Corporate Governance, Idiosyncratic Risk, and Information Flow*

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ABSTRACT

We study the relationship between measures of corporate governance policy related to anti-takeover restrictions and idiosyncratic risk in stock returns. Fewer governance provisions that restrict openness to control offers are associated with high levels of idiosyncratic risk, trading activity, private information flow, and more information about future earnings in stock prices. Trading interest by institutions, especially those active in merger arbitrage, accentuates the relationship of governance to idiosyncratic risk and information flow. Our results show that openness to the market for corporate control indicates more informative stock prices by encouraging collection of and trading on private information. Consistent with the information-flow interpretation of idiosyncratic risk, we find that the component of volatility unrelated to governance is associated with the efficiency of corporate investment.

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The effect of corporate governance mechanisms on equity prices and the distribution of returns is an important issue in modern corporate finance. Evidence suggests that governance mechanisms can directly influence equity value [e.g., Gompers, Ishii, and Metrick (2003a) and Cremers and Nair (2004)]. Constraints and incentives influencing management action are generally posited as the mechanism by which governance influences prices. However, any systematic effect on returns also requires a link from governance to investors' expectations or information flow. For example, Gompers et al. (2003a) argue that in the early 1990s investors might not have fully appreciated the extent of the agency costs engendered by weak governance mechanisms. Nonetheless, the link between governance and investors' information have been little studied.[1] This paper extends the current understanding by showing how governance mechanisms and informed trading interact to influence the incorporation of information into stock prices.

We develop and test a trading-link hypothesis showing how the specific aspects of governance that influence takeover vulnerability are linked to stock price informativeness. One specific route is the trading behavior of arbitrage-oriented institutional investors. We reason that the absence of anti-takeover governance provisions creates incentives to collect private information. When trading activity is generated, it further contributes to idiosyncratic volatility, or firm-specific stock return variation, and other indications of private information flow. Our reasoning is in the tradition of Grossman and Stiglitz (1980), who predict that improving the cost-benefit trade-off on private information collection leads to more extensive informed trading and to more informative pricing.

Fewer takeover restrictions could imply more private information collection via several paths. First, fewer restrictions might imply a higher probability of a takeover [Ambrose and Megginson (1992)], providing traders more incentive to speculate on the possibility of a takeover. Larcker and Lys (1987) show that speculators in takeover situations are better

[1] A recent important exception is the paper by Jin and Myers (2005). They show that country-level governance (investor protection) and disclosure are important for the holding-back of information by insiders, which can eventually lead to volatility via a crash.

informed about the likelihood of success, suggesting they have indeed collected private information. Jindra and Walkling (2004) show that offer prices are closer to market prices when there is a large price run-up prior to the offer - exactly what should occur if speculators are collecting and trading on private information in the pre-offer period. Second, fewer takeover restrictions could indicate that management is not expecting a control offer [Comment and Schwert (1995)], with the implication that speculators can profit from correctly anticipating that the probability of an offer is greater. Third, fewer takeover restrictions could indicate that the management/board would have limited bargaining power should a control offer occur [Comment and Schwert (1995) and Bebchuk, Coates, and Subramanian (2002)], thereby interesting speculators who would prefer to tender into an offer.[2] The probability of an offer is enhanced, albeit indirectly, due to the increased probability of success [Grinblatt and Titman (2002), page 725]. Offers are beneficial to speculators - even failed attempts often lead to increased value [Saffiedine and Titman (1999)]. Finally, strong investor protection, expressed by openness to takeovers, is associated with a lower possibility of insiders (controlling shareholders and managers) expropriating arbitrageurs and outside investors. Thus, openness can directly encourage uninformed ownership and trading, thereby providing more cover for privately informed trading and indirectly encouraging it. This final possibility, especially, suggests that governance provisions can affect information flow even when mergers are not imminent.

Our core empirical result is that there is a strong negative relation between the extent of a firm's anti-takeover provisions - the IRRC index used by Gompers et al. (2003a) - and the amount of firm-specific information impounded into stock prices as measured by idiosyncratic volatility (i.e., lack of synchronicity with other stocks in the market).

In view of the fact that the IRRC index includes some non-takeover-related governance characteristics, we corroborate our explanation's focus on anti-takeover measures by showing that a subset of powerful anti-takeover provisions is a particularly important empirical de-

[2] Kahan (2004) provides a dissenting view, although his evidence does not directly show whether takeover defenses affect a company's ability to resist offers.

terminant of idiosyncratic volatility. This subset of provisions is suggested by Cremers and Nair (2004) and includes, for example, the presence of a staggered board structure. These provisions are a deterrent to control offers in our sample, consistent with private information collection and trading being more attractive for stocks of companies that lack these provisions.[3]

Stock return variation has long been considered indicative of the rate of incorporation of private information into stock prices. Ross (1989) shows that price volatility is directly related to the rate of information arrival (or information flow), as an "important consequence of arbitrage-free economics." Price reactions to information occur as arbitrage opportunities are exploited, or else to prevent them. By comparing variances between business and non-business periods, French and Roll (1986) develop evidence that "private information causes most stock price changes." Strategic models from the microstructure literature establish that informed trade induces volatility [e.g., Kyle (1985) and Glosten and Milgrom (1985)]. Of course, in markets with limits to arbitrage, pricing errors and noise also manifest in volatility, and private information is likely more common with respect to industries and firms, rather than to the broad market. Roll (1988) therefore focuses specifically on idiosyncratic volatility, providing evidence that idiosyncratic price changes are driven at least seven times more by information than by noise even on days with no identifiable public information. Thus, idiosyncratic volatility is a good candidate as a summary measure of information flow, and especially for private information about firms.

Recent empirical evidence supports this informational interpretation of idiosyncratic volatility. High levels of idiosyncratic volatility are associated with more efficient capital allocation: U.S. industry-level evidence is provided by Durnev, Morck, and Yeung (2004), and international evidence by Wurgler (2000). Furthermore, U.S. industries with high levels

[3]Coates (2000) and Kahan (2004) argue that the development of corporate case law has rendered anti-takeover provisions into bargaining and delaying tools that are incapable of shutting down a takeover plan. In that case, they are more indicative of management's expectation of control activity than of entrenchment. The cited legal-scholar views are tempered by dissenting reasoning, as in Bebchuk et al. (2002) - but even if takeover restrictions can prevent a change of control, speculating on the possible outcome may still be of interest given that the attempt is more likely.

of idiosyncratic volatility have stock prices that are more informative about future earnings [Durnev, Morck, Yeung, and Zarowin (2003)]. Cross-country patterns of idiosyncratic volatility correspond to likely patterns of price informativeness. Morck, Yeung, and Yu (2000) find low idiosyncratic volatility in emerging markets but high firm-specific stock return variation in developed markets. Jin and Myers (2005) show that poor country-level governance together with opaqueness (lack of accounting transparency) explains low idiosyncratic volatility. Overall, a body of empirical evidence supports the use of idiosyncratic volatility as measure of stock price informativeness and of the extent to which stock prices accurately and timely incorporate firm-specific information.[4]

Buttressing an informational interpretation of our core result, we find that takeover vulnerability is similarly related to several alternative measures of private information flow. These include the stock's turnover, the PIN (probability of information-based trading) measure of Easley, Hvidkjaer, and O'Hara (2002), and the private information trading measure of Llorente, Michaely, Saar, and Wang (2002).[5] Additionally, we find that stock prices in industries with less anti-takeover provisions contain more information about future earnings.

We also provide some direct evidence on trading as a route for the relationship from governance to idiosyncratic risk. Specifically, that relationship persists and is more pronounced for stocks with intense trading by institutional investors, and in particular those that have recently been involved in risk arbitrage around takeovers. Thus, at least one of the links from governance to volatility and information flow is via arbitrageur institutions. To our

[4] One particular concern is that private information that is prevented from expressing itself sooner will express itself later with the same effect on volatility. Therefore, alleviating barriers to information transmission may not affect the average level of volatility if the underlying natural rate of information generation [Ross (1989)] is unchanged [see also Damodaran (1985)]. One response, explored by Durnev et al. (2004), is that the true stock value may be mean-reverting, with the implication that unexploited information depreciates. In the microstructure literature, Foster and Viswanathan (1995) posit conditions under which even short-lived private information that is publicly revealed at the end of each trading period will influence volatility via its influence on intra-period trading. Also Madhavan, Richardson, and Roomans (1997) develop a model in which asymmetric information and microstructure sources of volatility reinforce each other.

[5] Regarding turnover, rational expectations models where volume conveys information, such as Kim and Verrecchia (1991a), imply that expected volume and volatility are positively correlated or proportional. If in fact openness to the market for control is tied to the flow of privately-generated information, then governance also should be related to the intensity of the resulting trading.

knowledge, such a link has not been previously documented in the literature.

We develop the above results using a large panel of firms for the 1990-2001 period. We control for the possibility of spurious correlations. For example, larger firms might be both less volatile and also have more anti-takeover measures. In view of such possibilities, our study uses regressions with a large set of covariates (e.g., size, return on equity, and firm age) from the recent literature on idiosyncratic risk [e.g., Pastor and Veronesi (2003) and Wei and Zhang (2006)]. We also provide a variety of other robustness checks on our results and interpretations. For example, in view of the results of Bushee and Noe (2000), we control for the transparency of firms' financial reporting and find that accounting opaqueness is negatively correlated with idiosyncratic volatility - but that this does not substitute for or replace the governance-risk relation. More richly, we show how transparency interacts with governance to influence private information collection, extending the cross-country results of Jin and Myers (2005) to the firm level.

While the panel results establish a link between governance and idiosyncratic volatility, they do not directly address concerns of reverse-causality. To consider more directly whether takeover restrictions are economically causal, we provide a set of time series results. These focus directly on changes in idiosyncratic volatility plausibly induced by the adoption of takeover restrictions, comparing idiosyncratic volatility before and after takeover restrictions adoption. If we are simply documenting self-selection (that is, firms for which private information collection is light are more likely to adopt anti-takeover provisions), we should not find a significant change in idiosyncratic volatility around the event. In contrast to that view, we find that idiosyncratic volatility is lower after a firm adopts takeover restrictions and higher after a firm drops takeover restrictions. Furthermore, the relationship between governance and idiosyncratic volatility is stronger around the time of takeover events.[6]

[6] Core, Guay, and Rusticus (2005) find no link between governance and surprises in operating performance, raising questions about a governance-stock return link. One implication of our findings is that a key link is through second moments. Several papers consider a link between expected returns and idiosyncratic volatility [e.g., Goyal and Santa-Clara (2003), Ang, Hodrick, Xing, and Zhang (2006)]. Added to our findings, these suggest the possibility of a governance to volatility to expected returns link.

In addition to contributing to the understanding of the importance of corporate governance provisions, our results contribute to the broad literature on idiosyncratic risk determinants, establishing that governance variables are a determinant of idiosyncratic risk in addition to those previously uncovered. We also contribute to the more specific literature that emphasizes firm-specific information flow as a key driver of idiosyncratic volatility. As noted above, the information-flow interpretation emphasizes a particular aspect of idiosyncratic volatility, and is therefore somewhat tentative. Our results, which *positively* link arbitrageur trading and idiosyncratic risk, strengthen the information content interpretation of idiosyncratic risk.[7]

To conclude our analysis, we strengthen the interpretation of idiosyncratic risk as a measure of stock price informativeness by incorporating our governance-risk relation into an analysis of the link between idiosyncratic risk and the quality of corporate investment decisions. Durnev et al. (2004) have shown that corporate investment decision-making quality is increasing in idiosyncratic volatility. Since good capital budgeting is one expression of good governance, there is the possibility that the apparent relation of idiosyncratic risk to investing quality is statistical stand-in for a underlying economic relation of governance provisions to investing quality. Takeover restrictions might be important in that they could entrench current management [Bebchuk et al. (2002)] providing safety for poor investment decision makers. Other studies have found conflicting results, but have not taken into account the concurrent relation with volatility. For example, Larcker, Richardson, and Tuna (2004) find only weak relations of governance to the quality of investment decisions for a broad sample of companies.

We find that anti-takeover provisions are associated with a tilt toward more investment for firms that seem otherwise to be underinvesting. Moreover, with our finding of a governance-

[7]Shleifer and Vishny (1997), for example, take a different view in the context of delegated arbitrageurs akin to the institutions on which we focus. They reason that idiosyncratic volatility, interpreted as noise that may drive price further away from fundamentals before they converge, inhibits arbitrage. Investors, being unable to distinguish between recent losses due to incompetence and those due to temporarily increasing divergence from fundamentals, are tempted to limit arbitrageurs' capital at just the time when trading benefits may be greatest - which would tend toward to a volatility-trading link opposite to what we observe.

volatility relation, we have an instrument for volatility. That is, we can separate the level of volatility that is "expected" given the governance structure from the remaining "unexpected" volatility, and consider the relation of capital budgeting quality to each one. We find that "expected" volatility does not influence investing quality. The fundamental relation is thus between "unexpected" idiosyncratic volatility and investing quality, thus supporting our reliance on the information-flow interpretation of volatility. Our results imply that information flow dominates the effects of anti-takeover governance provisions for investment decision-making quality, though both are important.

The next section describes our data and key measures of idiosyncratic volatility, governance characteristics, and other control variables. Section II presents our core evidence on the relation of idiosyncratic volatility and anti-takeover governance provisions. Section III presents supporting evidence and additional interpretation considering endogeneity concerns, alternative measures of private information flow, and the trade-link hypothesis. Section IV considers the importance of accounting transparency for the information flow and presents robustness checks. Section V considers the relationship of governance provisions, idiosyncratic volatility, and the quality of corporate investing decision making. Section VI concludes.

I. Data and Measures

The data for our study is drawn from the Investor Responsibility Research Center (IRRC), Center for Research in Stock Prices (CRSP), Compustat, and also from Thomson Financial's institutional ownership database compiled from S.E.C. 13F filings. Our initial sample includes all companies in the IRRC database from 1990 to 2001. To enhance cross-sectional comparability, we omit financial firms and utilities (SIC 6000-6999 and 4900-4999). After imposing this requirement, the number of firm varies across time, but on average there are 1248 firms with a minimum of 1027 in 1992 and a maximum of 1526 in 1998.

For all sample firms, we construct our measure of idiosyncratic volatility and control variables. Idiosyncratic volatility is measured using daily returns from CRSP. The average number of firms with both governance rankings and idiosyncratic volatility is 1140 (the minimum is 943 in 1992 and the maximum is 1514 in 1998). Control variables are calculated either from CRSP, Compustat, and the 13F databases. Some control variables require a longer panel, starting prior to 1990. For these reason, we expand the sample backwards to 1987. We winsorize extreme observations at the bottom and top one-percent levels to avoid spurious inferences.

A. Idiosyncratic Volatility

We study idiosyncratic volatility for each stock, estimated for each month using daily return data. Our strategy to measure idiosyncratic volatility is based on a regression projection of equity returns (alternatively) on the returns of the market index, industry index, or other factors.

Consider the case of the market model first. The projection of stock i's excess return on the market is:

$$r_{id} = \alpha_i + \beta_i r_{md} + e_{id}, \tag{1}$$

with $\mathrm{E}(e_{id}) = \mathrm{Cov}(r_{md}, e_{id}) = 0$. In equation (1), r_{id} is the return of stock i on day d in excess of the risk-free rate and r_{md} is the value-weighted excess market index return defined as $r_{md} \equiv \sum_i w_{id} r_{id}$ with w_{id} the weight of firm i at day d. Then $\beta_i = \frac{\sigma_{im}}{\sigma_m^2}$, where $\sigma_{im} = \mathrm{Cov}(r_{id}, e_{md})$, and $\sigma_m^2 = \mathrm{Var}(r_{md})$. From this projection, idiosyncratic variance is estimated as:

$$\sigma_{ie}^2 \equiv \sigma_i^2 - \sigma_{im}^2/\sigma_m^2, \tag{2}$$

where $\sigma_i^2 =$Var(r_{id}) We use sums of squares of daily returns within each month t to estimate monthly variances σ_{it}^2 and σ_{mt}^2, and sums of cross-products to estimate covariances $\sigma_{im,t}^2$. Idiosyncratic volatility in equation (2) is equivalent to that in the regression strategy provided that coefficients and the residuals are calculated using data from the same time period.

