retained earnings and external debt financing, but are less likely to tap the market for new equity capital to fund further expansion and growth (e.g., Amoako-Adu and Smith 2001).

From Panel C, dual-class firms perform better in terms of *RNOA* than single-class firms (mean of 12 percent versus 9.6 percent), with both mean and median differences significant at the 1% level. Similarly, the mean (median) annual sales growth for the

dual-class firms is 12.3 percent (11.5 percent) compared to 10.9 percent (10.6 percent) for the single-class firms, the differences of which appear to be economically significant. Abnormal accruals represent 2.1 percent of lagged assets for the average dual-class firm and 2.8 percent of lagged assets for the average single-class firm with the mean difference significant at the 1% level. But dual-class firms have slightly more volatile cash flows. Moreover, average R&D and advertising expenses are 5.7 percent of revenue for dual-class firms, compared to 4.3 percent for matching single-class firms, which implies that single-class firms spend around 30 percent less than dual-class firms on R&D. The mean earnings reserve for dual-class firms is 0.293, compared to 0.249 for single-class firms. The differences between the two groups are statistically significant at the 1% level as indicated by t-statistic of mean difference and z-statistic of median difference. These results are consistent with prior research (DeAngelo and DeAngelo 1985) that without undue capital market pressures, managers in dual-class firms are more willing to invest in long-term projects compared to their single-class counterparts.

Panel D shows that the average market value is approximately $4.8 billion for dual-class firms, and slightly more than $5.1 billion for single-class sample firms. Dual-class stocks also suffer from reduced trading liquidity. The mean and median differences of stock returns across dual-class and single-class firms are not statistically significant, so univariate tests do not suggest that one group of firms outperforms the other group in the stock market.

Panel E reports descriptive statistics related to external monitoring and incentive compensation. Sample dual-class firms on average have about 24.6 percent of institutional ownership, compared to 27.8 percent for single-class firms, confirming that

some institutional investors disfavor dual-class structures.[16] Similarly, the average dual-class firm is followed by about 4.289 analysts, fewer than 6.042 analysts for single-class firms. Likewise, a smaller percentage of dual-class firms (93.8 percent vs. 97.8 percent for single-class firms) are audited by Big-Five audit firms. These results indicate that external monitoring of dual-class firms is weaker than monitoring of single-class firms. With respect to incentive compensation, both the mean and median of *Option* are smaller for dual-class firms (0.246 and 0 respectively) than for single-class firms (0.414 and 0.221 respectively) and the differences are significant at the 1% level. Dual-class firms also seem to be less likely to utilize cash bonuses for their managers. Finally, dual-class firms enjoy higher levels of managerial ownership. The mean (median) percentage of shares held by top managers is 5.8 percent (4.9 percent), while the mean (median) managerial ownership for single-class firms is 4.2 percent (3 percent).

The above analyses highlight salient features of dual-class firms relative to their single-class counterparts. Dual-class firms are more leveraged, and exhibit higher cash flow volatility. They use less stock-based compensation, but have higher levels of managerial equity ownership. They enjoy better operating performance, but are less monitored by outside shareholders. Most importantly, dual-class firms spend more on R&D, and boast higher levels of earnings reserves.

4.2 Results of H1: Propensity of meeting or just beating analysts' forecasts

I first provide descriptive evidence of irregularities in the distributions of earnings surprise. To test whether the empirical distribution of earnings surprises around the zero threshold is smooth, I adopt Burgstahler and Dichev's (1997) analysis by calculating the

[16] In practice, some institutional investors (e.g., TIAA-CREF and CalPERS) and corporate governance advocacy groups (e.g., Institutional Shareholder Service) view dual-class structures as providing low managerial accountability, and publicly oppose such structures.

standardized difference for the two intervals adjacent to zero earnings surprise.[17] Earnings surprises follow a smooth distribution when earnings management is absent. In that case, the standardized differences (z-statistics) will be distributed normal.

As shown in Panel A of Table 3, across both samples, more firm-years meet or just beat analysts' forecasts rather than just miss forecasts. Of 2,502 observations in the dual-class sample, 19.6 percent meet or just beat analysts' forecasts (two intervals just above zero), and 15.7 percent miss analysts' forecasts by only one or two cents. In contrast, 22.3 percent of the matching single-class firms meet or just beat analysts' forecasts, while only 11.8 percent miss analysts' forecasts by a small margin (two intervals just below zero). Single-class firms exhibit fewer small negative earnings surprises than expected (standardized difference of -4.86), but many more small positive earnings surprises (standardized difference of 3.41). The distribution of earnings surprises for the single-class sample exhibits a positive discontinuity around the threshold of zero earnings surprise, with both test statistics significant at the 1% level. By comparison, standardized differences of dual-class firms for intervals marginally below zero or marginally above zero are not significantly different from zero.

Table 3, Panel B presents results of the conditional logistic model (Equation 1). In Model 1, the probability of small positive earnings surprises is negatively and significantly associated with the dual-class dummy (*Dual*). In Model 2, the coefficient on *Wedge* is negative and significant at the 1% level. The pseudo R^2 increases slightly from 23.2 percent to 23.8 percent. Thus, my results support the capital market pressure

[17] The standardized difference for an interval is the difference between the observed and expected number of observations in the interval, standardized by the estimated standard deviation of the difference. The expected number of observations in each interval is the average of the number of observations in the two immediately adjacent intervals. I calculate the standard deviation of the standardized difference using the entire distribution of earnings surprises.

hypothesis. Furthermore, untabulated marginal effects suggest that one standard deviation increase of *Wedge* is associated with a reduction of the *MBE* probability by 5.3 percent, which is economically significant.

Results for the other determinants of *MBE* probability are generally consistent with those reported in prior research. Specifically, large firms, firms with high growth potential, firms with high leverage, firms with high *RNOA*, firms with higher R&D and advertising expenses, firms that award managers with more stock options and bonuses, firms followed by large numbers of sell-side analysts, and firms managing expectations, are more likely to meet or just beat analysts' forecasts. Conversely, firms with large net operating assets in the previous year, and firms that suffer losses from the prior year, are negatively related to small positive earnings surprises. Managerial ownership is positively related to the probability of meeting or just beating analysts' forecasts, but the relationship is not statistically significant. Both analyst following and institutional ownership exhibit positive associations with the *MBE* probability, but only the coefficient on *LnAnalysts* is significant. Finally, inconsistent with my expectation, financial leverage is negatively associated with the *MBE* probability.

For typical single-class firms, managers appear predisposed to managing earnings downward in years of good performance, which leads to increased reserving of current earnings to avoid future earnings disappointments (Cheng and Warfield 2005). So, in the untabulated analyses, I also run the logit regression of the probability of large positive earnings surprises (i.e., Prob ($MBE_{i,t} \geq \$0.04$) on dual-class structures and control variables. I find that *Dual* is positively, instead of negatively, associated with the probability of large positive earnings surprises. This piece of evidence corroborates the

notion that earnings at dual-class firms are less predictable, and managers at these firms are long-term oriented.

To summarize, the evidence in this section strongly supports the hypothesis that dual-class firms have a reduced likelihood of meeting or slightly beating analysts' earnings forecasts, even after controlling for the possibility of expectation management. My findings also confirm the notion that capital market pressures lead to managers' attempts to meet or just beat market expectations in single-class firms.

4.3 Results of H2: Propensity for accruals management

Table 4, Panel A classifies abnormal accruals according to the two outcomes (i.e., meeting or beating by small margins vs. all others), along with statistics from t-tests and Wilcoxon tests of mean and median differences across the two groups. The first row indicates that both mean and median abnormal accruals are not statistically different across the MBE and "all others" groups within the dual-class sample. In contrast, the second row shows that for single-class firms, both the mean and median of *AbnAccr* in the *MBE* group are significantly larger relative to those in the "all others" group, with t-statistic and z-statistic indicating highly significant mean and median differences. Thus, the descriptive evidence from Panel A confirms the existence of accruals management to meet or beat market expectations among single-class firms, while dual-class firms do not appear to exhibit similar opportunistic behaviors.

Panel B of Table 4 presents results of multivariate analyses for Hypothesis 2. For all model estimations, I report t-values based on Huber-White robust standard errors that control for heteroscedasticity and correlations in the error terms. In Model 1, the dummy variable *Dual* is negatively associated with the absolute value of abnormal accruals. The

coefficient estimate, at -0.224, is significant at p-value < 0.05. In Model 2, when I use *Wedge* in the regression, the resulting coefficient on *Wedge* is negative and significant at the 10% level.[18]

While many correlations between independent variables are significant, none of the variance inflation factors (VIF) of the independent variables is greater than 2.5, mitigating the concern of multicollinearity among control variables. The signs of coefficients on the control variables are generally consistent with expectation and prior literature. Specifically, the results from Panel B confirm that large firms, firms with high book-to-market ratios, more volatile cash flows, higher R&D and advertising expenses, more operating diversification, longer operating cycles, and firms that award managers with more stock options and bonuses, take less abnormal accruals. Consistent with the incentive alignment effect, managerial ownership is negatively correlated with abnormal accruals. And consistent with strong incentives to meet or just beat market expectations, the coefficient on *MBE* is significantly positive in the model.[19] However, the coefficients on *ExtFin* (dependence on external financing), *NOA*, *Loss* and *Herf* (industry concentration) are not statistically significant. In addition, the coefficient on *Lev* is negative and significant at the 1% level even though it is predicted to be positive.[20] Finally, the coefficients on *InstOwn* and *Auditor* are negative and significant at the 10% and 5% levels, consistent with external monitoring roles played by institutional investors and auditors.

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[18] In robustness checks, I use signed abnormal accruals as the dependent variable in the regressions and obtain qualitatively similar results.

[19] In further tests, I construct two variables that interact MBE with dual-class and single-class dummies respectively (i.e., *MBE*Dual* and *MBE*Single*). As expected, the coefficient on *MBE*Dual* is insignificant, while the coefficient on *MBE*Single* is positive and significant.

[20] Cheng and Warfield (2005) also report a negative relationship between leverage and abnormal accruals.

In sum, the empirical evidence in Table 4 is consistent with the capital market pressure hypothesis. Dual-class share structures diminish managerial propensity to opportunistically manage accruals. They further indicate that the absolute level of abnormal accruals is approximately a third lower for dual-class firms than for matching single-class counterparts.