We first present our results using idiosyncratic volatility estimated using the market model. We examine the robustness of our results using alternative models of idiosyncratic volatility: the Fama and French (1992) three-factor model and an industry model (instead of the market model). Estimation of the idiosyncratic volatility for multi-factor models is done similarly.

We also estimate each stock's residual volatility share, that is, the ratio of idiosyncratic volatility to total volatility $\sigma_{ie,t}^2/\sigma_{it}^2$, or relative idiosyncratic volatility. This is precisely $1 - R_{it}^2$ of equation (1). Given the bounded nature of R^2, we conduct regression tests using the logistic transformation of $1 - R_{it}^2$:

$$\Psi_{it} = \ln\left(\frac{1 - R_{it}^2}{R_{it}^2}\right) = \ln\left(\frac{\sigma_{ie,t}^2}{\sigma_{it}^2 - \sigma_{ie,t}^2}\right). \tag{3}$$

Thus, our dependent variable Ψ_{it} measures idiosyncratic volatility relative to market-wide variation.[8] One reason for constructing $\Psi_{i,t}$, i.e., scaling idiosyncratic volatility by the total variation in returns, is that some firms are more subject to economy-wide shocks than others, and firm-specific events in these industries may be correspondingly more intense. We also do this as needed for comparability to other studies; see, for example, Durnev et al. (2004) and Jin and Myers (2005).

Panel A of Table I presents univariate statistics for $\sigma_{ie,t}^2$ (annualized), $\sigma_{ie,t}^2/\sigma_{it}^2$, and Ψ_{it} over the entire sample period (January 1990 - December 2001). For this table, we have estimated volatility measures via regressions within each sample month t. The mean idiosyncratic variance (annualized) is 19.4×10^{-2}, which corresponds to a annualized standard deviation of 44%. Idiosyncratic volatility represents more than 85% of total individual stock volatility, on average.

[8] We can use the absolute idiosyncratic volatility $\sigma_{ie,t}^2$ as alternative to the relative idiosyncratic volatility Ψ_{it}. In this case, we need to control for systematic volatility. The results using this alternative specification (not tabulated here) are broadly consistent with out primary finding of a negative association between governance and idiosyncratic volatility. Moreover, alternative transformations of variance, such as logarithm of variance and standard deviation, also lead to qualitatively similar results.

B. Corporate Governance Index

A key independent variable in our work is provided by the IRRC database. This variable is the governance index, which we will denote as G, used in Gompers et al. (2003a).[9] The index G is constructed for each sample firm for 1990, 1993, 1995, 1998, and 2000 from observations on a set of mostly anti-takeover related governance provisions tracked by the IRRC. The provisions cover such things as tactics for delaying hostile bidders, voting rights, officer/director protection, and state laws limiting takeover bids. The index is formed by totaling up indicator variables for each of the 24 non-overlapping provisions for each firm in the IRRC universe in a given year. Larger values of the governance index G indicate a firm that is more insulated from takeovers and, in the judgment of Gompers et al. (2003a), less shareholder-friendly. The index has a potential range from zero to 24. Panel B of Table I presents summary statistics for G. The median G is 9.0 and standard deviation is 2.8.

For some of our tests we use the raw index G for each firm; for robustness, we always double-check the results against tests using a dummy-variable version, GD, which is coded as zero if the governance index is less than or equal to five (portfolio *open* to takeover activity in the sense of having few limiting provisions) and one if the index is greater than or equal to 14 (*closed* portfolio). When we use GD, we exclude firm/years with intermediate index values (i.e., between six and 13).

We also conduct tests using the Cremers and Nair (2004) anti-takeover index (ATI), which incorporates only three provisions, the ones more closely related to takeovers. This index depends on the subset of provisions in the IRRC-Gompers et al. (2003a) index that are thought to be most effective in deterring takeover activity and/or increasing target bargaining power. The Cremers and Nair (2004) index varies from zero to three, with one point being accorded for the blank-check preferred stock authorization, one point for a classified (staggered elections) board structure, and one point for limitations on shareholder's ability to call special meetings or act by written consent. Gompers et al. (2003a) call these

[9] We thank Andrew Metrick for providing data on the governance index linked to CRSP permnos.

"delay" provisions, since they might hold up a takeover. Ambrose and Megginson (1992) find that blank check preferred, a prerequisite for a poison pill defense, is negatively correlated with acquisition likelihood. We denote this index as ATI, for anti-takeover index.

With our convention that larger values for G, GD, and ATI correspond to more anti-takeover provisions, our indexes are inverse measures of firms' openness to the market for corporate control. When we need to specify a governance index for a particular month t, we use the most-recently announced level.

C. Alternative Measures of Private Information

To substantiate our informational interpretation of the governance-volatility relationship, we also test for the relation between governance and several dependent variables that are directly related to the level of trading activity and private-information trading. We consider three alternative proxies for the intensity of private information flowing to a stock's market.

We calculate a firm-level monthly turnover ($TURN$) time series by dividing monthly share volume by the number of shares outstanding. In separate tests, we use the annually-observed probability of information-based trading measure (PIN) of Easley et al. (2002).[10]

Additionally, we estimate the private information trading measure suggested by Llorente et al. (2002). Specifically, for each year, we estimate the firm-level time-series regression

$$r_{id} = b^a_{i0} + b^a_{i1} r_{i,d-1} + b^a_{i2} r_{i,d-1} V_{i,d-1} + \epsilon^a_{id}, \tag{4}$$

where V_{id} is log daily turnover detrended by subtracting a 200 trading day moving average. The b^a_{ij}'s are coefficients and ϵ^a_{id} is the residual. The amount of private information trading ($PRIVATE$) is given by the regression coefficient b^a_{i2} on the interaction variable. With this procedure, we have one observation on $PRIVATE$ for each firm/year. Panel C of Table I provides summary statistics on these alternative measures of private information.

[10] We thank the Soeren Hvidkjaer for making data on this measure available on his website.

D. Earnings Information in Stock Prices

We also consider the relation between governance and measures of the extent to which stock prices incorporate information about future earnings (annually measured), which are calculated at the industry level. Following Durnev et al. (2003), we estimate two variables of future earnings response. $FERC$ is the future earnings response coefficient, which is given by the sum of the coefficients $\sum_{\tau=1}^{3} b_{2,\tau}^{b}$ on future changes in earnings in the regression

$$r_{it} = b_0^b + b_1^b \Delta E_{it} + \sum_{\tau=1}^{3} b_{2,\tau}^{b} \Delta E_{i,t+\tau} + \sum_{\tau=1}^{3} b_{3,\tau}^{b} r_{i,t+\tau} + \epsilon_{it}^{b}, \tag{5}$$

where r_{it} represents the annual stock return for each firm i in a particular industry during year t, calculated from fiscal-year end share price plus dividends adjusted by stock splits and such, as reported on Compustat (annual items #199/#27 plus #26/#27); ΔE_{it} is annual change in earnings per share (earnings before interest, taxes, depreciation, and amortization, annual item #13) scaled by previous fiscal year end market capitalization (annual items #199 times #25).[11] The regression is performed annually on a two-digit SIC industry cross section of firms with at least ten observations. We use two-digit SIC industries instead of four-digit as in Durnev et al. (2003) because we are restricted to the smaller sample of firms for which we have G data.

$FINC$ is the future earning explanatory power increase which is given by the increase in the coefficient of determination (R^2) of the regression (5) relative to the base regression

$$r_{it} = b_0^c + b_1^c \Delta E_{it} + \epsilon_{it}^{c}. \tag{6}$$

Panel D of Table I provides summary statistics on this measures of future earnings information contained in current stock returns, $FERC$ and $FINC$.

[11] Note that the time subscript t in this regression pertains to annual increments. In most other settings, we use t to indicate monthly increments. We always note when we use the annual convention, to avoid both confusion and the proliferation of notation.

E. Institutional Risk Arbitrage Trading Activity

In explaining idiosyncratic risk and measures of private information flow, we control for the level of institutional trading as in Piotroski and Roulstone (2004), and, more to the point, for the trading of risk-arbitrage-oriented institutions. We also consider the interaction between the governance index and institutional trading/arbitrage trading.

Our proxy for institutional trading in a stock, $INST$, is the quarterly average of the absolute change in the number of shares held by each institutional investor as a fraction of the stock's annual trading volume. The holdings data is from the Thomson Financial's institutional ownership database compiled from S.E.C. 13F filings. When we need to match holdings quarterly data with monthly data, we use the most-recently observed quarterly figure.

We also require a proxy for the intensity of trading by arbitrage-intensive institutional investors. $INSTA$ is quarterly average of the absolute change in the number of shares held by each merger arbitrage-active institution as a fraction of annual trading volume. For each calendar quarter, we define an arbitrage-active institution as one that has bought into multiple merger situations in at least one of the previous eight calendar quarters. We define a "merger situation" as existing for the shares of any company when an acquisition offer is outstanding during the quarter, according the the Thomson Financial SDC Platinum M&A database. We define "buying into multiple merger situations" in terms of meeting *both* of the following two conditions. First, for at least three merger situations in the quarter, the institution either began the quarter owning at least one percent of a merger-situation firm's common shares outstanding and did not decrease its holdings over the quarter, or else, if it began the quarter with smaller holdings, increased its holdings by at least one-half percent of the shares outstanding. Second, the institution buys at least three percent of shares outstanding for at least one merger-situation. We have experimented with variations of these rules without much effect on the outcome. The point is to identify institutions that have recently been willing to take substantial positions speculating on M&A situations.

Panel E of Table I provides descriptive statistics for institutional risk arbitrage trading activity variables, $INST$ and $INSTA$.

F. Control Variables

A number of control variables are part of our empirical design. Recent studies, e.g. Pastor and Veronesi (2003) and Wei and Zhang (2006), use firm characteristics to explain the cross-section of individual firm idiosyncratic volatility. Our firm-specific control variables include return on equity, variance of return on equity over the past three years, leverage, market-to-book ratio, firm size (measured by market capitalization), and firm age since listing (proxied by CRSP appearance). These measures are drawn either from CRSP (in the case of age and market capitalization) or quarterly Compustat (in the case of accounting variables). When we need to match accounting quarterly data with monthly data, we use the most-recently observed quarterly figure. Quarterly earnings report date is used to determine when the information is available to investors (typically two months after the end of a firm's fiscal quarter).

Return on equity (ROE) for stock i in each month t is the stock's most recent quarterly earnings (earnings before extraordinary items, Compustat quarterly item #8) divided by the book value of equity (quarterly item #60). The quarterly ROE is multiplied by four to annualize.

The variance of ROE used in this study, $VROE$, for stock i in month t, is the sample variance of quarterly ROEs over the last three years. $VROE$ captures the uncertainty about future ROE's. Wei and Zhang (2006) find that both ROE and $VROE$ are powerful determinants of the cross-section of idiosyncratic volatility. They find that firms with lower ROE and higher $VROE$ have higher idiosyncratic volatility.[12]

[12] Following Wei and Zhang (2006), we estimate $VROE$ using only data available to investors at each time period by taking the sample variance of quarterly ROE over the last three years. We also have constructed an alternative $VROE$ variable as the residual standard deviation from an AR(1) process for firm-level quarterly ROE. This $VROE$ variable uses all data available for each firm (with a minimum of five years of consecutive ROE) and does not change over time. Panel regressions results (not tabulated here) using the alternative

LEV is the quarterly firm leverage defined as the ratio of long-term debt (quarterly item #51) to total assets (quarterly item #44). More highly levered firms are expected to have higher default risk and consequently higher idiosyncratic volatility. M/B is the quarterly log of the market-to-book equity ratio (end-of-quarter market value of equity is from CRSP and book value of equity is Compustat quarterly item #60). $SIZE$ is the monthly log of the firm i equity size given by its market capitalization in month t. DD is a quarterly dividend dummy which equals one if the firms pays dividends and zero otherwise (dividends given by Compustat quarterly item #20). AGE is the monthly firm age defined as the log of the number of years a firm has been in CRSP database, as a proxy for how long the firm has been listed. $DIVER$ is a quarterly dummy variable for firms operating in multiple business segments as given by Compustat. Following common practice in the literature, we winsorize these variables at bottom and top one-percent levels to avoid the possibility of results driven by extreme outliers.

In some tests, we also control for firm-level accounting transparency. We follow the accounting literature [e.g., Francis, LaFond, Olsson, and Schipper (2005)] in using several measures of abnormal accruals as an inverse measure of accounting transparency. Our procedure differs from Francis et al. (2005) in that we employ annual cross-sectional two-digit SIC industry regressions to benchmark accruals, whereas they work with time-series regressions.[13] Given the panel nature of our study, we require measures that can evolve over time. Using the same naming convention as in Francis et al. (2005), our first measure is called $EQ2$. $EQ2$ is defined as the absolute value of firm-specific residuals from an industry regression of total accruals on (the reciprocal of) total assets (Compustat annual item #6), revenue (annual item #12) growth, and fixed assets (annual item #7).[14] $EQ2$ is nearly the same as

$VROE$ are qualitatively the same.

[13]Technically, the calculation involves a near-residual, not the actual residual. The industry regression is calculated using income-statement revenues as a regressor. The fitted value that is netted from a firm's total accruals to form this near-residual is cash revenues (i.e., income statement revenues adjusted for the change in accounts receivable). The same adjustement is employed for the alternative measures. See Francis et al. (2005) for details.

[14]Total accruals is equal to the change in current assets (annual item #4) minus the change in current liabilities (annual item #5) plus the change in debt in current liabilities (annual item #34) minus the change

Teoh, Welch, and Wong (1998) measure of earnings quality. As an alternative measure, $EQ5$ is the absolute value of firm-specific residuals from an industry regression of total accruals on lagged, contemporaneous, and leading cash flow from operations [Dechow and Dichev (2002)].[15] Note that all these measures are direct indexes of opacity and inverse indexes of transparency, in that they increase in the magnitude of unexpected accruals. When we need to match these annual measures with monthly data, we use the most-recently observed annual figure.

Panel F of Table I provides descriptive statistics for the firm-specific control variables.

II. Governance, Idiosyncratic Volatility, and Information Flow

We present univariate statistics and panel regression evidence on the relation between governance and idiosyncratic volatility.

A. Univariate Statistics and Graphical Analysis

Figure 1 present a visual overview of our data for market-model idiosyncratic volatility σ_{ie}^2 (annualized idiosyncratic variance), and Figure 2 presents an overview for relative idiosyncratic volatility σ_{ie}^2/σ_i^2. Following Gompers et al. (2003a), we aggregate firms with a G index of less than or equal to five into an *open* portfolio that has relatively few takeover impediments in its governance structure. Firms with ratings of 14 or more are aggregated into a *closed* portfolio that is relatively insulated from potential takeovers. These portfolios correspond to our construction of the dummy variable GD above.

Panel A of Figure 1 plots the equal-weighted averages of idiosyncratic volatility over the full sample period (1990-2001) within each level of the governance index. A much higher level

in cash (annual item #2), and minus depreciation and amortization (annual item #14).

[15] Cash flow from operations is equal to earnings before extraordinary items (annual item #18) minus total current accruals (total accruals plus depreciation and amortization).

of idiosyncratic risk is associated with governance structures that are very open to control offers. For the full sample period 1990-2001, the open portfolio displays an idiosyncratic variance level of about 23.4×10^{-2}, which corresponds to an annualized standard deviation of 48.4%. The corresponding figure for the closed portfolio is about 13.5×10^{-2}, which corresponds to an annualized standard deviation of 36.7%. The difference between the two levels is highly statistically significant with a t-statistic of 27.74. The same comparison holds in every year, as shown in Panel B of Figure 1.

Panel A of Figure 2 similarly plots the average relative idiosyncratic volatility over the full sample period according to the level of the governance index. Our core result is clear in the figure: average idiosyncratic volatility for the open portfolio is clearly greater than for the closed portfolio. The open portfolio displays a relative idiosyncratic volatility of about 0.868. The corresponding figure for the closed portfolio is about 0.854. The difference between the two extreme portfolios is highly statistically significant with a t-statistic of 5.31. Moreover, all intermediate governance index portfolios present lower relative idiosyncratic volatility than the open portfolio, on average.

Panel B of Figure 2 plots the annual time series of the relative idiosyncratic volatility difference between the closed and open portfolio. Average relative idiosyncratic volatility is greater for the open portfolio than for the closed portfolio in every year, except at the very end of the sample period. The difference is statistically significant in the majority of the years.

B. Empirical Framework

The figures in the previous section establish that firms with governance structures that are open to control offers have greater relative idiosyncratic volatility, on average (absent of firm-level controls). These differences potentially could be driven by non-governance factors that are incidentally correlated with governance. In the remainder of this section, we establish that anti-takeover aspects of governance are at the core of the relation. To do so, we estimate the following regression equation:

$$\begin{aligned}\Psi_{it} &= c_0 + c_1 GOV_{i,t-1} + c_2 ROE_{i,t-1} + c_3 VROE_{i,t-1} + c_4 LEV_{i,t-1} + c_5 M/B_{i,t-1} \\ &\quad + c_6 SIZE_{i,t-1} + c_7 DD_{i,t-1} + c_8 AGE_{i,t-1} + c_9 DIVER_{i,t-1} + \epsilon_{it},\end{aligned} \tag{7}$$

where i indexes firms, t is a monthly time index, and GOV is a particular measure of governance provisions.[16] The additional regressors are volatility correlates as described earlier. Regressions include 2-digit SIC industry fixed effects that control for additional differences across industries. Most of our later tests are based on variations of this basic regression.

We alternatively set GOV equal to each of the IRRC-Gompers et al. (2003a) measures we have already described. The first, G, is simply their index of inverse governance quality or openness to takeover activity. The second, GD, is the indicator variable for anti-takeover provisions: open portfolio firms have $GD = 0$, whereas closed portfolio firms have $GD = 1$. When we use GD, we exclude firm/years with intermediate index values for which GD is undefined. Lower values of G and GD correspond to fewer anti-takeover provisions.

We also investigate whether the correlation between volatility and the IRRC-Gompers et al. (2003a) measure is tightly linked to anti-takeover provisions versus other aspects of governance. To do so, we set GOV equal to the Cremers and Nair (2004) anti-takeover provisions index, ATI.[17]

We are most centrally interested in the value of c_1 in each specification, as it provides information on the relationship of idiosyncratic risk to the measure of corporate governance. In this section, we estimate this regression as a pooled cross-sectional/time-series model. Later, in the robustness section, we also use Fama and MacBeth (1973) methodology. The

[16] Results in the robustness section show that annual or quarterly regressions lead to qualitatively similar results.

[17] The G index is mainly related to anti-takeover provisions, but some of the provisions are also related to other aspects of corporate governance such as board structure. Separate results (not tabulated here) show that proportion of insiders on the board does not have any significant correlation with the probability of a control offer and that it is not negatively correlated with idiosyncratic volatility. By ruling out the possibility that the G index is proxying for the board makeup, we further substantiate our belief that the relation of G and idiosyncratic risk is due to anti-takeover measures.

results are similar in each case.

Endogeneity is a well-known issue in governance regressions. As a first response, we always regress volatility on predetermined measures of governance characteristics and other variables. In the case of governance measures, we work with the most recent data on the Gompers et al. (2003a) index, which lags by up to three years. In the case of other variables, we use the most recent observation. In a later section, we provide time series tests that buttress our panel regression results, as one way to address remaining concerns about endogeneity.

C. Panel Regression Results

Table II presents estimates of the basic model in equation (7) in which the logistic transformed relative idiosyncratic volatility Ψ is the dependent variable. The table reports restricted versions of the basic model where a governance index is the only regressor, as well as full models with the complete set of control variables.

Columns (1), (3), and (5) display the restricted estimates. The consistent result is a significant positive relation between a governance stance open to takeovers, and idiosyncratic volatility. In column (1), for example, the regression coefficient on the G index is -0.0289 with a Newey-West robust t-statistic of -14.70. Higher levels of the index indicate less openness, so the relationship is clear. The same conclusion can be drawn from column (3), which uses GD, the dummy-variable indicator of openness. Here, the estimated coefficient is -0.2173 with a t-statistic of -7.29. Also, the same negative relation between anti-takeover provisions can be found when we use ATI, the anti-takeover index. The ATI estimated coefficient is -0.0875 with a t-statistic of -14.98.

Controlling for firm characteristics does not change the qualitative result, although the coefficients and robust t-statistics decrease. Estimates are reported in columns (2), (4), and (6) of table. The estimated coefficient on the G index, for example, is -0.0129 with a t-statistic of -5.51. We conclude that anti-takeover provisions are a strong statistical determinant of idiosyncratic volatility.

To add confidence that the ATI index actually does track restrictions on merger activity, we note that, for our sample, the index is empirically predictive of repressed activity in the market for corporate control. In a probit regression, we have used ATI, $G - ATI$, and the log of firm size (assets), in various combinations, to predict whether a firm will receive a control offer within a two year period.[18] We take data on offers from the SDC Platinum M&A database for the 1990-2001 sample period. ATI is always significant in the results, whatever the exact set of regressors, and only ATI is significant (p-value of 0.035) in the probit with all three regressors. Companies with large values of ATI are less likely to receive a control offer. Thus, ATI is, in fact, an index of anti-takeover intensity in our sample.

In this section, we have shown a strong negative connection between anti-takeover measures and idiosyncratic volatility. Durnev et al. (2004) have recently argued that idiosyncratic volatility is an index of information intensity in general, and in particular an index of the extent to which private information is revealed by trading. This interpretation is consistent with older studies by French and Roll (1986) and Roll (1988). In the alternative, idiosyncratic risk might be largely noise, though the alternative interpretation is less plausible given their systematic results. Within that interpretation, our findings indicate that a governance stance that includes openness to control offers results in more information flowing to market via trading on private information. In the following sections, we provide time series evidence that helps rule out reverse-causality, and tie these findings more directly to private information flow and arbitrage-active institutional trading.

[18] We include $G - ATI$ as a regressor, rather than simply G, because ATI is an additive component of G. We obtain similar results if we use G instead.

III. Substantiating and Interpreting the Relation Between Governance and Idiosyncratic Volatility

We discuss the results in several different respects, from changes in idiosyncratic volatility around governance events to measures of private information flow and finally institutional trading-link hypothesis.

A. Change in Idiosyncratic Volatility Following Governance Events

Endogeneity can be a serious concern with panel evidence of the type presented above. To address this concern, regression tests in this section focus on the changes in idiosyncratic volatility around our governance events, i.e., changes in the G index. Specifically, we compare idiosyncratic volatility before and after takeover restriction adoption or dropping for a given firm. If we are simply documenting self-selection, we should not find a significant change in idiosyncratic volatility around a change in G.

We consider one-year, two-year and three-year event windows (centered at the beginning of the event year), with monthly data, following a change in the governance index G. We include only observations during the window just before and just after a G change. For example, using a one-year window, we compare idiosyncratic volatility in the one-year period before the change in G with the one-year period after the change in G. Because the changes in G coincide with the releases of the G index, they can only happen in 1990, 1993, 1995, 1998, and 2000. Thus, the one-year window does not have overlapping events by construction. For 2-year or 3-year windows there may be overlapping events. In these cases, we exclude overlapping subsequent events. Results remain the same when we include overlapping events.

To register whether there has been a change in idiosyncratic volatility, we construct a post-dummy variable regressor which is one for the years after the G change, and zero for the years before. We denote the dummy variable $I_{\Delta G^+}$ for increases in G and $I_{\Delta G^-}$ for decreases in G. For example, the coefficient on the dummy variable $I_{\Delta G^+}$ gives an estimate of the difference in idiosyncratic volatility between the period following the adoption of anti-takeover provisions and the prior period. We also consider a variable that equals the change in G for the years after the firm enacts a G change, and zero for the years before. We denote this variable ΔG^{+} for increases in G and ΔG^{-} for decreases in G (for easier interpretation we take the absolute value of the negative changes in G). Hence, in contrast with the dummy variable that simply registers whether a change has occurred, this variable considers the actual magnitude of the G change.

There are more events of firms dropping than adopting takeover restrictions during our sample period (1990-2001). Using a one-year window, there are 1709 G-decrease events, while there are only 543 G-increase events. The vast majority of events correspond to a one-point change in G. More precisely, the number of one-point G changes corresponds to 72% of the total number of events in which G increases and 85% of the total number of events in which G decreases.

Panel A of Table III presents the results of estimations examining changes in idiosyncratic volatility by OLS, and alternatively using 2-digit SIC industry and year fixed effects.[19] OLS estimation using the dummy variable correspond to a difference-in-means test of logistic transformed relative idiosyncratic volatility Ψ between the periods immediately before and after the change in G. The results conform to the expectation that idiosyncratic volatility significantly decreases following adoption of anti-takeover provisions. For example, using a one-year window and OLS estimation [see column (1)], we find that idiosyncratic volatility drops 7.16 percentage points for an increase in G (adoption of anti-takeover provision) with a t-statistic of -2.77. Results are consistent across event windows, ranging from -0.0716 to -0.0892 for an increase in G in the OLS models. Results including industry and year fixed effects also confirm that idiosyncratic volatility decreases following the adoption of takeover restrictions, as do results using the variable that takes into account the magnitude of the G change.

[19] Results (not tabulated here) including firm-level controls are similar to the results including industry and year fixed effects.

Additionally, there is strong evidence that idiosyncratic volatility increases after the G index declines. We find that relative idiosyncratic volatility increases 20.14 percentage points for a one-point decrease in G with a t-statistic of 4.53 when using a one-year window around the event (decrease in G). This is true for all event windows with the coefficient ranging from 0.2014 to 0.3110 for a one-point decrease in G using OLS. Results including industry and year fixed effects or considering the magnitude of the G change confirm that idiosyncratic volatility increases when a firm drops of takeover restrictions.

Overall, we find that indeed idiosyncratic volatility is lower when a firm adopts takeover restrictions and higher when a firm drops takeover restrictions. The event study evidence indicates that our panel-based results are not likely driven by reverse-causality, i.e., firms with lower (higher) idiosyncratic risk are more likely to display (avoid) takeover defenses.

Panel B of Table III extends our event study evidence on governance changes by considering as event takeover situations instead of the changes in G. If in fact the negative relation between governance and idiosyncratic volatility arises of speculative trading due to the possibility of a takeover, we should find a stronger negative relationship around takeovers. We take data on offers from the SDC Platinum M&A database in which a sample firm with governance index is a target for the 1990-2001 period. We regress idiosyncratic volatility on a post-dummy variable ($EVENT$) that equals one in the months just before the takeover announcement and zero before this period. We also include an interaction between the $EVENT$ variable and the G index that capture the extent to which takeover differently impact idiosyncratic volatility according to the openness of the firm to takeovers.

We consider six-month, 12-month, and 18-month event windows. We exclude overlapping subsequent takeovers. In any case, the results remain the same when we include overlapping events. For example, using a six-month window, we compare idiosyncratic volatility in the six-month period just before the takeover announcement month with the immediately prior six-month period. We estimate the regressions alternatively by OLS or including 2-digit SIC industry and year fixed effects. Consistent with intuition, the results reveal an increase in idiosyncratic volatility during the takeover period in comparison to the prior period: the $EVENT$ dummy variable coefficient is positive and significant. The finding is consistent across windows and the coefficient ranges from 0.3274 to 0.4088 when including industry and year fixed effects. Furthermore, the regression coefficient on the interaction variable ($G \times EVENT$) is negative and significant supporting the hypothesis that the increase in idiosyncratic volatility in the takeover period mainly accrues for low G firms, i.e., firms that are open to the market for corporate control.

B. Governance and Private Information Trading

In this section, we study the cross-sectional determinants of trading activity and several targeted measures of information flow, focusing on the role of corporate governance provisions. We think of turnover as an alternative to idiosyncratic volatility in proxying for the intensity of private information flowing to a stock's market. The results support the proposition that governance is a driver of information flow. Trading is theoretically linked to the quality or extent of private information [e.g., Kyle (1985) and Blume, Easley, and O'Hara (1994)], and is thus a natural measure of private information flow.[20] A great deal of empirical research has linked trading volume (or turnover) to the intensity of price changes [e.g., Hasbrouck (1991)]. Recent research has provided several targeted information flow indexes (in particular, PIN and $PRIVATE$, both discussed earlier), which we also investigate in this section.

We begin with turnover, measured as unsigned trading activity. We calculate a firm-level monthly turnover ($TURN$) series by dividing monthly share volume by the number of shares outstanding.

[20] An alternative interpretation of trading activity is that it reflects disagreement [Verrecchia (1993) has a cogent discussion] or liquidity trading [as in Campbell, Grossman, and Wang (1993)]. In our context, this is similar to the notion that volatility reflects noise rather than information.

We estimate the following regression equation:

$$\begin{aligned} INFO_{it} \;=\; & c_0 + c_1 GOV_{i,t-1} + c_2 ROE_{i,t-1} + c_3 VROE_{i,t-1} + c_4 LEV_{i,t-1} \\ & + c_5 M/B_{i,t-1} + c_6 SIZE_{i,t-1} + c_7 DD_{i,t-1} + c_8 AGE_{i,t-1} \\ & + c_9 DIVER_{i,t-1} + \epsilon_{it}, \end{aligned} \tag{8}$$

where $GOV_{i,t-1} = \{G_{i,t-1} \vee GD_{i,t-1}\}$, and other regressors the same as in equation (7) for idiosyncratic volatility. $INFO_{it}$ refers to one of the measures just discussed: $TURN$, PIN or $PRIVATE$. $TURN$ is measured monthly and so t is a monthly time index; PIN and $PRIVATE$ are measured annually and so t refers to an annual index. We choose to use the same controls because our goal is not to fully explore the cross section of trading activity, but to control for influences on the extent of private-information trading. Nevertheless, our control regressors cover several categories of potential cross-sectional determinants of trading activity. We use the market-to-book ratio M/B as one proxy for a stock's visibility. Equity market value $SIZE$ is also used to capture a stock's visibility [Lo and Wang (2000)]. We also consider a proxy for information asymmetry and differences of opinion. A firm with excessive debt is considered riskier to investors than an equity-financed firm due to a high probability of financial distress and default as well as agency costs. Differences of opinion are potentially larger for more highly levered firms, and these differences could in turn influence trading activity. Thus, we include leverage LEV as an explanatory variable.

Columns (1) and (2) of Table IV report results for the turnover regressions. The coefficient on the takeover restrictions index G in column (1) is negative and significant. We use the alternative dummy variable version GD in column (2), and also report a negative and significant coefficient. Thus, the evidence is that trading activity is higher in stocks of firms that are open to control offers. Coefficients on control variables are mostly consistent with expectations.

While turnover provides a first test of the extent of information-based trading, it does not directly address the extent of private information trading. The annual probability of information-based trading (PIN) of Easley et al. (2002) and the annual amount of private information trading ($PRIVATE$) of Llorente et al. (2002) do so. Columns (3)-(6) of Table IV present estimates of equation (8) for each of these dependent variables. PIN is negatively related with the governance index, which supports our hypothesis that open firms are more subject to private information trading. The coefficient of the takeover restrictions index G in column (3) is negative and significant. We use the alternative dummy variable version GD in column (4), and also report a negative and significant coefficient. $PRIVATE$ is also negatively related with the governance index. The coefficient of the takeover restrictions index G in column (5) is negative and significant, and similarly with the alternative dummy variable version GD in column (6).

Thus, the evidence supports that trading activity and informed trading is higher in stocks of firms that are open to control offers. The results in this section buttress our earlier findings, and support the proposition that private information flow is facilitated by a takeover-open governance stance.

C. Governance and Earnings Information in Stock Prices

We next test whether in fact firms with fewer anti-takeover provisions have stock prices that contain more information about future earnings. We find that they do, consistent the informational interpretation of our earlier results.