4.4 Results of H3: Propensity of releasing earnings reserves

Table 5, Panel A estimates the expected level of earnings reserves. While lagged earnings reserve is the most significant determinant of current level of earnings reserve, sales growth, return on *NOA*, Tobin's Q, and firm's age also marginally explain current earnings reserves. Based on coefficients estimates from Model 2, I obtain expected earnings reserves for sample firms. Abnormal reserves are the difference between actual reserves and expected reserves.

Panels B and C of Table 5 present univariate and multivariate results of releasing abnormal earnings reserves across dual-class and single-class firms. Panel B gives the descriptive statistics of abnormal earnings reserves (*AbnER*) for firms that meet or just beat analysts' forecasts, compared to all other firms. For dual-class firms with small positive earnings surprises (i.e., $MBE \in [\$0, \$0.01]$), the mean (median) *AbnER* is 0.011 (-0.001), compared to the mean (median) *AbnER* of 0.012 (0.003) for all other dual-class firms. The resulting mean and median differences are not statistically significant at conventional level. In contrast, among single-class firms, the mean (median) *AbnER* is -0.009 (-0.006) when there are small positive earnings surprises, compared to the mean (median) *AbnER* of 0.011 (0.003) for all other single-class firms. The resulting mean and median differences between the two single-class samples are statistically significant at

the 5% level. Negative *AbnER* in response to meeting or beating analysts' consensus suggests that single-class firms release earnings reserves. From the descriptive evidence in Panel B, dual-class firms do not appear to release earnings reserves in order to meet or exceed market expectations.

Results from the logit regressions shed further light on the use of earnings reserves from prior periods to meet or beat market expectations. Model 1 of Panel C uses the dummy variable (*Dual*), along with other control variables, to explain the probability of $AbnER < 0$. The coefficient on *Dual* is significantly negative at -0.075 (p-value < 0.01). In Model 2, the coefficient on *Wedge* is negative at -0.082 and significant at the 1% level, with the marginal probability of 7.9 percent (untabulated). Combined, the results are consistent with Hypothesis 3. Dual-class firms are less likely to release reserves for the purpose of reporting higher short-term earnings.

Regarding the control variables, firms with more incentive compensation in terms of *Option* and *Bonus*, higher dependence upon external financing, strong incentive to meet or beat analysts' forecasts (*MBE*), and higher risk and volatility (*Lev* and *StdCFO*), tend to release earnings reserves.[21] I also find that the coefficient on *InstOwn* is negative and significant at the 1% level, consistent with existing research evidence (e.g., Bushee 1998; Bange and DeBondt 1998) that managers with higher total institutional ownership are less likely to release earnings reserves (i.e., by cutting R&D expense) to meet earnings targets. However, the coefficient on prior-period net operating assets is not significant at a conventional level. Taken together, the results show that unlike matching

[21] Similar to the test of accruals management, I also construct two variables that interact *MBE* with dual-class and single-class dummies respectively (i.e., *MBE*Dual* and *MBE*Single*). As expected, the coefficient on *MBE*Dual* is insignificant, while the coefficient on *MBE*Single* is positive and significant.

single-class firms, dual-class firms are less likely to respond to external pressure by releasing earnings reserves to reach for higher current earnings.

5. Additional Tests of the Capital Market Pressure Hypothesis

In this section, I provide additional evidence in support of the capital market pressure hypothesis. Specifically, I show that dual-class stocks attract a higher percentage of long-term institutional investors. Further confirming reduced capital market pressure for dual-class firms, when both dual-class and single-class firms miss analysts' forecasts, abnormal returns around the time of earnings announcements for dual-class firms are significantly less negative than those for their single-class counterparts. Moreover, as a consequence of long-term oriented corporate policies, dual-class firms enjoy higher operating and stock return performance. Finally, I repeat tests on the sample of dual-class firms only and on samples of firms that either initiated or abolished dual-class structures during the sample period.

5.1 Shareholder base of dual-class firms

Institutional investors are pivotal in the trading and pricing of stocks (Gompers and Metrick 2001). Institutional investors with different investment horizons may have different incentives to monitor firms' financial reporting practices. Short-term institutional investors, who exhibit high portfolio turnover, are more interested in a firm's short-run performance and thus create pressures on managers to meet short-term earnings targets (Bushee 1998, 2001).[22] In contrast, long-term institutional investors are willing to

[22] Prior studies provide evidence consistent with the notion that short-term institutional investors exert capital market pressures. Hotchkiss and Strickland (2003) find that stock price response to negative earnings news is more negative for firms with higher short-horizon institutional ownership. Shin (2005) finds that the boards of firms with higher short-term institutional ownership punish CEOs more severely in their annual bonuses when firms miss earnings targets.

engage in costly monitoring because they are likely to remain a firm's shareholders and to reap the corresponding benefits from monitoring (Froot, Perold, and Stein 1992). Therefore, managers will have a stronger (weaker) incentive to manipulate earnings when the firm has higher ownership by short-term (long-term) institutional investors.

To characterize shareholder types related to shareholder horizons, I use two firm-specific measures: the percentage of short-term and long-term institutional ownership (Bushee 1998) and a weighted average turnover measure of institutional investors (Gaspar, Massa, and Matos 2005). Bushee (1998) classifies institutional investors into transient, quasi-indexer, and dedicated institutional investors based on investment horizon and portfolio diversification.[23] The latter two groups of institutions are considered to be long-term institutional investors (*LTInstOwn*), while short-term institutional ownership (*STInstOwn*) consists of the percentage of shares held by transient institutional investors. Alternatively, the measurement of institutional investor turnover (*InstTurn*) is detailed in the footnote to Table 6.

Table 6, Panel A presents summary information related to the types of institutional ownership. The mean (median) institutional investor turnover is 22.4 percent (21.8 percent) for dual-class firms, while the mean (median) institutional turnover for single-class firms is 26.6 percent (26.2 percent), significantly higher than that of dual-class firms. Moreover, for a typical dual-class firm, long-term institutional investors own 18.4 percent of shares, while short-term institutional investors on average own only 6.2 percent of shares. In other words, the shareholder base of a typical dual-class firm

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[23] Transient institutional investors are characterized as having high portfolio turnover and highly diversified portfolio holdings. Quasi-indexing institutional investors have low turnover and diversified portfolios. Dedicated institutional investors have low turnover and more concentrated portfolios. I thank Brian Bushee for providing his institutional investor classification data.

consists of about 75 percent long-term institutional investors, and 25 percent short-term institutional investors. In comparison, about 34.5 percent of institutional investors who own single-class stocks are short-term oriented.

Next, based on a large literature examining the investment preferences of institutional investors with different investment horizons (e.g., Hessel and Norman 1992; Gompers and Metrick 2001), I specify the model of institutional investor turnover as follows:

$$\begin{aligned} InstTurn_{i,t} = {} & \lambda_0 + \lambda_1 Dual_{i,t} + \lambda_2 LnMV_{i,t-1} + \lambda_3 Liq_{i,t-1} + \lambda_4 LnMB_{i,t-1} \\ & + \lambda_5 RNOA_{i,t-1} + \lambda_6 Lev_{i,t-1} + \lambda_7 Beta_{i,t} + \lambda_8 IRisk_{i,t} + \lambda_9 Ret_{i,t-1} \\ & + \lambda_{10} Div_{i,t-1} + \sum \delta_T Year + \sum \delta_I Industry + \varepsilon_{i,t} \end{aligned} \tag{5}$$

where, for firm *i* at year *t*:

Liq	=	Liquidity of stock trading.
IRisk	=	Idiosyncratic stock return volatility.
Ret	=	Yearly buy-and-hold stock returns.
Div	=	Dividend yield.

All other variables are defined as before.

In Equation (5), I use two variables to proxy for institutional investors' preference for stocks with high levels of information and liquidity: size (*LnMV*) and liquidity (*Liq*). Institutional investors tend to invest in large firms, partly due to prudent-person standards (Del Guercio 1996). Moreover, for institutional investors with high portfolio turnover, stocks with greater liquidity are more attractive because the price impact of trading does not erode potential trading gains for such stocks (Falkenstein 1996). I also control for firms' growth prospects (*LnMB*), since many institutional investors follow either "value" or "growth" trading strategies. Likewise, return on net operating assets (*RNOA*) and financial leverage (*Lev*) serve as proxies on the safety of investment. I also control for prior-year stock returns (*SAR*) since some institutional investors follow the momentum

trading strategy (Grinblatt, Titman, and Wermers 1995). Furthermore, I include market beta (*Beta*) and idiosyncratic return volatility (*IRisk*) because there is evidence that institutional investors avoid stocks with high idiosyncratic volatility and tend to invest in high beta stocks (Bushee 2001). Finally, I control for dividend yield (*Div*), because institutional investors generally prefer stocks with low dividend yields and high capital gains potential (Cready 1994).

After controlling for all of these variables, I find that *InstTurn* is negatively associated with *Dual*, suggesting that dual-class firms attract a disproportionate presence of long-term investors. Separately, results in Panel B of Table 6 indicate that firms dominated with short-term institutional investors are usually smaller, more liquid, growth-oriented, have exhibited better performance in the past, and offer lower dividend yields. High institutional turnover is also associated with large systematic risk and high idiosyncratic volatility.

Recall that the regressions reported in Tables 3-5 use an aggregate measure of institutional ownership (*InstOwn*), but do not differentiate investors' horizons. It is possible that this pooling masks external pressures from short-term investors. I thus estimate Equations (2)-(4) again conditional on the investment horizons of institutional investors. Specifically, I investigate how the percentage of shares held by short-term and long-term institutional investors and the average turnover measure of institutional investors' entire portfolios are associated with firms' financial reporting quality. Untabulated results confirm that firms dominated by short-term institutional investors tend to actively manipulate earnings, while firms dominated by long-term institutional investors are associated with less earnings management. Furthermore, the effect of

monitoring from long-term institutional investors seems to be stronger among dual-class firms.

To summarize, a clientele effect exists such that dual-class firms tend to attract long-term institutional investors, who exert less short-term pressures but constrain managerial opportunism.

5.2 Abnormal returns around earnings announcements

In this section, I show that market reaction to missing expectations is significantly less negative for dual-class firms, rendering support to my proposition of reduced capital market pressures in dual-class firms.

I first compute univariate statistics for firms that report either good earnings news or bad earnings news. Within a sub-sample of firms that miss analysts' forecasts, the mean cumulative three-day market-adjusted return is –1.6 percent for dual-class firms, while it is –3.8 percent for single-class firms (untabulated). The difference in market reaction to negative earnings announcements between dual-class and single-class firms is economically significant. In contrast, the difference in market reaction to positive earnings news is not statistically significant between dual-class and single-class firms.