To do so, we estimate the following regression equations for our panel of annual observations at the two-digit SIC industry level:

$$\begin{aligned} FER_{kt} = \; & c_0 + c_1 G_{k,t-1} + c_2 ROE_{k,t-1} + c_3 VROE_{k,t-1} + c_4 LEV_{k,t-1} \\ & + c_5 M/B_{k,t-1} + c_6 SIZE_{k,t-1} + c_7 DD_{k,t-1} + c_8 AGE_{k,t-1} \\ & + c_9 DIVER_{k,t-1} + \epsilon_{kt}, \end{aligned} \tag{9}$$

where $FER_{kt} = \{FERC_{kt} \vee FINC_{kt}\}$ for industry k in year t, following Durnev et al. (2003), and other regressors are two-digit SIC industry averages of the same regressors in equation (7) for idiosyncratic volatility. The dependent variables here measure the amount of information about future earnings which is incorporate in current stock returns. $FERC$ is the annual future earnings response coefficient and $FINC$ is the annual future earnings incremental explanatory power, as described earlier.

Columns (7) and (8) of Table IV display regressions of industry average future earnings response measures ($FERC$ and $FINC$) on two-digit SIC industry average idiosyncratic volatility and control variables. Regression includes one-digit SIC industry fixed-effects, following Durnev et al. (2003). We find that the governance index G is negatively associated with future earnings response measures. The G coefficients is -2.1178 on $FERC$ regression and -0.1020 on $FINC$ regression and both coefficients are statistically significant at conventional levels.

D. Governance and Institutional Trading: The Trade-link Hypothesis

Institutional trading is an important channel through which information is incorporate into stock prices. Piotroski and Roulstone (2004) find that institutional trading is positively associated with idiosyncratic volatility. Also, Jiambalvo, Rajgopal, and Venkatachalam (2002) and Hartzell and Starks (2003) find that institutional investors contribute to private information collection and trading.[21]

We introduce institutional trading as an additional control in our basic model in equation (7). At least, this serves the purpose of checking the robustness of the relation between

[21] Other channels through which information is incorporate into stock prices include analysts and insiders activities. Evidence on the role of analysts is mixed and there is recent evidence that relative idiosyncratic volatility is negatively related with analyst coverage [Piotroski and Roulstone (2004) and Chan and Hameed (2005)]. Evidence on the role of insider trading is provided in Piotroski and Roulstone (2004). They find that insider trading is positively related to relative idiosyncratic volatility. However, Gompers et al. (2003a) present evidence that insider trading is uncorrelated with the governance index.

governance and idiosyncratic volatility after controlling for the level of institutional trading. Additionally, we test our trading-link hypothesis by including an interaction variable between governance and institutional trading. If in fact institutional trading contributes to the incorporation of information into stock prices of firms open to takeovers, we expect to find a negative coefficient on this interaction variable.

Specifically, we estimate the following idiosyncratic volatility monthly regression equation:

$$\begin{aligned}\Psi_{it} \quad = \quad & c_0 + c_1 G_{i,t-1} + c_2 ROE_{i,t-1} + c_3 VROE_{i,t-1} + c_4 LEV_{i,t-1} + c_5 M/B_{i,t-1} \\ & + c_6 SIZE_{i,t-1} + c_7 DD_{i,t-1} + c_8 AGE_{i,t-1} + c_9 DIVER_{i,t-1} \\ & + c_{10} INS_{i,t-1} + c_{11} INS_{i,t-1} \times G_{i,t-1} + \epsilon_{it},\end{aligned} \tag{10}$$

where $INS_{i,t-1} = \{INST_{i,t-1} \vee INSTA_{i,t-1}\}$, and other variables are as previously defined. $INST$ is the quarterly average of the absolute change in the number of shares held by institutions, as a fraction of annual trading volume. $INSTA$ is the quarterly average of the absolute change in the number of shares held by arbitrage-active institutions, as a fraction of annual trading volume, as defined in a previous section.

Columns (1) and (2) of Table V reports estimates of equation (10) using $INST$, the broad measure of institutional trading, respectively with and without an interaction regressor defined as $INST \times G$. The estimate of the governance coefficient is strongly significantly negative in both cases ($c_1 < 0$). Institutional trading is associated with more idiosyncratic volatility ($c_{10} > 0$). Note that the interaction regressor exerts a significantly negative effect when present ($c_{11} < 0$). Institutional trading adds to the statistical effect of governance on volatility, in that institutional trading accelerates the incorporation of firm-specific information into stock prices and, consequently increases idiosyncratic volatility.

Columns (3) and (4) of Table V report analogous results, but using $INSTA$, the targeted measure of arbitrage-active institutional trading defined earlier. The relation between governance and volatility remains strong after controlling for this focused type of institutional trading.

The results also support our trading-link hypothesis, in that the governance-idiosyncratic risk relation is stronger in the presence of high levels of institutional trading at the firm-level. The coefficient on the interaction variable is -0.0595 with a t-statistic of -2.30. The point is that governance and trading by institutions, especially by arbitrage-active institutions, is associated with an incrementally greater absolute influence of governance on idiosyncratic risk.

At a broad level, the trading-link nature of a negative relation between anti-takeover provisions and idiosyncratic risk comports with three sets of recent results. First, institutional investors are active in private information collection and trading, according to evidence in Chakravarty (2001), Jiambalvo et al. (2002) and Hartzell and Starks (2003). Piotroski and Roulstone (2004) specifically show that institutional trading is positively associated with idiosyncratic volatility. Second, there is a known connection between takeover defenses and institutions' decisions [Bethel, Liebeskind, and Opler (1998)], and between such decisions and the value effects of takeover defenses [Agrawal and Mandelker (1990)]. Third, arbitrage-oriented institutions may play a special role. Bushee and Noe (2000), for example, document a link from disclosure quality (a different aspect of information flow) to ownership by "transient" institutions and then to total volatility.[22] Hsieh and Walkling (2005) find that arbitrageur holdings are predictors of successful bids as well as of additional bids in takeover situations, which is evidence of private information collection around takeovers.[23] Butz (1994) emphasizes that even modest stakes by activist investors can impart significant

[22] In view of this work, subsequent tests in the robustness section control for the transparency of firms' financial reporting.

[23] This trading-link might become even more prominent in the future, if press claims that hedge funds are increasingly interested in speculating on possible mergers prove substantial. See, for example, Singer (2005). Also, anecdotal evidence suggests that increasing institutional ownership, particularly by hedge funds, results in more request for information from companies, which suggests that more information gathering is occurring ["Chaos in the Public Square" (2005)].

control, provided that the threat of a takeover is not completely empty and that information about firm value flows fairly unimpeded to the market.

Our paper implies that these conditions are mutually-reinforcing, and therefore perhaps more powerful. The stronger a firm's takeover defenses, the more impeded is the flow of information. And the effect is accentuated when arbitrage-oriented institutions avoid the stock.

IV. Robustness

We show that our primary findings are robust to controls for accounting transparency, idiosyncratic volatility measures, and several aspects of our empirical methodology.

A. Controlling for Accounting Transparency

Accounting disclosure is a central element of the information flow. More transparent or precise disclosures are usually thought to reduce the cost of acquiring private information [see, for example, the discussion in Verrecchia (1999)], but may also reduce the benefit [see, for example, Barth, Clinch, and Shibano (1999)]. Whether more transparency encourages the collection of private information or crowds it out depends on the balance of effects on these benefits and costs. In our context, if intense trading on private information underlies idiosyncratic volatility, then we might observe less volatility for stocks for which public information is more important, since more information is now flowing via lower-frequency accounting releases, or more volatility as additional information collection is encouraged [as in Kim and Verrecchia (1991b)].

This reasoning leads to testable hypotheses. First, if the governance index G is also indicative of the general level of corporate openness, it might not be statistically important for information flow once transparency level is directly considered. Therefore, we test whether G retains its negative coefficient in a idiosyncratic volatility regression once a measure of transparency is also included. Second, if, as we hypothesize, the idiosyncratic volatility measure Ψ reflects incentives for private information collection, then the coefficient on transparency in such a regression should be significant. Specifically, the presence of extensive transparency, controlling for the level of takeover openness, might lead to less information flow than otherwise (a crowding out effect) or more (an encouragement effect).

Our main result in this subsection is that governance remains a significant determinant of volatility, as in the previous subsection, even controlling for transparency. Firms with low levels of transparency also display lower levels of idiosyncratic risk, controlling for the level of governance openness - consistent with an encouragement effect.

We now describe our results in more detail. To measure transparency, we follow the accounting literature [e.g., Francis et al. (2005)], and focus on two measures based on the absolute size of abnormal accruals. Intuitively, larger accruals relative to what would be expected given a firm's activities are considered to be inverse indicators of accounting transparency. The measures of transparency we use are $EQ2$ and $EQ5$ described earlier. Note that all these measures are *inverse* indexes of transparency, in that they increase in the magnitude of unexpected accruals.

In Table VI we present idiosyncratic volatility regression results controlling for accounting transparency. The table is set up similarity to earlier tables, and reports variations on the same basic model in equation (7). The key difference is that the accruals-based measure of accounting quality EQ, where $EQ = \{EQ2 \vee EQ5\}$, is included as a regressor to explain relative idiosyncratic volatility Ψ. $EQ2$ is employed in columns (1) and (2), while $EQ5$ is used in columns (3) and (4). For all variations of the basic model that we estimate, our fundamental result is unchanged. The governance index is strongly correlated with idiosyncratic volatility, even after controlling for the transparency of accounting information.

Estimates of the coefficient on EQ, an inverse measure of accounting transparency, are negative and significant in every column [except in column (4)]. That is, the level of idiosyncratic volatility is greater in the presence of extensive transparency. Within the interpretation of idiosyncratic risk as private information flow, this is indicative of more information flowing to market via trading when accounting numbers are more transparent. This evidence is consistent with theoretical suggestions that high-quality disclosure could encourage the collection of private information, leading to more idiosyncratic volatility [Kim and Verrecchia (1991b)]. Interpreted within our paradigm, less accounting information apparently disproportionately inhibits efforts to collect more private information.

Our results in this section are related to recent work by Jin and Myers (2005), who develop a theory linking management opportunism, transparency, and idiosyncratic volatility. They provide cross-country evidence that low transparency level and poor investor protection result in low levels of relative idiosyncratic volatility (high R^2).

Our core results complement Jin and Myers (2005) work by showing that a particular form of firm-level governance - anti-takeover provisions - is associated with the incorporation of firm-specific information into stock prices. Fewer takeover restrictions apparently signal a governance setup and a company that is more open in general, facilitating the collection of private information. One plausible mechanism is that strong investor protection, expressed by openness to takeovers, is associated with a lower possibility of insiders (controlling shareholders and managers) expropriating arbitrageurs and outside investors. For such companies, ownership by outside investors is therefore encouraged. More such investors means more noise trading, providing cover and profits for those willing to invest in private information collection. Thus, fewer anti-takeover provisions can promote private information collection and trading by outside investors.

Additionally, our results on transparency support some of Jin and Myers (2005) predictions, showing that accounting transparency plays an important role in determining idiosyncratic volatility. Our firm-level results are confirmatory of their country-level work: we find evidence that low transparency is associated with low levels of idiosyncratic volatility, as well as evidence that poor firm-level corporate governance is associated with low levels of idiosyncratic volatility.

B. Additional Robustness Checks

In our panel regression setup, we have mitigated some potential for biases by using predetermined measures of the governance index. Nonetheless, the index measures governance characteristics related to takeover protections only with error. Therefore, the regressions in earlier tables have the potential of biases in t-statistics due to errors-in-variables and serial correlation. Considering this, we now confirm our results using Fama and MacBeth (1973) regression approach. Specifically, we estimate cross-sectional regressions for each month and compute the reported coefficients and t-statistics from the set of estimates for all months within a time period. Results are reported for full the sample period and subperiods (corresponding to the G index releases).

Table VII reports Fama and MacBeth (1973) regression estimates of our basic model in equation (7) using the takeover restriction index G a regressor and relative idiosyncratic volatility Ψ as dependent variable, and alternatively using the dummy variable GD for closed or open governance structure as a regressor. For each subperiod, we present results where, alternatively, regressors are: (1) governance only, and (2) governance and control variables. Similarly, the coefficient on the anti-takeover index G is negative in the full sample period and in almost all subperiods, and is nearly always significant. The coefficient on GD is negative and significant in the full sample and in almost all sub-periods. Thus, results in Table VII support our belief that idiosyncratic volatility is decreasing in the level of takeover protections. Overall, the Fama and MacBeth (1973) regressions are strongly confirmatory of our earlier findings.

Table VIII presents robustness evidence that our results are not due to the particular model of returns used to estimated idiosyncratic volatility, to the governance openness formulations used, or estimation methodology used. Results in columns (1)-(4) are estimated using a different version of the dependent variable in our basic panel model in equation (7), alternatively Fama and French (1992) three-factor model residual variance and industry factor model residual variance.

Columns (5)-(8) check the robustness of our results to different estimators of idiosyncratic volatility. Columns (5)-(6) use the French, Schwert, and Stambaugh (1987) estimator for variance:

$$\widehat{\sigma}_{it}^2 = \sum_{d=1}^{T} r_{id}^2 + 2\sum_{d=2}^{T} r_{id} r_{i,d-1}, \tag{11}$$

and the Scholes and Williams (1977) estimator for covariance:

$$\widehat{\sigma}_{ijt} = \sum_{d=1}^{T} r_{id} r_{jd} + \sum_{d=2}^{T} \left(r_{i,d-1} r_{jd} + r_{id} r_{j,d-1} \right), \tag{12}$$

where as before d indexes the days of month t. The additional terms adjust for biases that result from autocorrelation and cross-autocorrelations of daily returns. Columns (7) and (8) consider quarterly idiosyncratic volatility estimates instead of monthly. This potentially increases the precision of the estimates but reduces the number of data points. Moreover, it allows that the dependent variable is constructed with a frequency consistent with the majority of our control variables (e.g., accounting variables).

Columns (9) and (10) considers a variation of the measure of governance openness, only using the earliest observation (1990) present in the dataset to further guard against endogeneity concerns raised earlier. This also ensures that our results are not driven multiple observations on the same firms.

Columns (11) and (12) estimates the panel regression by including year fixed effects. This allows to control for time variation and trend in the series of idiosyncratic volatility.

Columns (13) consider the differences-in-differences approach by including both year and firm fixed effects [Bertrand and Mullainathan (2003)]. This serves as additional robustness check for endogeneity.

In all models, the coefficient on a predetermined measure of takeover restrictions remains negative and strongly significant. Our basic result is confirmed: more anti-takeover governance provisions are strongly associated with less idiosyncratic volatility.

V. Governance and the Market Valuation of Investment Decisions

This section extends our study of governance and idiosyncratic risk to consider their possible influence on investors' perceptions of the quality of corporate investment decision making. In a cross-sectional study, Durnev et al. (2004) find that industries with large idiosyncratic volatilities tend to exhibit marginal Tobin's q (hereinafter, $\dot{q}$) levels closer to the value of unity that they should possess in a full-information value-maximizing frictionless economy. Durnev et al. (2004) interpret $|\dot{q} - 1|$ as "measuring investors' aggregated opinions about corporate investment efficiency," emphasizing an informational view of $\dot{q}$. In this context, a negative relationship of relative idiosyncratic volatility, Ψ, a measure of firm-specific information intensity, to $|\dot{q} - 1|$ is natural. Durnev et al. (2004) also state "$\dot{q} > 1$ implies underinvestment and $\dot{q} < 1$ implies overinvestment," implying a judgment on the extent to which management is maximizing value. Anti-takeover protections, such as those measured by the G index, are thought to free a management to overinvest [Jensen (1986), Lambrecht and Myers (2005), and Masulis, Wange, and Xie (2005)], so it makes sense to consider how governance relates to such judgments. Frictions that prevent immediate full investment in strong opportunities would induce $\dot{q} > 1$; such frictions are likely to vary according to the nature of real investment and business conditions, and can thus be proxied by industry and control variables.

Prior literature on over-investment and free cash flow suggests that there is a link between governance and corporate investment. Chirinko and Schaller (2004), for example, show that firms with extensive financial slack overinvest by 7 to 22% of their asset base. Whether governance characteristics in particular are a strong driver of those incentives is yet at issue. Gompers et al. (2003a) and Larcker et al. (2004) find at least some evidence of a direct relation of anti-takeover measures and overinvestment. Gompers, Ishii, and Metrick (2003b) find that firms whose dual-class structures allow for more entrenchment overinvest more intensely. These studies do not consider a volatility link, nor do they consider the possibility of underinvestment.[24]

We have shown in previous sections that corporate governance is a direct determiner of idiosyncratic volatility. This raises a question as to whether the informational component of volatility is directly connected to corporate investing quality, or whether volatility is just a stand-in for the effect of corporate governance arrangements in Durnev et al. (2004) results. In this section, we present tests to sort out these relationships.

We establish two main results. First, idiosyncratic volatility is a driver of marginal Tobin's q, controlling for any influence of anti-takeover governance provisions. This result strengthens Durnev et al. (2004) conclusions, and adds credence to the information flow interpretation of idiosyncratic volatility. Second, the statistical relation of takeover protections to idiosyncratic volatility actually offsets part of the volatility-$\dot{q}$ relation. Takeover protections are positively correlated with $\dot{q}$ nearer one for firms that underinvest ($\dot{q} > 1$). Moreover, the component of idiosyncratic volatility that is negatively correlated with G is *also* positively related to $|\dot{q} - 1|$ for firms with $\dot{q} > 1$. Our evidence is consistent with the following economic interpretation. A limited flow of information allows for both corporate overinvestment and underinvestment, from the perspective of an investor, in that value is delinked from corporate decisions. At the same time (although perhaps not for the same firms), takeover protections that discourage the flow of information also encourage managerial spending, and actually militate underinvestment for some firms.

A. Industry-level Results

To begin, we require a measure of marginal Tobin's q. We follow the same general approach as Durnev et al. (2004). To obtain estimates useful in our panel data setting, we estimate $\dot{q}$ for each two-digit SIC industry, rather than using three-digit industries as in Durnev et al. (2004). This provides for more data for each industry in each year, at the cost of a richer

[24] Our earlier result on a governance-volatility relation together with the results of Durnev et al. (2004) on a volatility-investing quality relation suggest that the referenced results may be subject to bias, in that shareholder-oriented governance can operate through volatility rather than directly.

cross-section. To show this setup does not drive our results, we demonstrate that the primary findings also hold using a firm-specific $\dot{q}$. Durnev et al. (2004) use a recursive procedure to estimate replacement value aspects of their marginal q's, akin to that in Lewellen and Badrinath (1997). This is not practical in a panel setting, so we make more use of book values; however, since we are estimating marginal, not average, q's, and since we use values from cash flow statements whenever possible, the estimates should not be much affected as long as replacement values of assets-in-place do not change too much over a year's time. Full details on the procedure are provided in the Appendix.

In a nutshell, our $\dot{q}$ is based on estimates of the following regression for each two-digit SIC industry for each subsample of firms i in each year t:

$$\Delta EV_{it} = \lambda_0 + \dot{q}\Delta NFA_{it} + \lambda_1 D_{it} + \lambda_2 EV_{i,t-1} + u_{it}, \tag{13}$$

where ΔEV_{it} is firms i change in enterprise value in year t, ΔNFA_{it} is the change in net fixed assets, D_{it} is the flow of cash disbursements to all investors, and u_{it} is a residual term. All variables are scaled by the previous year's level of net fixed assets. Estimates of the coefficient $\dot{q}$ measure marginal Tobin's q at the industry level because they register the market value reaction to management's decisions on asset changes, controlling for extraneous value-change factors. Using each industry/year estimates of $\dot{q}$, we form a dataset on $|\dot{q}-1|$. This quantity is the subject of analysis by estimating the following annually regression equation:

$$|\dot{q}-1|_{kt} = \alpha_0+\alpha_1\Psi_{k,t-1}+\alpha_2 G_{k,t-1}+\alpha_3 LEV_{k,t-1}+\alpha_4 M/B_{k,t-1}+\alpha_5 SIZE_{k,t-1}+\alpha_6 DIVER_{k,t-1}+\epsilon_{kt}, \tag{14}$$

where regressors are two-digit SIC industry averages. Panel A of Table IX presents the results of this panel regressions where $|\dot{q}-1|$ is the dependent variable.[25]

[25] We have checked that similar results hold using the log absolute deviation as the dependent variable, which should better conform to OLS regression assumptions, though would not be as directly comparable to Durnev et al. (2004) results. Our results are also robust to the inclusion of one-digit SIC industry and year fixed-effects.

The results show that $|\dot{q}-1|$ is negatively related to relative idiosyncratic volatility Ψ [column (1)], consistent with Durnev et al. (2004) result. Strikingly, the governance index G is also negatively correlated with $|\dot{q}-1|$, whether or not idiosyncratic volatility is also included in the regression [columns (2) and (3)]. This seems to suggest that takeover protections improve capital budgeting quality, driving $\dot{q}$ toward unity.

Columns (4) and (5) report tobit regressions in which the dependent variable is $\dot{q}$ censored from below [column (4)] or from above [column (5)]. The results indicate that the positive relation between takeover protections and corporate investment efficiency is due to firms with $\dot{q} > 1$, i.e., firms which might be underinvesting. With censoring from below (above), unbiased regression coefficients are estimated based on the variation in $\dot{q}$ where $\dot{q} > 1$ ($\dot{q} < 1$). The negative relation of $|\dot{q}-1|$ to G is evident only for high-$\dot{q}$ observations. The implication is that anti-takeover protections are associated with a smaller (but still positive) degree of underinvestment.

The importance of idiosyncratic volatility in reducing $|\dot{q}-1|$ is therefore partially offset by the influence of the governance index G, which is also correlated with $|\dot{q}-1|$. To confirm that the offset is only partial, and that G and Ψ do in fact have their own independent effects, we present a final regression in column (6). In this regression, we follow the procedure used in Almeida and Wolfenzon (2005) for separating endogenous and exogenous effects. In our case, these are the effects of the relative idiosyncratic volatility Ψ on $|\dot{q}-1|$. We first regress Ψ on the governance index G, and retain the predicted values and residuals. The residuals contain information on Ψ that is statistically unrelated to G. We then regress $|\dot{q}-1|$ on both the predicted values and residuals, along with other control variables. The resulting coefficient on predicted value is significantly positive, and the coefficient on the residual is significantly negative. The predicted value is simply an inverse scaled version of G given that G and Ψ are negatively correlated, so the positive coefficient indicates a direct effect of the governance index on $|\dot{q}-1|$. We have confirmed via a Tobit, as above, that this effect is driven by the high-$\dot{q}$ observations, consistent with the idea that underinvestment is limited.

The negative residual coefficient is consistent with the Durnev et al. (2004) reasoning for the independent importance of idiosyncratic volatility.

B. Firm-level Results

As Durnev et al. (2004) argue, it is difficult to obtain precise estimates of marginal q for individual firms. Nonetheless, we can confirm that our results are robust to firm-level estimation. This is particularly useful because we then can employ industry-level indicator variables in a firm-level regression to control for the influence of the nature of an industry's investment and returns process. To accomplish this, we estimate industry-level $\dot{q}$'s in a single pooled version of the Durnev et al. (2004) equation, and investigate how governance and idiosyncratic volatility are related to firms' deviations from their industry norm by placing structure on the residual term. The regression setup is

$$\begin{aligned} \Delta EV_{it} &= \sum_{k=1}^{K} \lambda_0^k I_{it}^k + \dot{q}^{overall} \Delta NFA_{it} + \sum_{k=1}^{K} \dot{q}^k I_{it}^k \Delta NFA_{it} \\ &\quad + \lambda_1 D_{it} + \lambda_2 EV_{i,t-1} + u_{it}, \end{aligned} \tag{15}$$

and

$$\ln(u_{it}^2) = \alpha_0 + \alpha_1 \Psi_{it} + \alpha_2 G_{it} + \alpha_3 LEV_{it} + \alpha_4 M/B_{it} + \alpha_5 SIZE_{it} + \alpha_6 DIVER_{it} + \varepsilon_{it} \tag{16}$$

where k indexes industries, I_{it}^k refers to an indicator variable for firm i membership in the k-th industry in year t, $\dot{q}^k$ captures the industry-specific deviation from the full-sample estimate of $\dot{q}^{overall}$, and all other variables are as previously defined. In this system, we posit that the squared error term captures the tendency of any firm to deviate from its industry level of marginal q, and that this tendency may be related to the firm's governance index, relative idiosyncratic volatility, and other control variables. Alternatively, we also model the squared residual as a function of the components of share (i.e., G-related and independent, as we also did in the industry-level analysis).

Panel B of Table IX presents OLS estimates of this system, along with robust t-statistics. To save space, we suppress coefficients on industry dummies and industry interaction terms. Column (1) presents estimates of the equation relating market value changes (ΔEV) to management choices about assets (ΔNFA), which is the basis for marginal Tobin's q. Note especially that the estimate of $\dot{q}^{overall}$ is near unity (1.325) and it is very strongly statistically significant. Thus, the null hypothesis that $\dot{q}^{overall} = 1.0$ is not rejected by the data.

Columns (2) and (3) contain estimates of the determinants of the squared residual. In this specification, regressors associated with residuals nearer to zero will generate negative coefficient estimates. Residuals nearer to zero are economically interpreted as indications that the firm-specific marginal q is near to the industry marginal q. The economic implications of these regressions, which focus on intra-industry effects, are consistent with those in the previous subsection focusing on cross-industry effects.

In column (2), both the takeover protections index G and relative idiosyncratic volatility Ψ have negative coefficients - extensive takeover protections and high volatility both militate toward a firm $\dot{q}$ that is like the industry's. This is consistent with the economic interpretation in the previous section, that extensive takeover protections are associated with reduced underinvestment.

In column (3), the key regressors reflect the decomposition of idiosyncratic volatility, similar to those used in the previous section. To compute these results, we again follow Almeida and Wolfenzon (2005), and first decompose relative idiosyncratic volatility Ψ into a component linearly related to the governance index G and a residual component. As in the previous section, we refer to these as the fitted value and the error value of Ψ. We find that the error value portion of Ψ is associated with an attenuated level of firm-specific $\dot{q}$ deviation from the industry level, and that the G-related fitted value portion is associated with larger deviation. This results is consistent with the findings in the industry level section.[26]

The results of this firm-specific investigation extend the conclusions of the industry investigation. Table IX establishes that, in fact, there is a fundamental relation between the market's perception of investing quality and idiosyncratic volatility, even when the relation between volatility and governance is taken into account. Governance stance itself appears to have an effect on this practical business outcome which partially offsets the effect of idiosyncratic volatility in that takeover protections discourage underinvestment. Overall, it is the information flow rather than governance that exerts the dominant influence.

VI. Conclusion

We find that corporate governance characteristics that bear on the likelihood of takeovers are tightly correlated with idiosyncratic risk. Idiosyncratic risk is decreasing in firms' degree of insulation from takeovers. The result is robust to several firm-level controls such as size and accounting transparency, and to a range of empirical methodologies.

Within the interpretation of idiosyncratic risk as an index of information flow, our finding implies a tight link between openness to the market for corporate control and openness of private information flow to the market. Openness to the market for control links to information flow in a way not captured by the openness or opaqueness of the books.

Buttressing an informational interpretation of our core result, we find that takeover vulnerability is similarly related to several alternative measures of private information flow and trading. Additionally, we find that stock prices in industries with less anti-takeover provisions contain more information about future earnings.

We next show that an institutional trading link is one route for the relationship of governance to idiosyncratic risk. Specifically, that relationship persists and is more pronounced for stocks with intense trading by institutional investors, and especially those that have re-

[26] Note that the ceofficients on the control variables are the same across columns (2) and (3) because they correlate in exactly the same way with G and Ψ as they do with the components.

cently been involved in risk arbitrage around takeovers. Thus, at least one of the links from governance to volatility and information flow is via arbitrageur institutions.

Finally, considering the connections among corporate governance characteristics, idiosyncratic risk, and investing quality. The quality of investing is a business outcome that is influenced by the alignment of managerial and shareholder interests, which is influenced by corporate governance characteristics. We find a positive correlation between investing quality and idiosyncratic risk. Upon decomposing volatility into governance-related and non-governance related components, we find that only non-governance related idiosyncratic volatility is associated with the quality of investment decision-making. Our finding means that it is the information flow rather than governance that is more important for this practical business outcome.

A. Appendix: Estimating Marginal Tobin's q

Consider a firm's summarized market value balance sheet.

Economic Assets	**Claims**
Net Working Capital (NWC)	Short term Debt (STD)
+ Net Fixed Assets (NFA)	+ Long term Debt (LTD)
+ Going Concern Value (GCV)	+ Equity (E)
= Total Economic Assets (A)	= Total Capital (TC)

Define Enterprise Value (EV) as $EV \equiv E + LTD + STD - NWC$, and note that $EV = NFA + GCV$. Then, over time, the change (Δ) in enterprise value can be thought of as the sum of changes in its components: $\Delta EV = \Delta NFA + \Delta GCV$, provided that outside investors do not contribute additional capital nor do they withdraw value via dividends, interest, or the like; we add consideration of such complications below.

Conceptually, Tobin's q relates the firm market value to the replacement value of its physical assets (i.e., total net economic assets less going concern value). Marginal Tobin's q (hereinafter $\dot{q}$) is therefore the multiplier that must be applied to the marginal decision on physical assets, ΔNFA, to account for ΔEV, i.e., $\dot{q} \equiv \frac{\Delta EV}{\Delta NFA}$. This definition suggests that $\dot{q}$ could be estimated by regressing the change in enterprise value (a market value concept) on the change in physical assets (a replacement value concept). This is the idea that underlies the Durnev et al. (2004) suggestion for estimating marginal q.

To implement this suggestion, several additional considerations must be factored into the regression. We should control for any expected changes in NFA and EV, including depreciation. We should account for withdrawals of value via dividends and repurchases, and allow for any tax effects of the withdrawals. Durnev et al. (2004) discuss these issues and propose an augmented regression equation [their equation (11)] to estimate $\dot{q}$. In our notation, that regression is

$$\Delta EV_{it} = \lambda_0 + \dot{q}\Delta NFA_{it} + \lambda_1 D_{it} + \lambda_2 EV_{i,t-1} + \varepsilon_{it}, \tag{A1}$$

where ΔEV_{it} is firms i change in enterprise value, ΔNFA_{it} is the change in net fixed assets, D_{it} is the flow of cash disbursements to all investors, and ω_{it} is a residual term.

In the total value of disbursements via dividends and share repurchases; interest payments can be ignored if, following Durnev et al. (2004)'s suggestion, we think of the debt as effectively perpetual. Durnev et al. (2004) suggest scaling all variables by a lagged measure of replacement value, such as the previous year NFA. In this context, the extra regression coefficients register the additional effects noted above. Specifically, λ_0 registers the difference between the expected rate of increase in physical assets and their depreciation rate, λ_1 registers the tax effect on disbursements, and λ_2 registers the expected return on an investment in the firms enterprise value.

Durnev et al. (2004) construct a data set for estimating this regression for each three-digit SIC industry from Compustat data, using recursive procedures of the type suggested in Lewellen and Badrinath (1997) to adjust book values to replacement values. Their goal is to obtain a single estimate of $\dot{q}$ for each three-digit industry for the early-to-mid 1990s. For our purposes, we require panel data on $\dot{q}$ over the 1990s. Furthermore, for comparison to our other results, we prefer some information on $\dot{q}$ at the firm level. With these goals in mind, we modify Durnev et al. (2004) procedures somewhat. The need for a panel in which time-series observations reflect mainly incremental information precludes the extensive use of recursive procedures. The trade-off is that our estimates are then more dependent on book values. We mitigate this issue by relying on data from statements of cash flows wherever possible, since these are inherently at current values, and by factoring in consideration the effect of inventory valuation method. In order to have a broader sample for each industry at each point in time and thereby enhance statistical precision, we estimate $\dot{q}$ for two-digit SIC industries rather than three-digit industries. The trade-off is a greater chance of forcing a single $\dot{q}$ estimate for firms that actually have disparate $\dot{q}$'s. This is one of the reasons for the analysis of residual terms that we report in the main text.