Following prior literature (e.g., Collins and Kothari 1989; Easton and Zmijewski 1989; Skinner and Sloan 2002), I control for unexpected earnings and other determinants that affect market reaction to earnings surprises, such as size (*LnMV*), leverage (*Lev*), growth prospect (*LnMB*), and systematic risk (*Beta*). I estimate the following model on the sample of dual-class and single-class firms:

$$
\begin{aligned}
CAR_{i,t} = {} & \theta_0 + \theta_1 Dual_{i,t} + \theta_2 Good_{i,t} + \theta_3 Bad_{i,t} + \theta_4 Dual_{i,t} * Good_{i,t} + \theta_5 Dual_{i,t} * Bad_{i,t} \\
& + \theta_6 UE_{i,t} + \theta_7 Good_{i,t} * UE_{i,t} + \theta_8 Bad_{i,t} * UE_{i,t} + \theta_9 Dual_{i,t} * Good_{i,t} * UE_{i,t} \\
& + \theta_{10} Dual_{i,t} * Bad_{i,t} * UE_{i,t} + \theta_{11} LnMV_{i,t} + \theta_{12} LnMB_{i,t} + \theta_{13} Lev_{i,t} + \theta_{14} InstOwn_{i,t} \\
& + \theta_{15} LnMV_{i,t} * UE_{i,t} + \theta_{16} LnMB_{i,t} * UE_{i,t} + \theta_{17} Lev_{i,t} * UE_{i,t} + \theta_{18} Beta_{i,t} * UE_{i,t} \\
& + \sum \delta_T Year + \sum \delta_I Industry + \varepsilon_{i,t}
\end{aligned} \tag{6}
$$

where, for firm *i* at year *t*:

CAR = Three-day market-adjusted stock returns.
UE = The difference between actual earnings per share and consensus forecasts, then scaled by stock price at the fiscal year end.
Good = 1 if *UE* is positive, and 0 otherwise.
Bad = 1 if *UE* is negative, and 0 otherwise.
Beta = Market beta.
All other variables are defined as before.

Table 7 reports results from pooled cross-sectional and time-series OLS regressions. Model 1 presents a baseline model, where I examine the relation between abnormal returns and unexpected earnings without controls. Model 2 controls for various determinants of market reaction. Since the results from Models 1 and 2 are similar, I focus on the results from Model 2. Consistent with prior studies, the negative coefficient on the dummy *Bad* indicates that disappointing earnings news elicits strong negative market reactions. Moreover, investors react towards the magnitude of earnings surprise, as captured by the variable *UE*. Although the coefficient on *Dual* is insignificant, the coefficient on *Dual*Bad* is significantly positive at 0.014. For sample firms that miss market expectations, the above results suggest that abnormal returns for dual-class firms are 1.4 percent less negative than those for single-class firms.

In summary, the multivariate regression analyses confirm that after controlling for various determinants of market response to earnings announcements, the market response for dual-class firms is more subdued. While this particular finding is consistent with Francis et al. (2005), my result points to a clientele that focuses more on dual-class firms'

long-term performance.[24]

5.3 Future operating and stock return performance of dual-class firms

Several recent studies (e.g., Bhojraj, Hribar, and Picconi 2005; Gunny 2005) document that using accruals or discretionary expenditure (including R&D expenses) to meet or beat analysts' forecasts results in long-term underperformance relative to firms that do not manage earnings. If managers at dual-class firms indeed focus on the long-term, I expect that managers at dual-class firms will deliver stronger firm performance than managers who succumb to short-term market pressures.

Univariate results in Section 4.1 provide only weak evidence of superior performance by dual-class firms. However, univariate analyses do not take into account other factors that may potentially affect a firm's long-term performance. I thus conduct multivariate regression tests where I examine the association between dual-class proxies and subsequent performance. Given that new IPO firms often sustain losses in the first few years of going public and have a higher likelihood of bankruptcy (Fama and French 2004), I require each firm-year observation to be at least five years after a firm's IPO. Specifically, I perform the following pooled cross-sectional and time-series regressions:

$$RNOA_{i,t+k} = \eta_0 + \eta_1 DualProxy_{i,t} + \eta_2 RNOA_{i,t} + \eta_3 LnMV_{i,t} + \eta_4 LnMB_{i,t} + \eta_5 Accr_{i,t} + \eta_6 AbnER_{i,t} + \sum \delta_T Year + \sum \delta_I Industry + \varepsilon_{i,t} \quad (7)$$

where, for firm i in year t, $RNOA_{i,t+k}$ equals to one-year ahead and three-year ahead (k = 1, 3) return on net operating assets; and $Accr$ is total accruals scaled by lagged total assets. All other variables are defined as before.

[24] As an alternative explanation, earnings disappointments for dual-class firms may be viewed by investors as being more transitory and thus have smaller price responses. I thank Ray Pfeiffer for making this point.

In Equation (4), the variables of interest are *Dual* and *Wedge*. I examine the incremental effects of dual-class structures on future profitability after controlling for factors that are likely to be associated with future performance. I control for current-period *RNOA*, because firms' profitability is strongly correlated over time (e.g., Freeman, Ohlson, and Penman 1982). Firm size (*LnMV*), and market-to-book ratio (*LnMB*) are also included to control for market expectations of future performance growth. Accruals (*Accr*) are included to control for the differential persistence of accruals and cash flow components of earnings (Sloan 1996). Abnormal earnings reserves (*AbnER*), due to conservative accounting treatments of inventory and R&D, will positively affect future *RNOA*. Results from Panel A of Table 8 show that the variable *Dual* is positively associated with one-year ahead *RNOA* (at a 5% significance level) and three-year ahead *RNOA* (at a 1% significance level). Using *Wedge* as a proxy for dual-class status generates consistent, but weaker results.[25]

Next, in the following multivariate test for one-year ahead and cumulative three-year ahead size-adjusted returns, I control for several risk proxies and pricing factors:

$$SAR_{i,t+k} = \phi_0 + \phi_1 DualProxy_{i,t} + \phi_2 Beta_{i,t} + \phi_3 LnMV_{i,t} + \phi_4 LnBM_{i,t} + \phi_5 Lev_{i,t} + \phi_6 EP_{i,t} + \phi_7 SAR_{i,t} + \phi_8 Accr_{i,t} + \phi_9 AbnER_{i,t} + \sum \delta_T Year + \sum \delta_I Industry + \varepsilon_{i,t} \quad (8)$$

where, for firm i in year t (unless otherwise indicated):

$SAR_{i,t+k}$ = One-year ahead and cumulative three-year ahead ($k = 1, 3$) size-adjusted stock returns.

LnBM = The natural logarithm of book value of equity divided by market value of equity at the fiscal year end.

EP = The earnings-to-price ratio.

All other variables are defined as before.

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[25] Because performance measures based on cash flows are less susceptible to accruals management than earnings-based performance measures such as *RNOA*, I also run regressions with future operating cash flow (*CFO*) as the dependent variable. Factors associated with future *CFO* include size, market-to-book ratio, abnormal earnings reserves, current-period *CFO*, and cash flow volatility. I find that dual-class structures are positively and significantly associated with future cash flow from operations.

Following the risk factors identified by Fama and French (1992), I control for firm size (*LnMV*), the log of book-to-market ratio (*LnBM*), systematic risk (*Beta*), and financial leverage (*Lev*). Further, I include the earnings-to-price ratio (*EP*) since Basu (1977) documents that earnings-to-price ratio is systematically associated with future returns. Current year's stock returns ($SAR_{i,t}$) is used to capture any price momentum impacts (Jegadeesh and Titman 1993). Moreover, I control for two accounting factors: total accruals and abnormal earnings reserves. Results are presented in Panel B of Table 8. Notably, conditional upon risk proxies and pricing factors, the coefficients on *Dual* (Model 1) and *Wedge* (Model 2) are insignificant when the independent variable is one-year ahead *SAR*. However, when the independent variable is cumulative three-year ahead *SAR*, the coefficients on *Dual* (Model 3) and *Wedge* (Model 4) are positive and significant. For example, in Model 3, the coefficient on *Dual* is 0.165, indicating that if all else remains the same, dual-class stocks will outperform single-class stocks by 16.5 percent over the subsequent three years since the construction of portfolio.

Consistent with the capital market pressure hypothesis, analyses in this section provide evidence that subsequent operating performance (in terms of *RNOA* and *CFO*) and stock returns (*SAR*) are significantly higher for dual-class firms than for matched single-class firms, after controlling for other determinants of operating and stock return performance.

5.4 Within-sample analyses of dual-class share structure on earnings quality

While my main analyses are based on a matched sample of dual-class and single-class firms, there is substantial variation of cash flow and voting rights among dual-class firms. For example, as shown in Panel A of Table 2, the standard deviation of insiders'

cash flow rights is 0.147, and the standard deviation of insiders' voting rights is 0.202. So I repeat the main analyses, but with several modifications. First, I constrain the sample to dual-class firms only. Second, to provide a more robust result of both incentive and entrenchment effects (Adams and Ferreira 2007), I include insiders' cash flow rights (*CFRights*) in the regressions. And finally, I examine whether the relation between financial reporting and the wedge of cash flow and voting rights is non-linear. So I add two more variables in the regressions: the square of the wedge ($Wedge^2$) and the square of cash flow rights ($CFRights^2$).

The results are reported in Table 9. Three general conclusions are derived from these results. First, significant and negative coefficients on the variable *Wedge* indicate that the higher the wedge, the less likely dual-class firms are to engage in earnings management. Second, consistent with the incentive alignment effect, dual-class firms are less likely to engage in earnings management with higher insiders' cash flow rights. Finally, there appears to be a U-shaped relation between absolute abnormal accruals and the wedge, since the coefficient on $Wedge^2$ in Panel B is positive and significant. The result suggests that dual-class firms with the wedge greater than 57.8% start to report higher abnormal accruals. A similar U-shaped relation also exists between absolute abnormal accruals and insiders' cash flow rights.

5.5 Financial reporting quality for firms initiating or abolishing dual-class structures

If dual-class structures lead to higher financial reporting quality, I would expect to observe a change in financial reporting behavior for firms that initiate or abolish dual-class structures. In the sample period, some single-class firms recapitalized into dual-class share structures ("recapitalization"), and other firms abolished dual-class structures

("unification"). For the capital market pressure hypothesis to hold, I expect that higher earnings quality is associated with dual-class status, but not single-class status, for affected sample firms.