Our regressors for each firm/year are computed using the following Compustat annual data items:

$EV =$ Item #25$\times$ Item #199+ Item #56+ Item #9+ Item #34$-$ Item #4.

$\Delta NFA =$ Item #128+ Item #129 $+$ ΔItem #240$-$ Item #303; if Item #303 is missing then we add ΔItem #3 instead.

$D =$ Item #127+ Item #115.

Each regression variable is scaled by the previous year net fixed assets. We include inventory in the calculation of the change in net fixed assets, even though it is formally a current asset, to reflect that inventory decisions are decisions about assets whose value is judged by investors in the stock market.

We require 20 firms in with complete data in each industry/year in order to estimate equation (A1). Following the general approach of Durnev et al. (2004), we eliminate some firms with extreme data from the estimation and we eliminate some extreme outliers from the results. Specifically, we eliminate firms whose change in enterprise value is more than 300 percent in absolute value, and we eliminate industry/years with $\dot{q}$ estimates of more than 6 in absolute value.

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Table I
Descriptive Statistics

This table reports mean, median, standard deviation, maximum, and minimum of variables. Panel A presents descriptive statistics for monthly idiosyncratic volatility measures. σ_e^2 is the monthly idiosyncratic variance (annualized). σ_e^2/σ_i^2 is the monthly relative idiosyncratic volatility. Ψ is the monthly logistic transformed relative idiosyncratic volatility. Panel B presents descriptive statistics for G, the IRRC-Gompers et al. (2003a) governance index, which is based on 24 anti-takeover provisions, and for ATI, the Cremers and Nair (2004) anti-takeover provisions index, which incorporates only three anti-takeover provisions. Panel C presents descriptive statistics for alternative measures of information flow and informed-based trading. $TURN$ is the monthly share volume divided by shares outstanding. PIN is the annual probability of information-based trading of Easley et al. (2002). $PRIVATE$ is the annual amount of private information trading of Llorente et al. (2002). Panel D presents descriptive statistics for future earnings response variables. $FINC$ is the annual future earnings incremental explanatory power of two-digit SIC industries. $FERC$ is the annual futures response coefficient of two-digit SIC industries. Panel E presents descriptive statistics for the institutional trading activity variables. $INST$ is the quarterly absolute change in the number of shares held by institutions, as a fraction of annual trading volume. $INSTA$ is the quarterly absolute change in the number of shares held by takeover arbitrage institutions, as a fraction of annual trading volume. Panel F presents descriptive statistics for the time series of other control variables. ROE is the quarterly average of return-on-equity (annualized) across all firms. $VROE$ is the sample variance of quarterly ROEs over the last three years. LEV is the quarterly average of leverage across all firms defined as the ratio of long-term debt to total assets. M/B is quarterly average of the log of the market-to-book equity ratio across all firms. $SIZE$ is the monthly average of log market capitalization. DD is the quarterly average of the dividend dummy. AGE is the monthly average of log age across all firms defined as the number of months (divided by 12) since the stock listing. $DIVER$ is the quarterly average of a dummy variable that equals one when a firm operates in multisegments and zero otherwise. $EQ2$ is the annual average of a measure of accounting opaqueness defined as the absolute value of firm-specific residuals from an industry regression of current accruals on (the reciprocal of) assets and revenue growth. $EQ5$ is the annual average of a measure of accounting opaqueness defined as the absolute value of firm-specific residuals from an industry regression of total accruals on lagged, contemporaneous, and lead cash flow from operations. The sample period is from 1990 to 2001. Financial and utilities industries are omitted (SIC 6000-6999 and 4900-4999). All variables are winsorized at the bottom- and top-1%.

		Mean	Median	Std Dev	Maximum	Minimum	N
Panel A: Idiosyncratic Volatility Variables (firm-level)							
Idiosyncratic volatility	σ_e^2	0.194	0.099	0.280	2.471	0.008	161691
Relative idiosyncratic volatility	σ_e^2/σ^2	0.854	0.907	0.155	1.000	0.071	161691
Logistic relative idiosyncratic volatility	Ψ	2.731	2.261	2.198	19.552	-2.574	160456
Panel B: Corporate Governance Variables (firm-level)							
Governance index	G	8.971	9.000	2.833	19.000	1.000	6043
Anti-takeover index	ATI	1.830	2.000	0.944	3.000	0.000	6043
Panel C: Alternative Information Flow Variables (firm-level)							
Turnover	$TURN$	0.101	0.065	0.107	0.771	0.004	159599
Probability of information-based trading	PIN	0.162	0.156	0.053	0.353	0.067	9953
Amount of private information trading	$PRIVATE$	-0.002	0.001	0.098	0.253	-0.274	13662
Panel D: Future Earnings Response Variables (2-digit SIC industries)							
Future earnings response coefficient	$FERC$	1.209	1.048	5.914	18.792	-19.601	160
Future earnings incremental explanatory power	$FINC$	0.373	0.362	0.187	0.838	0.012	160
Panel E: Institutional Trading Activity Variables (firm-level)							
Institutional trading	$INST$	0.133	0.099	0.123	0.810	0.000	45719
Arbitrage institutional trading	$INSTA$	0.093	0.066	0.089	0.536	0.000	44774
Panel F: Control Variables (firm-level)							
Return-on-equity	ROE	0.097	0.123	0.269	1.576	-2.345	50824
Volatility of return-on-equity	$VROE$	0.196	0.002	2.115	55.783	0.000	48720
Leverage	LEV	0.228	0.210	0.162	0.890	0.001	45989
Market-to-book	M/B	0.805	0.753	0.695	3.091	-0.928	49451
Market capitalization	$SIZE$	13.749	13.694	1.515	17.750	9.945	157278
Dividend dummy	DD	0.588	1.000	0.492	1.000	0.000	54411
Firm age	AGE	3.143	3.497	0.796	3.930	-2.485	161687
Diversification dummy	$DIVER$	0.436	0.000	0.496	1.000	0.000	56616
Earnings quality	$EQ2$	0.053	0.034	0.060	0.433	0.002	10004
Earnings quality	$EQ5$	0.094	0.064	0.093	0.567	0.005	9920

Table II
Panel Regression of Idiosyncratic Volatility on Corporate Governance

This table reports estimates of coefficients of the monthly time-series cross-sectional firm-level regression

$$\Psi_{it} = c_0 + c_1 GOV_{i,t-1} + c_2 ROE_{i,t-1} + c_3 VROE_{i,t-1} + c_4 LEV_{i,t-1} + c_5 M/B_{i,t-1} + c_6 SIZE_{i,t-1} + c_7 DD_{i,t-1} + c_8 AGE_{i,t-1} + c_9 DIVER_{i,t-1} + \epsilon_{it},$$

where Ψ is the logistic transformed relative idiosyncratic volatility. GOV is alternatively: G, the IRRC-Gompers et al. (2003a) governance index; GD, which is zero if the governance index is less than or equal to five (open portfolio) and one if the index is greater than or equal to 14 (closed portfolio); and ATI, the anti-takeover provisions index, which incorporates only three anti-takeover provisions from the governance index. Firm/years with intermediate index values are not included when using GD. ROE is the return-on-equity. $VROE$ is the sample variance of ROE over the last three years. LEV is the ratio of long-term debt to total assets. M/B is the log of market-to-book equity ratio. $SIZE$ is the log market capitalization. DD is a dummy to identify dividend-paying firms. AGE is log number of years since listing. $DIVER$ is a diversified-firm dummy, which equals one when a firm operates in multiple segments and zero otherwise. Regressions include two-digit SIC industry fixed effects. The sample period is from January 1990 to December 2001. Financial and utilities industries are omitted (SIC 6000-6999 and 4900-4999). All variables are winsorized at the bottom- and top-1%. Newey-West t-statistics with three lags are in parentheses. Coefficients significant at the 5% level are in boldface.

	(1)	(2)	(3)	(4)	(5)	(6)
G	**-0.0289**	**-0.0129**				
	(-14.70)	(-5.51)				
GD			**-0.2173**	**-0.2694**		
			(-7.29)	(-7.20)		
ATI					**-0.0875**	**-0.0145**
					(-14.98)	(-2.19)
ROE		**0.0645**		0.0154		**0.0652**
		(2.23)		(0.22)		(2.25)
$VROE$		-0.0006		0.0190		-0.0007
		(-0.10)		(1.07)		(-0.12)
LEV		**0.2345**		**0.3391**		**0.2365**
		(5.31)		(3.13)		(5.35)
M/B		**-0.0630**		**-0.0946**		**-0.0625**
		(-5.29)		(-3.22)		(-5.25)
$SIZE$		**-0.3800**		**-0.3714**		**-0.3796**
		(-72.42)		(-30.01)		(-72.07)
DD		**0.0998**		**0.2585**		**0.0906**
		(6.45)		(6.32)		(5.90)
AGE		**0.0395**		**0.0803**		**0.0266**
		(3.93)		(3.42)		(2.73)
$DIVER$		**-0.0893**		**-0.0545**		**-0.0923**
		(-6.69)		(-1.61)		(-6.91)
R^2	1.31%	7.90%	2.09%	8.10%	1.32%	7.88%
N	160456	119541	28216	21315	160456	119541

Table III
Change in Idiosyncratic Volatility Following Corporate Governance Events

Panel A reports estimates of an event-study regression for Ψ, the logistic transformed relative idiosyncratic volatility on changes in the IRRC-Gompers et al. (2003a) governance index G. The event window includes, alternatively, the one-, two-, and three years before and after the year of the event. We run separate regressions for positive and negative G changes, using only data within the window. In each case, the regressor is one of several event dummy variables. $I_{\Delta G^+}$ is a dummy variable that takes a value of one for the years that fall on and after the firm the positive change in G, and zero for the years that fall before the increase in G. $I_{\Delta G^-}$ is a dummy variable that takes a value of one for the years that fall on and after the negative change in G, and zero for the years that fall before the decrease in G. ΔG^+ is a variable that takes a value equal to the positive change in G for the years that fall on and after the increase in G, and zero for the years that fall before the increase in G. ΔG^- is a variable that takes a value equal to the absolute value of the negative change in G for the years that fall on and after the decrease in G, and zero for the years that fall before the decrease in G. Panel B reports estimates of an event-study regression for Ψ on a takeover-target event dummy variable and an interaction variable. The window includes the six-, 12- and 18-months just prior to the takeover period that ends with the announcement date. $EVENT$ is a dummy variable that takes the value one for the months that fall on just before the announcement date, and zero for the prior months. $EVENT \times G$ is an interaction variable, which takes the value of the firm G index on and after the announcement, zero otherwise. Regressions in columns (2), (4), and (6) include two-digit SIC industry and year fixed effects. The sample period is from January 1990 to December 2001. Financial and utilities industries are omitted (SIC 6000-6999 and 4900-4999). All variables are winsorized at the bottom- and top-1%. Newey-West t-statistics with three lags are in parentheses. Coefficients significant at the 5% level are in boldface.

	(1)	(2)	(3)	(4)	(5)	(6)
	1-year		2-year		3-year	
	Panel A: Governance Index Changes (ΔG)					
$I_{\Delta G^+}$	**-0.0716**	**-0.0703**	**-0.0892**	**-0.0893**	**-0.0793**	**-0.1000**
	(-2.77)	(-2.74)	(-3.22)	(-3.21)	(-3.36)	(-4.23)
ΔG^+	**-0.0467**	**-0.0489**	**-0.0282**	**-0.0393**	**-0.0400**	**-0.0552**
	(-3.47)	(-3.62)	(-2.30)	(-3.15)	(-3.71)	(-5.01)
Number of events	543		394		341	
$I_{\Delta G^-}$	**0.2014**	**0.1986**	**0.2049**	**0.1987**	**0.3110**	**0.3127**
	(4.53)	(4.51)	(5.53)	(5.43)	(9.33)	(9.48)
Abs(ΔG^-)	**0.0560**	**0.0576**	**0.0541**	**0.0584**	**0.1039**	**0.1108**
	(2.18)	(2.23)	(2.46)	(2.63)	(5.30)	(5.66)
Number of events	1709		1333		1197	
	6-month		12-month		18-month	
	Panel B: Takeover Events - G Firms Targets					
$EVENT$	**0.2893**	**0.3274**	**0.3829**	**0.3921**	**0.3362**	**0.4088**
	(3.54)	(3.93)	(5.36)	(5.40)	(5.16)	(5.99)
$EVENT \times G$	**-0.0171**	**-0.0264**	**-0.0296**	**-0.0344**	**-0.0356**	**-0.0389**
	(-2.11)	(-3.22)	(-4.13)	(-4.82)	(-5.51)	(-5.96)
Number of events	1259		1073		942	
Industry fixed effects	No	Yes	No	Yes	No	Yes
Year fixed effects	No	Yes	No	Yes	No	Yes

Table IV
Panel Regression of Alternative Information Measures and Future Earnings Response on Corporate Governance Index

Columns (1)-(6) report estimates of coefficients of the time-series cross-sectional firm-level regression

$$\begin{aligned} INFO_{it} &= c_0 + c_1 GOV_{i,t-1} + c_2 ROE_{i,t-1} + c_3 VROE_{i,t-1} + c_4 LEV_{i,t-1} + c_5 M/B_{i,t-1} \\ &\quad + c_6 SIZE_{i,t-1} + c_7 DD_{i,t-1} + c_8 AGE_{i,t-1} + c_9 DIVER_{i,t-1} + \epsilon_{it}, \end{aligned}$$

where $INFO$ is, alternatively: $TURN$, the monthly share volume divided by shares outstanding; PIN, the annual probability of information-based trading of Easley et al. (2002); and $PRIVATE$, the annual amount of private information trading of Llorente et al. (2002). $INFO$ regressions are annually (i.e. subscript t refers to years) except the one for $TURN$, which is monthly, and include two-digit SIC industry fixed effects. GOV is, alternatively: G, the IRRC-Gompers et al. (2003a) governance index; and GD, which is zero if the governance index is less than or equal to five (open portfolio) and one if the index is greater than or equal to 14 (closed portfolio); Firm/years with intermediate index values are not included when using GD. ROE is the return-on-equity. $VROE$ is the sample variance of ROE over the last three years. LEV is the ratio of long-term debt to total assets. M/B is the log of market-to-book equity ratio. $SIZE$ is the log market capitalization. DD is a dummy variable to indentify dividend paying firms. AGE is log number of years since listing. $DIVER$ is a firm-diversification dummy, which equals one when a firm operates in multisegments and zero otherwise. Columns (7)-(8) report estimates of coefficients of the time-series cross-sectional industry-level regression

$$\begin{aligned} FER_{kt} &= c_0 + c_1 G_{k,t-1} + c_2 ROE_{k,t-1} + c_3 VROE_{k,t-1} + c_4 LEV_{k,t-1} + c_5 M/B_{k,t-1} \\ &\quad + c_6 SIZE_{k,t-1} + c_7 DD_{k,t-1} + c_8 AGE_{k,t-1} + c_9 DIVER_{k,t-1} + \epsilon_{kt}, \end{aligned}$$

where FER for each industry k is, alternatively, $FERC$ the annual future earnings response coefficient; and $FINC$ the annual futures earnings incremental explanatory power. $FINC$ and $FERC$ regressions are estimated at the two-digit SIC industry-level and include one-digit SIC industry fixed effects. Regressors are two-digit SIC industry averages. The sample period is from 1990 to 2001. Financial and utilities industries are omitted (SIC 6000-6999 and 4900-4999). All variables are winsorized at the bottom- and top-1%. Newey-West t-statistics with three lags are in parentheses. Coefficients significant at the 5% level are in boldface.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	$TURN$		PIN		$PRIVATE$		$FERC$	$FINC$
G	**-0.0013**		**-0.0008**		**-0.0006**		**-2.1178**	**-0.1020**
	(-9.46)		(-4.71)		(-2.41)		(-2.26)	(-3.28)
GD		**-0.0102**		**-0.0084**		**-0.0128**		
		(-7.43)		(-2.94)		(-3.15)		
ROE	**-0.0205**	**-0.0151**	0.0019	0.0004	**-0.0063**	**-0.0044**	3.0095	0.2017
	(-13.32)	(-3.86)	(0.84)	(0.06)	(-2.16)	(-0.57)	(0.32)	(0.66)
$VROE$	**-0.0011**	**-0.0003**	**0.0009**	**0.0008**	**0.0009**	**0.0004**	-2.6691	-0.0476
	(-4.06)	(-0.32)	(3.95)	(2.80)	(3.65)	(0.94)	(-1.72)	(-0.91)
LEV	**-0.0138**	**-0.0308**	**-0.0191**	**-0.0175**	-0.0076	**0.0336**	16.3156	0.1214
	(-7.25)	(-5.89)	(-5.84)	(-2.09)	(-1.62)	(3.23)	(1.13)	(0.31)
M/B	**0.0151**	**0.0239**	**-0.0034**	-0.0033	-0.0006	**-0.0064**	6.8113	-0.0445
	(27.36)	(16.65)	(-3.90)	(-1.50)	(-0.48)	(-2.13)	(1.54)	(-0.41)
$SIZE$	**0.0125**	**0.0110**	**-0.0226**	**-0.0226**	**-0.0064**	**-0.0033**	-3.3128	0.1063
	(58.26)	(21.77)	(-58.77)	(-21.54)	(-10.72)	(-2.29)	(-1.20)	(1.84)
DD	**-0.0713**	**-0.0797**	**0.0070**	**0.0053**	0.0029	0.0019	7.5963	0.0651
	(-98.47)	(-44.29)	(6.04)	(1.42)	(1.71)	(0.47)	(1.25)	(0.35)
AGE	**-0.0208**	**-0.0175**	0.0002	-0.0003	**0.0055**	**0.0065**	3.9986	0.0242
	(-44.16)	(-17.15)	(0.35)	(-0.21)	(5.47)	(2.92)	(1.90)	(0.34)
$DIVER$	**-0.0128**	**-0.0087**	**-0.0030**	-0.0015	**-0.0079**	-0.0009	1.4557	-0.1346
	(-23.77)	(-6.17)	(-3.20)	(-0.59)	(-5.24)	(-0.26)	(0.23)	(-0.63)
R^2	25.78%	29.76%	44.59%	44.97%	3.95%	9.13%	12.75%	17.37%
N	119358	20958	8155	1316	10347	1847	160	160

Table V
Panel Regression of Idiosyncratic Volatility on Corporate Governance and Institutional Trading

This table reports estimates of coefficients of the monthly time-series cross-sectional firm-level regression

$$\Psi_{it} = c_0 + c_1 G_{i,t-1} + c_2 ROE_{i,t-1} + c_3 VROE_{i,t-1} + c_4 LEV_{i,t-1} + c_5 M/B_{i,t-1} + c_6 SIZE_{i,t-1} + c_7 DD_{i,t-1} + c_8 AGE_{i,t-1} + c_9 DIVER_{i,t-1} + c_{10} INS_{i,t-1} + c_{11} INS_{i,t-1} G_{i,t-1} + \epsilon_{it},$$

where Ψ is the logistic transformed relative idiosyncratic volatility. G is the IRRC-Gompers et al. (2003a) governance index. ROE is the return-on-equity. $VROE$ is the sample variance of ROE over the last three years. LEV is the ratio of long-term debt to total assets. M/B is the log of market-to-book equity ratio. $SIZE$ is the log market capitalization. DD is a dummy variable to identify dividend-paying firms. AGE is log number of years since listing. $DIVER$ is a firm-diversification dummy, which equals one when a firm operates in multisegments and zero otherwise. INS alternatively refers to: $INST$, the quarterly average of the absolute change in the number of shares held by institutions, as a fraction of annual trading volume; and $INSTA$, the quarterly average of the absolute change in the number of shares held by takeover arbitrage institutions, as a fraction of annual trading volume. Regressions include two-digit SIC industry fixed effects. The sample period is from January 1990 to December 2001. Financial and utilities industries are omitted (SIC 6000-6999 and 4900-4999). All variables are winsorized at the bottom- and top-1%. Newey-West t-statistics with three lags are in parentheses. Coefficients significant at the 5% level are in boldface.

	(1)	(2)	(3)	(4)
G	**-0.0136**	**-0.0074**	**-0.0136**	**-0.0077**
	(-5.33)	(-2.05)	(-5.25)	(-2.13)
ROE	**0.0900**	**0.0891**	**0.1027**	**0.1019**
	(2.84)	(2.81)	(3.22)	(3.19)
$VROE$	0.0012	0.0012	0.0005	0.0004
	(0.15)	(0.15)	(0.06)	(0.05)
LEV	**0.2968**	**0.2935**	**0.2705**	**0.2666**
	(6.05)	(5.98)	(5.45)	(5.37)
M/B	**-0.0738**	**-0.0733**	**-0.0763**	**-0.0757**
	(-5.63)	(-5.58)	(-5.76)	(-5.71)
$SIZE$	**-0.3689**	**-0.3692**	**-0.3641**	**-0.3643**
	(-63.29)	(-63.34)	(-61.96)	(-62.01)
DD	**0.1402**	**0.1394**	**0.1364**	**0.1353**
	(8.35)	(8.30)	(8.02)	(7.95)
AGE	**0.0340**	**0.0338**	**0.0356**	**0.0357**
	(3.09)	(3.08)	(3.21)	(3.22)
$DIVER$	**-0.0837**	**-0.0837**	**-0.0877**	**-0.0878**
	(-5.78)	(-5.78)	(-5.98)	(-5.99)
$INST$	0.0990	**0.5003**		
	(1.77)	(2.80)		
$INST \times G$		**-0.0438**		
		(-2.38)		
$INSTA$			**0.4360**	**0.9817**
			(5.67)	(3.89)
$INSTA \times G$				**-0.0595**
				(-2.30)
R^2	7.78%	7.79%	7.73%	7.74%
N	101169	101169	99075	99075

Table VI
Panel Regression of Idiosyncratic Volatility on Corporate Governance and Accounting Transparency

This table reports estimates of coefficients of the monthly time-series cross-sectional firm-level regression

$$\Psi_{i,t} = c_0 + c_1 GOV_{i,t-1} + c_2 ROE_{i,t-1} + c_3 VROE_{i,t-1} + c_4 LEV_{i,t-1} + c_5 M/B_{i,t-1} + c_6 SIZE_{i,t-1} + c_7 DD_{i,t-1} + c_8 AGE_{i,t-1} + c_9 DIVER_{i,t-1} + c_{10} EQ_{i,t-1} + \epsilon_{i,t},$$

where Ψ is the logistic transformed relative idiosyncratic volatility. GOV is alternatively: G, the IRRC-Gompers et al. (2003a) governance index; and GD, which is zero if the governance index is less than or equal to five (open portfolio) and one if the index is greater than or equal to 14 (closed portfolio); and ATI, the anti-takeover provisions index, which incorporates only three anti-takeover provisions. Firm/years with intermediate index values are not included when using GD. ROE is the return-on-equity. $VROE$ is the sample variance of ROE over the last three years. LEV is the ratio of long-term debt to total assets. M/B is the log of market-to-book equity ratio. $SIZE$ is the log market capitalization. DD is a dummy variable to identify dividend-paying firms. AGE is log number of years since listing. $DIVER$ is a firm-diversification dummy, which equals one when a firm operates in multisegments and zero otherwise. EQ is alternatively: $EQ2$, the annual measure of accounting opaqueness defined as the absolute value of firm-specific residuals from an industry regression of current accruals on (the reciprocal of) assets and revenue growth; and $EQ5$, the annual measure of accounting opaqueness defined as the absolute value of firm-specific residuals from an industry regression of total accruals on lagged, contemporaneous, and lead cash flow from operations. Regressions include two-digit SIC industry fixed effects. The sample period is from January 1990 to December 2001. Financial and utilities industries are omitted (SIC 6000-6999 and 4900-4999). All variables are winsorized at the bottom- and top-1%. Newey-West t-statistics with three lags are in parentheses. Coefficients significant at the 5% level are in boldface.

	(1)	(2)	(3)	(4)
G	**-0.0087**		**-0.0101**	
	(-3.24)		(-3.80)	
GD		**-0.2182**		**-0.2010**
		(-5.04)		(-4.71)
ROE	0.0309	-0.0275	0.0194	-0.0046
	(0.93)	(-0.35)	(0.59)	(-0.06)
$VROE$	0.0037	0.0203	0.0034	0.0191
	(0.46)	(1.13)	(0.43)	(1.07)
LEV	**0.2626**	0.2092	**0.2751**	0.1647
	(4.96)	(1.60)	(5.16)	(1.27)
M/B	**-0.0872**	**-0.0914**	**-0.0872**	**-0.0805**
	(-6.16)	(-2.63)	(-6.19)	(-2.33)
$SIZE$	**-0.3769**	**-0.3636**	**-0.3666**	**-0.3719**
	(-61.18)	(-24.47)	(-59.68)	(-25.87)
DD	**0.1269**	**0.2579**	**0.1132**	**0.2487**
	(7.22)	(5.53)	(6.43)	(5.31)
AGE	**0.0379**	**0.0862**	**0.0333**	**0.0752**
	(3.31)	(3.23)	(2.89)	(2.78)
$DIVER$	**-0.0922**	-0.0465	**-0.0983**	-0.0281
	(-6.01)	(-1.20)	(-6.41)	(-0.73)
$EQ2$	**-0.4632**	**-0.7336**		
	(-3.62)	(-2.29)		
$EQ5$			**-0.3207**	-0.0933
			(-3.88)	(-0.48)
R^2	7.94%	7.67%	7.73%	7.88%
N	93118	16676	91964	16719

Table VII
Fama-MacBeth Regressions of Idiosyncratic Volatility on Corporate Governance

This table mean estimates of coefficients of the monthly cross-sectional firm-level regression

$$\begin{aligned}\Psi_{it} &= c_0 + c_1 GOV_{i,t-1} + c_2 ROE_{i,t-1} + c_3 VROE_{i,t-1} + c_4 LEV_{i,t-1} + c_5 M/B_{i,t-1} \\ &\quad + c_6 SIZE_{i,t-1} + c_7 DD_{i,t-1} + c_8 AGE_{i,t-1} + c_9 DIVER_{i,t-1} + \epsilon_{it},\end{aligned}$$

where Ψ is the logistic transformed relative idiosyncratic volatility, i indexes firms and t indexes months. GOV is alternatively: G, the IRRC-Gompers et al. (2003a) governance index; and GD, which is zero if the governance index is less than or equal to five (open portfolio) and one if the index is greater than or equal to 14 (closed portfolio). Firm/years with intermediate index values are not included when using GD. ROE is the return-on-equity. $VROE$ is the sample variance of ROE over the last three years. LEV is the ratio of long-term debt to total assets. M/B is the log of market-to-book equity ratio. $SIZE$ is the log market capitalization. DD is a dummy to identify dividend-paying firms. AGE is log number of years since listing. $DIVER$ is a firm-diversification dummy, which equals one when a firm operates in multisegments and zero otherwise. Regressions include two-digit SIC industry fixed effects. The sample period is from January 1990 to December 2001. Financial and utilities industries are omitted (SIC 6000-6999 and 4900-4999). All variables are winsorized at the bottom- and top-1%. Fama-MacBeth t-statistics adjusted for heteroskedasticity and third-order autocorrelation are in parentheses. Coefficients significant at the 5% level are in boldface.

	Const	G	GD	ROE	$VROE$	LEV	M/B	$SIZE$	DD	AGE	$DIVER$
1990	**2.9952**	**-0.0288**									
-2001	(39.62)	(-5.76)									
	8.1292	**-0.0136**		-0.0334	-0.0028	**0.2783**	**-0.0584**	**-0.3951**	**0.1971**	0.0073	-0.0308
	(27.92)	(-6.00)		(-0.65)	(-0.22)	(4.84)	(-2.76)	(-18.19)	(6.39)	(0.67)	(-1.49)
1990	**3.0834**	**-0.0554**									
-1992	(26.00)	(-12.39)									
	9.5397	**-0.0175**		-0.1766	0.0086	0.1248	-0.0247	**-0.5293**	**0.2987**	-0.0043	0.0933
	(24.38)	(-3.87)		(-1.66)	(0.41)	(1.66)	(-0.66)	(-15.84)	(7.77)	(-0.29)	(3.09)
1993	**3.0834**	**-0.0554**									
-1994	(30.21)	(-4.71)									
	8.0094	**-0.0145**		-0.0439	0.0294	0.0134	0.0297	**-0.3798**	**0.2617**	0.0319	-0.0213
	(20.60)	(-3.19)		(-0.57)	(0.67)	(0.08)	(0.65)	(-12.32)	(6.75)	(1.82)	(-0.46)
1995	**3.1522**	**-0.0328**									
-1997	(30.20)	(-3.59)									
	8.9213	-0.0102		-0.0877	-0.0098	0.2650	-0.0249	**-0.4202**	0.0944	**-0.0550**	-0.0620
	(20.20)	(-1.76)		(-1.10)	(-0.73)	(2.53)	(-0.87)	(-14.99)	(1.21)	(-2.84)	(-2.27)
1998	**2.9949**	**-0.0276**									
-1999	(18.20)	(-4.00)									
	7.5745	**-0.0112**		-0.0522	-0.0286	**0.4669**	**-0.1203**	**-0.3410**	**0.1039**	0.0103	-0.0254
	(27.69)	(-2.96)		(-1.23)	(-0.68)	(4.47)	(-3.19)	(-16.64)	(2.02)	(0.69)	(-1.03)
2000	2.3526	0.0153									
-2001	(17.05)	(1.65)									
	5.4995	**-0.0143**		0.2919	-0.0155	**0.6049**	**-0.1854**	**-0.2258**	**0.2270**	**0.0902**	**-0.1853**
	(13.86)	(-3.01)		(1.85)	(-1.42)	(9.53)	(-3.91)	(-6.70)	(3.03)	(5.75)	(-7.12)
1990	**2.9283**		**-0.2010**								
-2001	(41.87)		(-4.61)								
	7.8554		**-0.2253**	-0.1350	-0.2088	**0.3633**	-0.0419	**-0.3981**	**0.2659**	**0.1017**	-0.0202
	(24.05)		(-6.95)	(-1.49)	(-0.75)	(3.53)	(-1.13)	(-16.11)	(6.47)	(3.66)	(-0.54)
1990	**2.8775**		**-0.3153**								
-1992	(23.24)		(-7.84)								
	9.6014		**-0.1354**	-0.2519	-0.4208	0.0801	0.0002	**-0.5268**	**0.3999**	-0.0534	**0.1220**
	(21.86)		(-2.01)	(-1.31)	(-0.75)	(0.47)	(0.00)	(-14.83)	(7.13)	(-1.28)	(2.03)
1993	**3.1965**		**-0.2590**								
-1994	(30.21)		(-4.71)								
	7.3808		**-0.2914**	-0.0510	0.1354	0.1059	-0.0094	**-0.3621**	**0.2592**	**0.1732**	0.0044
	(14.21)		(-4.05)	(-0.27)	(0.17)	(0.45)	(-0.10)	(-7.54)	(2.46)	(4.66)	(0.05)
1995	**3.0434**		**-0.2512**								
-1997	(28.80)		(-2.76)								
	8.3237		**-0.2726**	**-0.3762**	-0.7715	**0.8971**	0.0630	**-0.4369**	**0.2373**	**0.1355**	0.0106
	(17.76)		(-4.14)	(-2.10)	(-1.49)	(5.85)	(1.08)	(-13.52)	(2.36)	(2.55)	(0.15)
1998	**2.9795**		**-0.2675**								
-1999	(16.81)		(-3.47)								
	7.8754		**-0.2500**	-0.2479	0.0611	**0.5132**	-0.1220	**-0.3739**	**0.1603**	0.0895	-0.0610
	(24.40)		(-4.09)	(-1.54)	(0.53)	(2.21)	(-1.51)	(-14.46)	(2.13)	(1.92)	(-1.01)
2000	**2.5128**		**0.1701**								
-2001	(15.28)		(1.94)								
	4.9887		**-0.1985**	**0.4307**	0.3393	0.0951	**-0.2147**	**-0.2070**	**0.2198**	**0.2247**	**-0.2639**
	(9.37)		(-2.83)	(2.75)	(0.41)	(0.56)	(-2.51)	(-4.05)	(2.68)	(3.32)	(-4.03)