Out of the initial sample I compile a list of 91 firms that unified their share structures, and 72 firms that underwent dual-class recapitalization for the period of 1990-2006. Using the same variables as in Equations (1), (2) and (4), I estimate the firm-specific time-series fixed-effects models (e.g., Himmelberg, Hubbard, and Palia 1999).[26] Unlike prior sections that examine variation in financial reporting quality across dual-class and single-class firms, the panel-structure regressions in this section are intended to compare financial reporting behavior of a firm under the dual-class regime with that of the *same* firm under the single-class regime.

Results are reported in Table 10. I focus on the coefficients on *Dual*, which is a dummy that equals one if a sample firm is under the dual-class regime, and zero otherwise. For the sample of dual-class recapitalization, the coefficients on *Dual* in the three tests (i.e., the propensity to meet or just beat forecasts, the propensity of accruals management, and the propensity to release earnings reserves) are negative and significant at conventional levels. Thus, after these sample firms recapitalize into dual-class structures, they report lower abnormal accruals, are less prone to reporting small positive earnings surprises, and are less likely to release reserves. However, for the sample of share unification, the coefficient on *Dual* is insignificant in two out of three tests (see the right columns of Panels A-C). The only negative and significant coefficient on *Dual* is for the test of propensity to release earnings reserves in Panel C. One reason for the weak

[26] A number of firms (78 firms among 163 firms) do not have compensation data at ExecuComp, so I hand-collected the data from proxy statements.

results in the "unification" sample could be the following. Although these firms abolish dual-class share structures, other forms of anti-takeover protection (e.g., poison pills, and staggered boards) remain in effect. These anti-takeover devices, accompanied by other factors (such as close monitoring by large blockholders), continue to drive managers to focus on firms' long-term value (see Hauser and Lauterbach 2004). Therefore I find little difference in financial reporting behaviors for the "unification" sample.

6. Further Robustness Checks

I test the sensitivity of the main results in the following un-tabulated specification checks.

Accounting for potential endogeneity of dual-class structures. In the main tests, I control for significant differences across single-class and dual-class firms. I then follow Hausman (1978) and test for the possible endogeneity of dual-class structures. This test does not reject the null of exogeneity of dual-class structures with a p-value of 0.16. Nonetheless, I conduct a two-stage analysis to account for potential endogeneity of dual-class share structures, with the first stage being the probit selection model of dual-class structures.[27] Following Gompers et al. (2005), the determinants of dual-class status include firm characteristics such as size (measured by *LnSales*), leverage, profitability (measured by *RNOA*), growth potential (proxied by growth in sales), and a firm's inclusion in a media industry (with 2-digit SIC codes of 48 and 78). The pseudo R^2 indicates that the selection model explains 48 percent of the cross-sectional variation in the choice to be a dual-class firm within the sample. I then use the estimates from the

[27] The dependent variable in the first-stage probit model is *Dual*. The two-stage analysis is based on the rationale that a firm's choice to embrace a dual-class or single-class share structure is continuous. In other words, firms always have the option to change share structure at any point in time.

probit model to compute the inverse Mills ratio for each sample firm (Heckman 1979). In the second stage, I estimate Equations (1), (2) and (4) with the inverse Mills ratio as a control. I obtain qualitatively similar results. That is to be expected, however, because I have included in the main tests various firm characteristics that correlate with both financial reporting quality and dual-class structures.

Separating family firms from other dual-class firms. I adopt a broad definition of family firms, in which founders or their descendents are key executives, directors, or blockholders (e.g., Anderson and Reeb 2003).[28] About 43.6 percent of dual-class firms in my sample are regarded as family firms. Further tests show that the main results hold across family and non-family dual-class firms.[29] Thus, results in the main sections are not driven by the superior performance of family firms (Anderson and Reeb 2003; Ali, Chen, and Radhakrishnan 2007; Wang 2006).

Results from analyses of sub-periods. The passage of the Sarbanes-Oxley Act ushered in more stringent corporate governance and financial reporting requirements, so I test whether the primary findings are robust under different sub-periods. I split the sample period into pre- and post-Sarbanes Oxley Act (SOX) periods. The pre-SOX period covers 1994 to 2002, and the post-SOX period spans 2003-2005. In association with abnormal accruals for the post-SOX sample period, the coefficient on *Dual* remains negative, but only significant at the 10% level. This could be due to a sharp decline in the overall use of accrual manipulations following the passage of the SOX (Cohen, Dey,

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[28] Following Anderson and Reeb (2003), I manually collected data on family ownership, family board representation, and CEO attributes from proxy statements. I also manually examined corporate histories for each firm in the sample. Histories were gleaned from Gale Business Resources, Hoovers, and from company press releases and websites.

[29] The main results also hold after I classify family firms into three groups with different CEO attributes: founder CEO, descendant CEO, and professional CEO.

and Lys 2008). Resultantly it is becoming increasingly difficult to detect the differential level of abnormal accruals across dual-class and single-class firms. Other results remain unaltered.

Results from yearly regressions. In Table 4, cross-sectional correlation in the residuals could induce biased inferences. To mitigate this concern, I also perform Fama-MacBeth (1973) procedures based on thirteen annual regressions. The Fama-MacBeth t-statistics are slightly smaller than those reported in Table 4 but still significant at $p < 0.10$ or lower. The inference that dual-class structures are associated with less accruals management is consistent with the results reported in Section 4.3.

Controlling for a richer set of governance attributes. I try to control for key governance data from the IRRC database. Wherever the governance data is available, I include a management entrenchment index (E-score) of Bebchuk, Cohen, and Ferrell (2004) in the regressions. I also add in the regressions G-score, another index of shareholder rights compiled by Gompers, Ishii, and Metrick (2003). Controlling for additional governance attributes does not materially change my main conclusions.

7. Concluding Remarks

This study develops and tests predictions about differences in earnings management behaviors across a sample of publicly-traded U.S. dual-class and single-class firms. The differing intensity of capital market pressures observed between dual-class and single-class firms has important implications for firms' financial reporting strategy, profitability, and growth. Capital market pressures can drive companies to sacrifice the long-term interests of their shareholders in order to meet short-term market expectations. I provide evidence that dual-class firms, as opposed to single-class firms,

are less likely to manage reported earnings to meet or marginally beat analysts' earnings forecasts. Furthermore, managers at dual-class firms are less likely to engage in accrual management and/or to release earnings reserves relative to managers at single-class firms, who otherwise experience greater capital market pressures.

On a cautionary note, the outcome derived from the U.S. setting may not apply to foreign markets. My study takes advantage of two important U.S. institutional features. One is obvious: dual-class structures reduce otherwise intense capital market pressures laid on managers of U.S. public firms. Less apparent is the other institutional factor that drives the results, i.e., strong shareholder protection afforded by the more rigorous financial reporting requirements in the U.S. To capital market regulators, this study implies that restricting dual-class share structures seems to be unnecessary,[30] as dual-class structures are not necessarily value-reducing to inferior-class shareholders in the U.S. Current U.S. laws and regulations have shown themselves to be sufficient in curbing insiders' expropriation activities.

[30] The New York Stock Exchange (NYSE) banned dual-class shares in the 1920s. The discrimination of dual-class shares lasted until the late 1980s.

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Table 1
Distribution of the Sample

Panel A: Sample distribution by year

Year	Number of initial dual-class firms	Number of dual-class firms used in the tests
1994	110	78
1995	400	141
1996	444	165
1997	485	192
1998	504	231
1999	489	240
2000	482	210
2001	434	228
2002	362	198
2003	412	155
2004	422	194
2005	458	252
2006	426	218
Total	5,002	2,502

Panel B: Sample distribution by industry groups

Industry description	Percentage of firms in Compustat	Percentage of dual-class firms
Agriculture and food	2.17%	2.48%
Mining and construction	2.77%	1.22%
Textiles and publishing	3.92%	5.62%
Chemicals	1.81%	1.79%
Pharmaceuticals	3.88%	3.05%
Extractive and refining	4.23%	1.83%
Durable manufacturers	17.26%	20.53%
Transportation	6.25%	12.30%
Utilities	2.99%	1.63%
Retail	9.23%	6.47%
Banking or financial services	21.73%	17.48%
Services	9.46%	11.18%
Computers	12.95%	14.03%
Others	1.35%	0.40%
Total	100.00%	100.00%

Table 1 gives the sample distribution by year (Panel A) and by industry groups (Panel B) over the period 1994-2006. Industry membership is determined by SIC codes as follows: Agriculture and food (0100-0999 and 2000-2199), Mining and construction (1000-1999, excluding 1300-1399), Textiles and printing/publishing (2200-2799), Chemicals (2800-2824, 2840-2899), Pharmaceuticals (2830-2836), Extractive and refining (2900-2999, 1300-1399), Durable manufacturers (3000-3999, excluding 3570-3579 and 3670-3679), Transportation (4000-4899), Utilities (4900-4999), Retail (5000-5999), Banking or financial services (6000-6999), Services (7000-8999 excluding 7370-7379), Computers (7370-7379, 3570-3579, 3670-3679), and Other (>9000). The industry classification scheme is based on Barth, Beaver, and Landsman (1998).

Table 2
Firm Characteristics of Dual-Class Firms and Single-Class Firms

Panel A: Cash flow rights and voting rights for dual-class firms

	Mean	25%	Median	75%	Std. Dev.
CFRights	0.404	0.123	0.425	0.512	0.147
VoteRights	0.617	0.265	0.630	0.748	0.202
Wedge	0.213	0.143	0.205	0.236	0.121

Panel B: Fundamental firm characteristics

	Dual-class firms			Single-class firms			Test for difference	
Variables	Mean	Median	Std. Dev.	Mean	Median	Std. Dev.	t-stat	z-stat
Assets ($M)	5365.704	1054.202	8644.660	5371.409	1057.702	8633.304	-0.42	-0.23
Sales ($M)	5835.345	1573.111	6744.043	5973.420	1648.928	7464.981	-1.04	-1.39
NOA ($M)	2414.915	1985.843	2611.432	1993.678	1811.488	2651.245	1.45	0.72
Div	0.019	0.013	0.016	0.012	0.008	0.015	1.84*	1.92*
Lev	0.256	0.215	0.246	0.244	0.211	0.191	2.03**	1.27
ExtFin	0.412	0.413	1.149	0.408	0.394	0.978	0.47	0.61
Herf	0.032	0.026	0.012	0.037	0.034	0.012	-0.85	-0.73
Segment	0.386	0.337	0.311	0.482	0.377	0.332	-8.56***	-8.62***
OpCyc	136.231	134.562	124.035	134.701	124.954	121.485	0.21	1.36

Panel C: Operating performance

	Dual-class firms			Single-class firms			Test for difference	
Variables	Mean	Median	Std. Dev.	Mean	Median	Std. Dev.	t-stat	z-stat
RNOA	0.120	0.099	0.133	0.096	0.084	0.123	2.62***	2.51***
GSales	0.123	0.115	0.252	0.109	0.106	0.250	2.22**	1.86*
CFO	0.099	0.093	0.112	0.099	0.096	0.113	0.03	-0.97
StdCFO	0.085	0.051	0.080	0.067	0.041	0.091	2.74***	2.01**
FCF	0.042	0.045	0.115	0.026	0.038	0.116	4.32***	4.01***
Accr	-0.031	-0.022	0.089	0.051	0.050	0.088	-0.89	-0.72
AbnAccr	0.021	0.015	0.124	0.028	0.019	0.144	-2.62***	-1.97**
Loss	0.106	0.000	0.242	0.109	0.000	0.265	-0.27	-0.15
RDAdv	0.057	0.000	0.261	0.043	0.000	0.303	3.78***	2.86***
ER	0.293	0.202	0.309	0.249	0.178	0.333	5.49***	6.48***
AbnER	0.014	0.000	0.033	-0.006	0.003	0.041	1.99**	-0.33

Panel D: Stock performance

	Dual-class firms			Single-class firms			Test for difference	
Variables	Mean	Median	Std. Dev.	Mean	Median	Std. Dev.	t-stat	z-stat
MV ($M)	4820.435	1120.275	11081.305	5107.698	1257.601	13063.900	-2.55***	-3.56***
MB	3.541	2.356	4.722	3.379	2.284	4.031	1.23	1.12
EP	0.034	0.044	0.184	0.035	0.046	0.166	-0.21	-0.11
TobinQ	1.687	1.304	1.194	1.696	1.251	1.431	-1.12	0.99
Ret	0.152	0.121	0.835	0.150	0.096	1.032	1.05	1.72*
SAR	0.014	0.013	0.451	0.013	0.014	0.436	0.22	-0.14
Beta	1.114	1.050	0.564	1.087	1.032	0.512	1.44	0.89
IdRisk	0.023	0.018	0.020	0.019	0.016	0.021	0.86	0.24
Liq	2.132	2.070	0.838	2.540	2.321	0.889	-2.83***	-2.75***

Panel E: Governance attributes

	Dual-class firms			Single-class firms			Test for difference	
Variables	Mean	Median	Std. Dev.	Mean	Median	Std. Dev.	t-stat	z-stat
Analysts	4.289	4.284	0.842	6.042	5.204	0.892	-9.24***	-9.51***
InstOwn	0.246	0.241	0.731	0.278	0.279	0.674	-1.85*	-1.82*
Auditor	0.938	1.000	0.242	0.978	1.000	0.123	-6.82***	-6.94***
Option	0.246	0.000	0.412	0.414	0.221	0.476	-9.67***	-14.22***
Bonus	0.081	0.000	0.162	0.141	0.050	0.184	-8.69****	-11.45***
MgmtOwn	0.058	0.049	0.052	0.042	0.030	0.061	2.12**	2.14**

Table 2 presents descriptive statistics for 2,502 observations of dual-class firms and the same number of matched single-class firms over the period 1994-2006. The paired t-statistics report differences of means, and the z-statistics of the Wilcoxon signed rank tests report differences of medians. ***, **, * indicate, respectively, the significance levels at 1%, 5%, and 10%.

The variables are defined as follows:

Cash Flow Rights and Voting Rights:

CFRights = The total percentage of cash flow ownership by officers and directors.
VoteRights = The total percentage of votes owned by officers and directors across classes.
Wedge = The difference between voting and cash flow ownership by officers and directors.

Fundamental Firm Characteristics:

Assets = Total assets (Compustat #6).
Sales = Sales (#12). The natural logarithm of sales is used in the tests.
NOA = Net operating assets are measured as market value shareholders' equity (#60) less cash and marketable securities (#1) plus total debt (#9 + #34). I use net operating assets divided by lagged sales in the tests.
Div = Dividend yield, measured as dividend (#21) / market value of equity at the fiscal year end.
Lev = Financial leverage, measured as total debt (#9 + #34) / total assets (#6).

ExtFin = External financing dependence, = -1.002*cash flow (#14 + #18) / lagged total assets (#6) – 39.368*cash dividends (#21 + #19) / lagged total assets – 1.315*cash balance (#1) / lagged total assets + 3.139*leverage ((#9 + #34) / #6) + 0.283*Tobin's Q.

Herf = Industry concentration (the Herfindahl index), measured as the sum of squared market shares in the industry. Market share is measured as sales of a firm as a percentage of total sales in the same industry.

Segment = Operating segment diversification, calculated as the within-firm Herfindahl index, i.e., the sum of squared segment-level sales within a firm.

OpCyc = The length of operating cycle, where operating cycle = 360 x [(#2 / #12) + (#3 / #41)]. The natural logarithm of operating cycle is used in the tests.

Operating Performance:

RNOA = Return on net operating assets, measured as net income before extraordinary items (#18) divided by net operating assets at year *t*-1 (#6).

GSales = Annual growth in sales (#12).

CFO = Cash flow from operations ((#308 - #124) scaled by lagged total assets (#6).

StdCFO = Standard deviation of cash flow from operations (#308 - #124), scaled by lagged total assets (#6), over the rolling prior 5 years.

FCF = Free cash flow as cash flow from operations (#308) less capital expenditure (#128), and then scaled by lagged total assets (#6).

Accr = Total accruals, measured as earnings before extraordinary items (#123) minus cash flow from operating activities (#308), and scaled by lagged total assets (#6).

AbnAccr = Abnormal accruals. The estimation method is detailed below.

Loss = A dummy variable that equals 1 for firms reporting losses before extraordinary items (#18), 0 otherwise.

RDAdv = R&D and advertising expense scaled by sales ((#45 + #46) / #12).

ER = Earnings reserve measure developed by Penman and Zhang (2002), measured as (inventory reserve + R&D reserve + advertising reserve) / net operating assets. Inventory reserve equals the LIFO reserve reported in footnotes; R&D reserve is calculated as the estimated amortized R&D assets as if R&D had not been expensed; and advertising reserve is calculated as estimated brand assets created by advertising expenditures.

AbnER = Abnormal earnings reserves, as estimated from Equation (3). See Section 3.4 for details.

Stock Performance:

MV = Market value of equity at the fiscal year end. The number of shares outstanding for dual-class firms includes traded and non-traded shares. The number of non-traded shares for dual-class firms is obtained from proxy statements. I assume that non-traded shares are priced the same as traded shares.

MB = Market value of equity at the fiscal year end divided by book value of equity (#60).

EP = The earnings-to-price ratio, measured as net income before extraordinary items (#18) / market value of equity at the fiscal year end.

TobinQ = Tobin's Q, defined as (market value of equity + book value of preferred stock + long-term debt + short-term debts) / total assets, = (#199 x #25 + #130 + #9 + #34) / #6.

Ret = Yearly buy-and-hold stock returns, starting from four months after the fiscal year-end. For the dual class firms, returns are based on the inferior-class shares.

SAR = Size-adjusted stock returns, measured as the difference between the buy-and-hold return for a firm starting from four months after the fiscal year end and the corresponding buy-and-hold return of the size decile portfolio to which the firm belongs.

Beta = Market beta, measured as the slope in the regression of a rolling window of five years of monthly returns against the monthly CRSP value-weighted market index.

IRisk = Idiosyncratic stock return volatility, measured as the average monthly variance of market-adjusted returns for a fiscal year. Market adjusted returns are the excess of daily individual stock return over the daily return on the value-weighted market portfolio.

Liq = Liquidity, measured as the natural logarithm of average monthly trading volume divided by shares outstanding.

Governance Attributes:

Analysts = The number of analysts issuing annual earnings forecasts. The natural logarithm of *Analysts* is used in the tests.

InstOwn = The percentage of total outstanding shares held by institutional investors. Institutional percentage shareholding is averaged over four quarters each year.

Auditor = A dummy variable that is set to 1 if the firm is audited by a Big-5 audit firm or successor (#149), 0 otherwise.

Option = CEO's stock option incentives, measured as the ratio of vested options held by the CEO scaled by his firm-specific wealth. CEO's firm-specific wealth is the sum of salary, bonus and annual compensation, stock ownership, vested and unvested options.

Bonus = Bonuses paid to the CEO scaled by CEO's firm-specific wealth.

MgmtOwn = The percentage of stock holdings (including restricted stock) held by top managers.

To estimate abnormal accruals, I first estimate ordinary least square (OLS) regressions for all non-sample firms for every industry classified by its 2-digit SIC code listed on the Compustat universe with at least 20 observations in year *t*. Cross-listed international firms, closed-end funds, and REITs are removed from the estimation sample. Specifically, the following regression is estimated for each industry-year:

$$Accr_{j,t} = \gamma_0 \frac{1}{Assets_{j,t-1}} + \gamma_1 \frac{\Delta Sales_{j,t} - \Delta AR_{j,t}}{Assets_{j,t-1}} + \gamma_2 \frac{PPE_{j,t}}{Assets_{j,t-1}} + \varepsilon_{j,t}$$

where, for firm *j* in year *t* (unless otherwise indicated):

Assets = Total assets (#6) for year *t*-1.

$\Delta Sales$ = The change in revenues (#12) between year *t*-1 and *t*.

ΔAR = The change in accounts receivable (#2) between year *t*-1 and *t*.

PPE = Gross property, plant, and equipment (#7).

The coefficient estimates from the above equation are used to estimate the firm-specific expected accruals for the sample firms. Abnormal accruals are the difference between total accruals and expected accruals. I then match each sample firm with the control firm on industry, year, and return on assets, to arrive at the performance-matched abnormal accruals.

Table 3
Propensity to Meet or Just Beat Analysts' Forecasts
across Dual-Class and Single-Class Firms

Panel A: Firm-year observations with earnings surprises in the intervals just below/above zero

Share structure type	Earnings surprise interval [-$0.02, -$0.01]			Earnings surprise interval [$0.00, $0.01]		
	Actual number of observations	Expected number of observations	Standardized difference	Actual number of observations	Expected number of observations	Standardized difference
Dual-class firms	392	436	-1.51	490	448	1.45
Single-class firms	295	422	-4.86***	558	434	3.41***

Panel B: The logit model of propensity to meet or just beat analysts' forecasts

$$\begin{aligned}\text{Prob}(MBE_{i,t} \in [\$0, \$0.01]) = F(&\alpha_0 + \alpha_1 DualProxy_{i,t} + \alpha_2 LnMV_{i,t} + \alpha_3 LnMB_{i,t} + \alpha_4 Lev_{i,t} \\ &+\alpha_5 Option_{i,t} + \alpha_6 Bonus_{i,t} + \alpha_7 MgmtOwn_{i,t-1} + \alpha_8 NOA_{i,t-1} + \alpha_9 RNOA_{i,t-1} + \alpha_{10} Loss_{i,t-1} \\ &+\alpha_{11} RDAdv_{i,t-1} + \alpha_{12} StdCFO_{i,t} + \alpha_{13} LnAnalysts_{i,t} + \alpha_{14} InstOwn_{i,t} + \alpha_{15} ExpMgmt_{i,t} \\ &+\sum \delta_T Year + \sum \delta_I Industry + \varepsilon_{i,t})\end{aligned} \quad (1)$$

Variables	Pred. Sign	Model 1		Model 2	
		Coef. Est.	p-value	Coef. Est.	p-value
Intercept	?	-2.683	0.001	-2.486	0.001
Dual	-	-0.038	0.042		
Wedge	-			-0.055	0.008
LnMV	+	0.142	0.014	0.145	0.016
LnMB	+	0.451	0.001	0.445	0.001
Lev	+	-0.632	0.005	-0.642	0.005
Option	+	0.193	0.004	0.188	0.005
Bonus	+	0.362	0.005	0.352	0.005
MgmtOwn	-	0.456	0.182	0.481	0.128
NOA	-	-0.012	0.062	-0.010	0.054
RNOA	+	1.633	0.001	1.644	0.001
Loss	-	-0.220	0.037	-0.252	0.043
RDAdv	+	0.011	0.005	0.005	0.005
StdCFO	-	-0.020	0.009	-0.015	0.017
LnAnalysts	+	0.263	0.001	0.260	0.001
InstOwn	-	0.025	0.152	0.021	0.229
ExpMgmt	+	0.421	0.001	0.418	0.001
Year and Industry		Yes		Yes	
N		5,004		5,004	
Pseudo R^2		0.232		0.238	
Likelihood ratio		261.433		265.428	

Table 3 presents results of the propensity to meet or just beat analysts' forecasts over the period 1994-2006. ***, **, * indicate, respectively, the significance levels at 1%, 5%, and 10%.

In Panel A, the standardized difference is the difference between the observed and expected number of observations in an interval, divided by the estimated standard deviation of the difference.

In Panel B, *MBE* is a dummy variable that equals 1 if the difference between actual earnings per share and consensus analysts' forecasted earnings per share is within [$0, $0.01], and 0 otherwise. The variable *DualProxy* consists of *Dual*, and *Wedge*. All other variables are defined in Table 2. For parsimony, the coefficients on year dummies and industry dummies are not reported.

Table 4
Propensity for Accruals Management across Dual-Class and Single-Class Firms

Panel A: Abnormal accruals for firms that meet or just beat analysts' forecasts, compared to all other firms

Share structure type	Firms with $MBE \in [\$0, \$0.01]$			All other firms in the sample			Test for difference	
	N	Mean	Median	N	Mean	Median	t-stat	z-stat
Dual-class sample	490	0.024	0.019	2,012	0.022	0.016	0.97	0.53
Single-class sample	558	0.036	0.027	1,944	0.027	0.021	2.50***	1.98**

Panel B: OLS regressions of absolute abnormal accruals on dual-class proxies and other controls

$$\begin{aligned}|AbnAccr_{i,t}| &= \beta_0 + \beta_1 DualProxy_{i,t} + \beta_2 LnMV_{i,t} + \beta_3 LnMB_{i,t} + \beta_4 Lev_{i,t} + \beta_5 StdCFO_{i,t} \\ &+ \beta_6 Option_{i,t} + \beta_7 Bonus_{i,t} + \beta_8 MgmtOwn_{i,t-1} + \beta_9 ExtFin_{i,t} + \beta_{10} MBE_{i,t} + \beta_{11} NOA_{i,t-1} \\ &+ \beta_{12} Loss_{i,t-1} + \beta_{13} RDExp_{i,t-1} + \beta_{14} Herf_{i,t} + \beta_{15} Segment_{i,t} + \beta_{16} LnOpCyc_{i,t} \\ &+ \beta_{17} InstOwn_{i,t} + \beta_{18} Auditor_{i,t} + \sum \delta_T Year + \sum \delta_I Industry + \varepsilon_{i,t}\end{aligned} \quad (2)$$

Variables	Pred. Sign	Model 1		Model 2	
		Coef. Est.	t-stat	Coef. Est.	t-stat
Intercept	?	1.022	3.12***	1.090	3.01***
Dual	-	-0.224	-2.22**		
Wedge	-			-0.077	-1.72*
LnMV	-	-0.122	-2.82***	-0.084	-2.30**
LnMB	+	0.157	3.16***	0.152	3.15***
Lev	+	-0.382	-2.61***	-0.383	-2.67***
StdCFO	+	0.921	2.33**	0.926	2.22**
Option	+	0.016	2.15**	0.017	2.10**
Bonus	+	0.023	2.04**	0.024	2.11**
MgmtOwn	-	-0.231	-2.16**	-0.239	-2.13**
ExtFin	+	0.019	1.10	0.018	1.01
MBE	+	0.025	2.01**	0.022	1.99**
NOA	-	0.004	0.47	0.007	0.45
Loss	-	0.008	0.36	0.007	0.33
RDAdv	+	0.006	2.10**	0.003	2.09**
Herf	+	-0.482	-1.53	-0.475	-1.55
Segment	-	-0.120	-2.11**	-0.115	-2.07**
LnOpCyc	+	0.051	1.73*	0.051	1.81*
InstOwn	-	-0.023	-1.81*	-0.021	-1.84*
Auditor	-	-0.032	-2.05**	-0.031	-2.09**
Year and Industry		Yes		Yes	
N		5,004		5,004	
Adj. R^2		0.123		0.127	

Table 4 presents the results of the association between dual-class structures and accruals management over the period 1994-2006. ***, **, * indicate, respectively, the significance levels at 1%, 5%, and 10%.

In Panel A, the t-statistics report differences of means, and the z-statistics of the Wilcoxon rank score tests report differences of medians.

In Panel B, the variable *DualProxy* consists of *Dual* and *Wedge*. All other variables are defined in Table 2. For parsimony, the coefficients on year dummies and industry dummies are not reported. The t-statistics are based on Huber-White adjusted standard errors.

Table 5
Propensity for Releasing Earnings Reserves across Dual-Class and Single-Class Firms

Panel A: Estimating expected earnings reserves

$$ER_{j,t} = \gamma_0 + \gamma_1 ER_{j,t-1} + \gamma_2 \Delta ER_{j,t-1} + \gamma_3 GSales_{j,t} + \gamma_4 RNOA_{j,t} + \gamma_5 FCF_{j,t-1} + \gamma_6 TobinQ_{j,t} + \gamma_7 Age_{j,t} + \varepsilon_{j,t} \quad (3)$$

		Model 1		Model 2	
Variables	Pred. Sign	Coef. Est.	t-stat	Coef. Est.	t-stat
Intercept	?	0.006	2.12**	0.010	1.43
$ER_{j,t}$	+	0.857	18.32***	0.853	20.22***
ΔER	-	-0.122	-1.93*	-0.081	-1.76*
GSales	+			0.293	1.83*
RNOA	+			-0.091	-1.93*
FCF	+			0.032	0.95
TobinQ	+			0.558	2.78***
Age	-			-0.485	-2.01**
Average N		1,824		1,824	
Adj. R^2		0.772		0.814	

Panel B: Abnormal earnings reserves for firms that meet or just beat analysts' forecasts, compared to all other firms

Share structure type	Firms with $MBE \in [\$0, \$0.01]$			All other firms in the sample			Test for difference	
	N	Mean	Median	N	Mean	Median	t-stat	z-stat
Dual-class sample	490	0.011	-0.001	2,012	0.012	0.003	-0.47	-1.27
Single-class sample	558	-0.009	-0.006	1,944	0.011	0.003	-2.23**	-1.97**

Panel C: The logit model of probability of releasing earnings reserves

$$\text{Prob}(AbnER_{i,t} < 0) = F(\kappa_0 + \kappa_1 DualProxy_{i,t} + \kappa_2 Option_{i,t} + \kappa_3 Bonus_{i,t} + \kappa_4 MgmtOwn_{i,t-1} + \kappa_5 ExtFin_{i,t} + \kappa_6 MBE_{i,t} + \kappa_7 Lev_{i,t} + \kappa_8 StdCFO_{i,t} + \kappa_9 NOA_{i,t-1} + \kappa_{10} InstOwn_{i,t} + \sum \delta_T Year + \sum \delta_I Industry + \varepsilon_{i,t}) \quad (4)$$

		Model 1		Model 2	
Variables	Pred. Sign	Coef. Est.	p-value	Coef. Est.	p-value
Intercept	?	-1.422	0.001	-1.531	0.001
Dual	-	-0.075	0.001		
Wedge	-			-0.082	0.005
Option	+	0.075	0.032	0.072	0.035
Bonus	+	0.512	0.001	0.493	0.004
MgmtOwn	-	-1.102	0.004	-1.011	0.012
ExtFin	+	0.332	0.001	0.332	0.001
MBE	+	0.084	0.038	0.088	0.027
Lev	+	-0.988	0.001	-0.974	0.001
StdCFO	+	0.242	0.003	0.239	0.008
NOA	-	-0.018	0.117	-0.014	0.210
InstOwn	-	0.022	0.214	0.043	0.202
Year and Industry		Yes		Yes	
N		5,004		5,004	
Pseudo R^2		0.144		0.151	
Likelihood ratio		334.893		336.743	

Table 5 presents results of the propensity for releasing abnormal earnings reserves over the period 1994-2006. ***, **, * indicate, respectively, the significance levels at 1%, 5%, and 10%.

Panel A presents regression results for all firms with available data on the Compustat universe for every industry with at least 10 observations in year *t*. ΔER is the annual change in earnings reserves. *Age* is the log of the number of years the firm has been listed on CRSP as of the start of the year.

In Panel B, the t-statistics report differences of means, and the z-statistics of the Wilcoxon rank score tests report differences of medians.

In Panel C, the variable *DualProxy* consists of *Dual* and *Wedge*. All other variables are defined in Table 2. For parsimony, the coefficients on year dummies and industry dummies are not reported.

Table 6
The Interaction of Shareholder Base and Dual-Class Structures on Earnings Quality

Panel A: Comparing types of institutional investors across dual-class and single-class firms

	Dual-class firms			Single-class firms			Test for difference	
Variables	Mean	Median	Std. Dev.	Mean	Median	Std. Dev.	t-stat	z-stat
InstTurn	22.40%	21.80%	0.784	26.60%	26.20%	0.645	-2.88***	-3.12***
LTInstOwn	18.40%	17.20%	0.477	18.20%	17.80%	0.643	0.11	-0.69
STInstOwn	6.20%	7.00%	0.328	9.60%	10.10%	0.421	-2.66***	-3.25***

Panel B: The Relationship between institutional investor turnover and dual-class firms

$$InstTurn_{i,t} = \lambda_0 + \lambda_1 Dual_{i,t} + \lambda_2 LnMV_{i,t-1} + \lambda_3 Liq_{i,t-1} + \lambda_4 LnMB_{i,t-1} + \lambda_5 RNOA_{i,t-1}$$
$$+\lambda_6 Lev_{i,t-1} + \lambda_7 Beta_{i,t} + \lambda_8 IRisk_{i,t} + \lambda_9 Ret_{i,t-1} + \lambda_{10} Div_{i,t-1} \quad (5)$$
$$+\sum \delta_T Year + \sum \delta_I Industry + \varepsilon_{i,t}$$

Variables	Pred. Sign	Coef. Est.	t-stat
Intercept	?	0.284	30.34***
Dual	+	-0.004	-2.01**
LnMV	+	0.007	2.96***
Liq	+	0.008	12.53***
LnMB	+	0.009	3.50***
RNOA	+	0.004	1.60
LEV	-	0.000	0.58
Beta	+	0.005	6.19***
IRisk	+	0.002	2.01**
Ret	+	0.006	2.01**
Div	-	-0.204	-7.05***
Year and Industry		Yes	
N		4,568	
Adj. R^2		0.120	

Table 6 presents the association of various types of institutional ownership and dual-class firms. Types of institutional investors include the following: the weighted average of the average total portfolio churn rates of a firm's institutional investors over four quarters (*InstTurn*), percentage holdings by long-term institutional investors (*LTInstOwn*), and percentage holdings by short-term institutional investors (*STInstOwn*). Long-term institutional investors include quasi-indexers and dedicated institutional investors, and short-term institutional investors consist of transient institution investors, as classified by Bushee (1998). Institutional investor turnover (*InstTurn*) is calculated as follows. I first calculate for each institutional investor a measure of how frequently he rotates his positions on all of the stocks in his portfolio for a given quarter. Define the turnover rate (*Turnover*) of institutional investor *j* at quarter *q* as,

$$Turnover_{j,q} = \frac{\sum_{i=1}^{Q} N_{i,j,q}P_{i,q} - N_{i,j,q-1}P_{i,q-1} - N_{i,j,q-1}\Delta P_{i,q-1}}{\sum_{i=1}^{Q}(N_{i,j,q}P_{i,q} + N_{i,j,q-1}P_{i,q-1})/2},$$

where $P_{i,q}$ and $N_{i,j,q}$ represent price and number of shares, respectively, of firm i held by institutional investor j at quarter q; and Q is the set of firms held by investor j. Then, I calculate institutional investor turnover for firm i as the weighted average of the turnover rates of all of its institutional investors over the most recent four quarters:

$$InstTurn = \sum_{j \in S} w_{i,j,q}\left(\frac{1}{4}\sum_{r=1}^{4} Turnover_{j,q-r-1}\right),$$

where S is the set of institutional investors in firm i, and $w_{i,j,q}$ as the weight of investor j in the total percentage held by institutional investors at quarter q in firm i. Year-specific *InstTurn* is reported as the last quarter of institutional investor turnover for every firm.

In Panel A, the t-statistics report differences of means, and the z-statistics of the Wilcoxon rank score tests report differences of medians.

In Panel B, all control variables are defined in Table 2. For parsimony, the coefficients on year dummies and industry dummies are not reported. The t-statistics are based on Huber-White adjusted standard errors. ***, **, * indicate, respectively, the significance levels at 1%, 5%, and 10%.

Table 7
Abnormal Returns around Earnings Announcements

$$CAR_{i,t} = \theta_0 + \theta_1 Dual_{i,t} + \theta_2 Good_{i,t} + \theta_3 Bad_{i,t} + \theta_4 Dual_{i,t} * Good_{i,t} + \theta_5 Dual_{i,t} * Bad_{i,t}$$
$$+\theta_6 UE_{i,t} + \theta_7 Good_{i,t} * UE_{i,t} + \theta_8 Bad_{i,t} * UE_{i,t} + \theta_9 Dual_{i,t} * Good_{i,t} * UE_{i,t}$$
$$+\theta_{10} Dual_{i,t} * Bad_{i,t} * UE_{i,t} + \theta_{11} LnMV_{i,t} + \theta_{12} LnMB_{i,t} + \theta_{13} Lev_{i,t} + \theta_{14} InstOwn_{i,t} \quad (6)$$
$$+\theta_{15} LnMV_{i,t} * UE_{i,t} + \theta_{16} LnMB_{i,t} * UE_{i,t} + \theta_{17} Lev_{i,t} * UE_{i,t} + \theta_{18} Beta_{i,t} * UE_{i,t}$$
$$+\sum \delta_T Year + \sum \delta_I Industry + \varepsilon_{i,t}$$

		Model 1		Model 2	
Variables	Pred. Sign	Coef. Est.	t-stat	Coef. Est.	t-stat
Intercept	?	0.020	2.24**	0.015	1.82*
Dual	?	-0.005	-0.80	-0.004	-0.63
Good	+	0.034	4.37***	0.030	4.40***
Bad	-	-0.040	-5.49***	-0.036	-5.31**
*Dual*Good*	+	0.007	0.72	0.006	0.65
*Dual*Bad*	+	0.015	2.36**	0.014	2.32**
UE	+	0.086	6.22***	0.075	6.11***
*Good*UE*	+	0.004	0.52	0.003	0.46
*Bad*UE*	+	0.011	1.77*	0.010	1.76*
*Dual*Good*UE*	+	0.009	1.49	0.013	1.54
*Dual*Bad*UE*	-	-0.022	-2.36**	-0.020	-2.16**
LnMV	?			-0.003	-1.79*
LnMB	?			-0.014	-2.16**
Lev	?			0.004	1.90*
InstOwn	?			0.012	2.71***
*LnMV*UE*	?			0.022	2.84***
*LnMB*UE*	?			0.062	3.32***
*Lev*UE*	?			-0.007	-0.87
*Beta*UE*	?			-0.014	-1.81*
Year and Industry		Yes		Yes	
N		4,992		4,992	
Adj. R^2		0.026		0.038	

Table 7 presents multivariate regression results of the three-day market-adjusted abnormal returns for dual-class and single-class firms around earnings announcements. The dependent variable is three-day market-adjusted stock returns (*CAR*). *UE* is the difference between actual earnings per share and consensus forecasts, then scaled by stock price at the fiscal year end. *Good* (*Bad*) is a dummy variable that equals to 1 if *UE* is positive (negative), and 0 otherwise. All other variables are defined in Table 2. For parsimony, the coefficients on year dummies and industry dummies are not reported. The t-statistics are based on Huber-White adjusted standard errors. ***, **, * indicate, respectively, the significance levels at 1%, 5%, and 10%.

Table 8
Future Performance of Dual-Class Firms vs. Single-Class Firms

Panel A: Future operating performance measured by *RNOA*

$$RNOA_{i,t+k} = \eta_0 + \eta_1 DualProxy_{i,t} + \eta_2 RNOA_{i,t} + \eta_3 LnMV_{i,t} + \eta_4 LnMB_{i,t} + \eta_5 Accr_{i,t} + \eta_6 AbnER_{i,t} + \sum \delta_T Year + \sum \delta_I Industry + \varepsilon_{i,t} \quad (7)$$

		One-year ahead $RNOA_{i,t+1}$ as the dependent variable				Three-year ahead $RNOA_{i,t+3}$ as the dependent variable			
		Model 1		Model 2		Model 3		Model 4	
Variables	Pred. Sign	Coef. Est.	t-stat	Coef. Est.	t-stat	Coef. Est.	t-stat	Coef. Est.	t-stat
Intercept	?	-0.022	-2.33**	-0.020	-2.50***	-0.020	-2.57***	-0.022	-2.72***
Dual	+	0.022	2.12**			0.027	2.92***		
Wedge	+			0.015	1.79*			0.012	1.42
RNOA	+	0.658	21.72***	0.658	22.01***	0.514	20.83***	0.512	20.28***
LnMV	+	0.004	2.34**	0.003	2.40**	0.005	3.83***	0.004	3.23***
LnMB	+	0.012	2.86***	0.012	2.84***	0.009	2.52***	0.008	2.61***
Accr	-	-0.152	-5.83***	-0.134	-5.99***	-0.132	-4.12***	-0.135	-4.32***
AbnER	+	0.185	3.24***	0.181	3.32***	0.121	2.48***	0.124	2.51***
Year		Yes		Yes		Yes		Yes	
Industry		Yes		Yes		Yes		Yes	
N		4,064		4,064		3,682		3,682	
Adj. R^2		0.468		0.484		0.321		0.315	

Panel B: Future stock performance measured by *SAR*

$$SAR_{i,t+k} = \phi_0 + \phi_1 DualProxy_{i,t} + \phi_2 Beta_{i,t} + \phi_3 LnMV_{i,t} + \phi_4 LnBM_{i,t} + \phi_5 Lev_{i,t} + \phi_6 EP_{i,t} + \phi_7 SAR_{i,t} + \phi_8 Accr_{i,t} + \phi_9 AbnER_{i,t} + \sum \delta_T Year + \sum \delta_I Industry + \varepsilon_{i,t} \quad (8)$$

		One-year ahead $SAR_{i,t+1}$ as the dependent variable				Cumulative three-year ahead $SAR_{i,t+3}$ as the dependent variable			
		Model 1		Model 2		Model 3		Model 4	
Variables	Pred. Sign	Coef. Est.	t-stat	Coef. Est.	t-stat	Coef. Est.	t-stat	Coef. Est.	t-stat
Intercept	?	0.082	0.76	0.074	0.96	0.067	1.35	0.059	1.38
Dual	+	0.112	1.75*			0.165	2.53***		
Wedge	+			0.055	1.22			0.151	2.04**
Beta	+	0.039	1.22	0.027	1.02	0.044	1.27	0.055	1.32
LnMV	-	-0.046	-1.72*	-0.022	-1.75*	-0.092	-1.79*	-0.091	-1.59
LnBM	+	0.136	2.50***	0.119	2.57***	0.182	2.58***	0.139	2.56***
Lev	+	-0.247	-2.87***	-0.238	-2.81***	-0.173	-1.37	-0.191	-1.28
EP	+	0.443	1.36	0.437	1.34	-1.384	-1.12	-1.232	-0.95
SAR	+	0.103	1.45	0.114	1.70*	-0.229	-1.58	-0.235	-1.92*
Accr	-	-0.144	-0.46	-0.132	-0.74	-0.157	-0.54	-0.143	-0.35
AbnER	+	1.445	3.57***	1.427	2.78***	1.845	2.18**	1.458	1.71*
Year		Yes		Yes		Yes		Yes	
Industry		Yes		Yes		Yes		Yes	
N		4,064		4,064		4,064		4,064	
Adj. R^2		0.238		0.242		0.402		0.407	

Table 8 presents results of future operating and stock performance over the period 1994-2006. All variables are defined in Table 2. For parsimony, the coefficients on year dummies and industry dummies are not reported. The t-statistics are based on Huber-White adjusted standard errors. ***, **, * indicate, respectively, the significance levels at 1%, 5%, and 10%.

Table 9
Within-Sample Analyses of Voting and Cash Flow Rights on Financial Reporting Quality

Panel A: The logit model of propensity to meet or just beat analysts' forecasts

$$\text{Prob}(MBE_{i,t} \in [\$0, \$0.01]) = F(\alpha_0 + \alpha_1 Wedge_{i,t} + \alpha_2 Wedge_{i,t}^2 + \alpha_3 CFRights_{i,t} + \alpha_4 CFRights_{i,t}^2 + \sum_{k=5}^{17} \alpha_k Controls_{i,t} + \varepsilon_{i,t}) \quad (9)$$

		Model 1		Model 2	
Variables	Pred. Sign	Coef. Est.	p-value	Coef. Est.	p-value
Intercept	?	-2.021	0.001	-1.888	0.001
Wedge	-	-0.069	0.011	-0.072	0.014
Wedge^2	?			0.042	0.063
CFRights	-	-0.322	0.001	-0.322	0.001
CFRights^2	?			0.052	0.114
Control Variables		Not Reported		Not Reported	
N		2,502		2,502	
Pseudo R^2		0.244		0.261	
Likelihood ratio		272.867		291.662	

Panel B: OLS regressions of abnormal accruals on voting and cash flow rights

$$|AbnAccr_{i,t}| = \beta_0 + \beta_1 Wedge_{i,t} + \beta_2 Wedge_{i,t}^2 + \beta_3 CFRights_{i,t} + \beta_4 CFRights_{i,t}^2 + \sum_{k=5}^{20} \beta_k Controls_{i,t} + \varepsilon_{i,t} \quad (10)$$

		Model 1		Model 2	
Variables	Pred. Sign	Coef. Est.	t-stat	Coef. Est.	t-stat
Intercept	?	1.243	3.20***	1.046	3.26***
Wedge	-	-0.078	-1.92*	-0.085	-1.80*
Wedge^2	?			0.147	2.62***
CFRights	-	-0.322	-2.83***	-0.319	-2.81***
CFRights^2	?			0.409	3.35***
Control Variables		Not Reported		Not Reported	
N		2,502		2,502	
Adj. R^2		0.104		0.109	

Panel C: The logit model of probability of releasing earnings reserves

$$\text{Prob}(AbnER_{i,t} < 0) = F(\kappa_0 + \kappa_1 Wedge_{i,t} + \kappa_2 Wedge_{i,t}^2 + \kappa_3 CFRights_{i,t} + \kappa_4 CFRights_{i,t}^2 + \sum_{k=5}^{12} \kappa_k Controls_{i,t} + \varepsilon_{i,t}) \quad (11)$$

		Model 1		Model 2	
Variables	Pred. Sign	Coef. Est.	p-value	Coef. Est.	p-value
Intercept	?	-1.760	0.001	-1.644	0.001
Wedge	-	-0.092	0.004	-0.082	0.008
Wedge^2	?			0.038	0.151
CFRights	-	-1.281	0.001	-1.013	0.001
CFRights^2	?			0.212	0.139
Control Variables		Not Reported		Not Reported	
N		2,502		2,502	
Pseudo R^2		0.142		0.172	
Likelihood ratio		367.632		381.885	

Table 9 presents results of earnings quality of dual-class firms over the period 1994-2006. *Wedge*, the square of *Wedge*, *CFRights*, and the square of *CFRights* are main test variables. *Controls* include all control variables in Equations (1)-(3) except *MgmtOwn*. All variables are defined in Table 2. ***, **, * indicate, respectively, the significance levels at 1%, 5%, and 10%.

Table 10
Financial Reporting Quality Following Changes in Share Structure

Panel A: The logit model of propensity to meet or just beat analysts' forecasts

$$\text{Prob}(MBE_{i,t} \in [\$0, \$0.01]) = F(\alpha_i + \alpha_1 Dual_{i,t} + \sum_{k=2}^{15} \alpha_k Controls_{i,t} + \varepsilon_{i,t}) \quad (1')$$

		Dual-class recapitalization		Share unification	
Variables	Pred. Sign	Coef. Est.	p-value	Coef. Est.	p-value
Dual	-	-0.238	0.022	-0.075	0.142
LnMV	+	0.249	0.014	0.184	0.028
LnMB	+	0.343	0.001	0.250	0.001
Lev	+	-0.594	0.041	-0.381	0.055
Option	+	0.232	0.021	0.146	0.010
Bonus	+	0.332	0.018	0.339	0.012
MgmtOwn	-	0.425	0.089	0.412	0.129
NOA	-	0.122	0.018	-0.052	0.068
RNOA	+	0.811	0.024	0.622	0.024
Loss	-	-0.253	0.048	-0.249	0.058
RDAdv	+	0.019	0.012	0.015	0.015
StdCFO	-	-0.031	0.006	-0.034	0.001
LnAnalysts	+	0.218	0.001	0.279	0.009
InstOwn	-	0.055	0.211	-0.017	0.161
ExpMgmt	+	0.443	0.004	0.433	0.001
N		476		836	
Pseudo R^2		0.280		0.252	
Likelihood ratio		381.532		322.024	

Panel B: OLS regressions of abnormal accruals on dual-class dummy and other controls

$$| AbnAccr_{i,t} | = \beta_i + \beta_1 Dual_{i,t} + \sum_{k=2}^{18} \beta_k Controls_{i,t} + \varepsilon_{i,t} \quad (2')$$

		Dual-class recapitalization		Share unification	
Variables	Pred. Sign	Coef. Est.	t-stat	Coef. Est.	t-stat
Dual	-	-0.241	-2.20**	-0.017	-1.22
LnMV	-	-0.083	-2.21**	-0.083	-2.43**
LnMB	+	0.123	3.01***	0.151	3.47***
Lev	+	-0.323	-2.94***	-0.383	-2.89***
StdCFO	+	0.067	1.38	0.901	2.24**
Option	+	0.024	1.85*	0.061	2.23**
Bonus	+	0.028	2.32**	0.070	2.14**
MgmtOwn	-	-0.234	-2.27**	-0.231	-2.11**
ExtFin	+	0.062	1.49	0.044	1.07
MBE	+	0.047	1.98**	0.070	2.31**
NOA	-	0.008	0.22	0.013	0.49
Loss	-	0.011	0.39	0.015	0.84
RDAdv	+	0.016	2.18**	0.009	2.39**
Herf	+	-0.443	-1.23	-0.232	-1.18
Segment	-	-0.165	-1.02	-0.142	-2.04**
LnOpCyc	+	0.075	1.86*	0.051	1.76*
InstOwn	-	-0.015	-1.88*	-0.033	-2.11**
Auditor	-	-0.094	-2.56***	-0.027	-2.25**
N		476		836	
Adj. R^2		0.172		0.148	

Panel C: The logit model of probability of releasing earnings reserves

$$\text{Prob}(AbnER_{i,t} < 0) = F(\kappa_i + \kappa_1 Dual_{i,t} + \sum_{k=2}^{10} \kappa_k Controls_{i,t} + \varepsilon_{i,t}) \quad (4')$$

		Dual-class recapitalization		Share unification	
Variables	Pred. Sign	Coef. Est.	p-value	Coef. Est.	p-value
Dual	-	-0.134	0.001	-0.122	0.022
Option	+	0.074	0.121	0.068	0.090
Bonus	+	0.472	0.007	0.481	0.001
MgmtOwn	-	-1.132	0.004	-1.045	0.005
ExtFin	+	0.333	0.001	0.314	0.001
MBE	+	0.089	0.042	0.072	0.031
Lev	+	-0.874	0.001	-0.226	0.081
StdCFO	+	0.218	0.034	0.239	0.005
LnMV	-	-0.038	0.118	-0.024	0.213
InstOwn	-	0.013	0.173	-0.019	0.148
N		476		836	
Pseudo R^2		0.147		0.159	
Likelihood ratio		314.247		321.383	

Table 10 presents results of earnings quality of firms that either underwent dual-class recapitalization or abolish dual-class structures over the period 1990-2006. A fixed-effect panel structure is employed and firm-specific fixed effects are not reported. To be included in the sample, a firm must have at least three years' data before and after the recapitalization/unification event. *Controls* include all control variables in Equations (1)-(3), respectively. All variables are defined in Table 2. ***, **, * indicate, respectively, the significance levels at 1%, 5%, and 10%.