Table VIII
Robustness Checks of Regression of Idiosyncratic Volatility on Corporate Governance

This table reports estimates of coefficients of the monthly time-series cross-sectional firm-level regression

$$\Psi_{it} = c_0 + c_1 GOV_{i,t-1} + c_2 ROE_{i,t-1} + c_3 VROE_{i,t-1} + c_4 LEV_{i,t-1} + c_5 M/B_{i,t-1} + c_6 SIZE_{i,t-1} + c_7 DD_{i,t-1} + c_8 AGE_{i,t-1} + c_9 DIVER_{i,t-1} + \epsilon_{it},$$

where Ψ is the logistic transformed relative idiosyncratic volatility. GOV is alternatively: G, the IRRC-Gompers et al. (2003a) governance index; and GD, which is zero if the governance index is less than or equal to five (open portfolio) and one if the index is greater than or equal to 14 (closed portfolio). Firm/years with intermediate index values are not included when using GD. All other regressors are as defined for Table II (coefficients not shown in the table) and regressions include two-digit SIC industry fixed effects. Columns (1) and (2) report results using idiosyncratic volatility estimates from an industry model of returns according to the 48 industry SIC classification scheme in Fama and French (1997). Columns (3) and (4) report results using idiosyncratic volatility estimates from Fama-French three-factor model of returns. Columns (5) and (6) report results using idiosyncratic volatility estimates with the autocorrelation correction in French et al. (1987) for variance and in Scholes and Williams (1977) for covariance. Columns (7) and (8) report results using quarterly time-series cross-sectional firm-level regression. Columns (9) and (10) report results only considering the earliest (1990) governance index G. Columns (11) and (12) report results including year fixed effects. Column (13) reports results including firm fixed effects and year fixed effects (differences-in-differences). Newey-West t-statistics with three lags are in parentheses.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)
	Industry Model		Fama-French Model		Autocorrelation Correction		Quarterly		First G		Year Fixed Effects		Dif-in-differences
G	**-0.0067**		**-0.0024**		**-0.0054**		**-0.0192**		**-0.0102**		**-0.0142**		**-0.0412**
	(-4.58)		(-2.01)		(-2.09)		(-5.39)		(-3.95)		(-6.18)		(-5.08)
GD		**-0.0727**		**-0.0692**		**-0.2124**		**-0.4058**		**-0.1717**		**-0.2596**	
		(-3.08)		(-3.67)		(-5.20)		(-7.14)		(-4.11)		(-7.09)	
ROE	0.0174	-0.0181	0.0004	0.0010	0.0336	-0.0460	0.0290	-0.1216	-0.0080	-0.1416	0.0549	0.0234	-0.0469
	(0.96)	(-0.40)	(0.03)	(0.03)	(1.06)	(-0.56)	(0.69)	(-0.93)	(-0.22)	(-1.54)	(1.91)	(0.34)	(-1.53)
$VROE$	0.0019	0.0120	0.0059	0.0118	0.0079	0.0111	0.0097	0.0063	0.0115	0.0065	0.0020	0.0251	-0.0159
	(0.55)	(1.67)	(1.92)	(1.55)	(1.12)	(0.95)	(1.80)	(0.43)	(1.32)	(0.57)	(0.34)	(1.44)	(-1.94)
LEV	**0.0653**	**0.1887**	0.0223	0.0395	**0.2060**	**0.3833**	**0.3463**	**0.4749**	**0.2581**	**0.2904**	**0.2214**	**0.3849**	-0.0260
	(2.36)	(2.78)	(1.03)	(0.78)	(4.19)	(3.22)	(5.13)	(2.81)	(4.86)	(2.08)	(5.08)	(3.60)	(-0.31)
M/B	**-0.0198**	**-0.0770**	-0.0011	**-0.0350**	**-0.0761**	**-0.1024**	**-0.0787**	**-0.1343**	-0.0216	-0.0542	**-0.0921**	**-0.1174**	**-0.0841**
	(-2.63)	(-4.15)	(-0.19)	(-2.39)	(-5.72)	(-3.13)	(-4.37)	(-2.86)	(-1.52)	(-1.46)	(-7.78)	(-4.02)	(-4.04)
$SIZE$	**-0.4079**	**-0.4124**	**-0.2025**	**-0.2000**	**-0.2177**	**-0.1998**	**-0.4989**	**-0.5006**	**-0.3832**	**-0.3841**	**-0.3867**	**-0.3781**	**-0.3604**
	(-120.78)	(-51.92)	(-77.80)	(-30.92)	(-37.00)	(-14.50)	(-60.99)	(-25.42)	(-65.22)	(-25.80)	(-74.59)	(-31.07)	(-20.59)
DD	**0.1406**	**0.2887**	**0.0769**	**0.1534**	**0.0602**	**0.2163**	**0.1471**	**0.3507**	**0.1102**	**0.2083**	**0.1470**	**0.2556**	**0.0794**
	(14.51)	(11.49)	(9.91)	(7.78)	(3.49)	(4.88)	(6.19)	(5.49)	(5.66)	(4.37)	(9.45)	(6.27)	(2.15)
AGE	**0.0510**	0.0209	**0.0104**	0.0002	**0.0411**	**0.0778**	**0.0610**	**0.1108**	**0.0770**	**0.0832**	**0.0248**	**0.0766**	**-0.1601**
	(8.03)	(1.43)	(2.10)	(0.02)	(3.70)	(3.02)	(4.02)	(3.08)	(5.59)	(2.53)	(2.49)	(3.29)	(-2.72)
$DIVER$	-0.0110	-0.0127	**-0.0533**	**-0.0592**	**-0.0787**	**-0.0802**	**-0.1134**	-0.0940	**-0.0947**	-0.0630	**-0.0695**	-0.0389	**-0.1680**
	(-1.31)	(-0.60)	(-8.04)	(-3.46)	(-5.28)	(-2.16)	(-5.64)	(-1.83)	(-6.11)	(-1.57)	(-5.25)	(-1.15)	(-7.62)
R^2	17.86%	19.14%	7.99%	8.82%	2.59%	2.81%	16.04%	17.78%	8.04%	8.53%	11.25%	11.33%	19.82%
N	119341	21256	119609	21318	115103	20335	40625	7259	90617	16154	119541	21315	119541

Table IX
Panel Regression of Capital Budgeting Quality on Corporate Governance and Idiosyncratic Volatility

Panel A reports estimates of coefficients of the annual time-series cross-sectional industry-level regression

$$|\dot{q}-1|_{kt} = \alpha_0 + \alpha_1 \Psi_{k,t-1} + \alpha_2 G_{k,t-1} + \alpha_3 LEV_{k,t-1} + \alpha_4 M/B_{k,t-1} + \alpha_5 SIZE_{k,t-1} + \alpha_6 DIVER_{k,t-1} + \epsilon_{kt},$$

where $|\dot{q}-1|$ is the absolute deviation of the two-digit SIC industry marginal Tobin's q relative to one (see the Appendix for full details). Regressors are two-digit SIC industry averages. Ψ is the is the logistic transformed relative idiosyncratic volatility. G is the IRRC-Gompers et al. (2003a) governance index. LEV is the ratio of long-term debt to total assets. M/B is the log of market-to-book equity ratio. $SIZE$ is the log market capitalization. $DIVER$ is a firm-diversification dummy, which equals one when a firm operates in multisegments and zero otherwise. In column (6) of Panel A, we split idiosyncratic volatility Ψ into two components, $\Psi^{predicted}$ and $\Psi^{residual}$, using a linear regression of Ψ on G. Columns (1) amd (2) of Panel B reports estimates of coefficients of the annual time-series cross-sectional firm-level regressions

$$\begin{aligned} \Delta EV_{it} &= \sum_{k=1}^{K} \lambda_0^k I_{it}^k + \dot{q}^{overall} \Delta NFA_{it} + \sum_{k=1}^{K} \dot{q}^k I_{it}^k \Delta NFA_{it} + \lambda_1 D_{it} + \lambda_2 EV_{i,t-1} + u_{it}, \\ \ln(u_{it}^2) &= \alpha_0 + \alpha_1 \Psi_{it} + \alpha_2 G_{it} + \alpha_3 LEV_{it} + \alpha_4 M/B_{it} + \alpha_5 SIZE_{it} + \alpha_6 DIVER_{it} + \varepsilon_{it}, \end{aligned}$$

where ΔEV is the change in enterprise value, ΔNFA_t is the change in net fixed assets, $I_{i,t}^k$ is an indicator variable for the firm's industry (k), and D_t is the flow of cash disbursements to all investors. In column (3) of Panel B, we split idiosyncratic volatility Ψ into two components, $\Psi^{predicted}$ and $\Psi^{residual}$, using a linear regression of Ψ on G. Newey-West t-statistics with three lags are in parentheses. Coefficients significant at the 5% level are in boldface.

	(1)	(2)	(3)	(4)	(5)	(6)
	Panel A: Industry-level Regressions					
	OLS	OLS	OLS	Tobit-Left Censored	Tobit-Right Censored	OLS
Ψ	**-0.292**		**-0.298**	-0.106	0352	
	(2.17)		(2.16)	(0.51)	(1.35)	
G		-0.101	-0.104	**-0.188**	-0.155	
		(1.86)	(1.92)	(2.18)	(1.43)	
$\Psi^{predicted}$						1.867
						(1.69)
$\Psi^{residual}$						**-0.298**
						(2.16)
LEV	**-1.816**	**-2.724**	**-2.253**	-1.911	0.879	**-2.253**
	(2.25)	(3.27)	(2.65)	(1.35)	(0.48)	(2.65)
M/B	0.234	0.153	0.195	-0.151	-0.589	0.195
	(1.20)	(0.78)	(1.01)	(0.55)	(1.71)	(1.01)
$SIZE$	-0.145	-0.093	-0.147	0.115	**0.462**	-0.147
	(1.47)	(0.95)	(1.49)	(0.79	(2.52)	(0.149)
$DIVER$	-0.252	-0.059	-0.093	-0.489	-0.710	-0.093
	(0.85)	(0.22)	(0.35)	(1.12)	(1.34)	(0.35)
Constant	**4.361**	**4.213**	**5.390**	2.640	-3.465	1.051
	(3.09)	(2.91)	(3.48)	(1.13)	(1.24)	(0.49)
R^2	0.06	0.06	0.07			0.07
N	236	236	236	236	236	236
	Panel B: Firm-level Regressions					
Dependent Variable	ΔEV	$\ln(u^2)$	$\ln(u^2)$			
$\dot{q}^{overall}$	1.325					
	(3.42)					
Ψ		**-0.108**				
		(2.32)				
G		**-0.083**				
		(7.73)				
$\Psi^{predicted}$			-0.071			
			(0.79)			
$\Psi^{residual}$			**-0.113**			
			(2.10)			
LEV		**-1.554**	**-1.615**			
		(6.65)	(6.86)			
M/B		**1.214**	**1.249**			
		(20.22)	(20.72)			
$SIZE$		**-0.164**	**0.190**			
		(5.73)	(6.50)			
$DIVER$		-0.065	**-0.122**			
		(1.01)	(1.92)			
λ_1	-2.14					
	(0.91)					
λ_2	0.019					
	(0.12)					
Constant		0.227	-0.217			
		(0.48)	(0.46)			
R^2		0.46	0.46			
N		6790	6790			

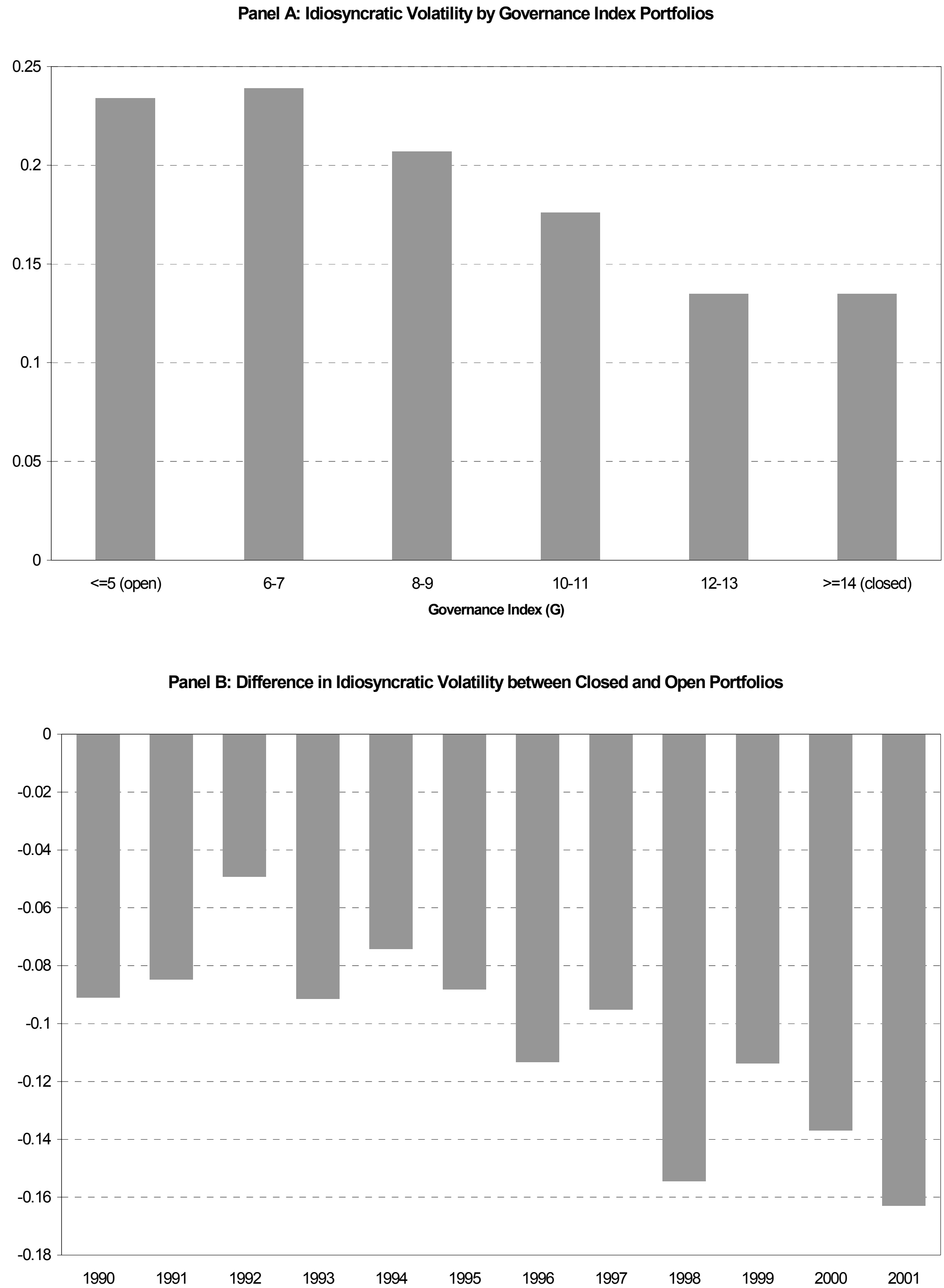

Figure 1. Idiosyncratic Volatility by Government Index. Panel A plots averages of annualized idiosyncratic variance by governance index (G) groups for the 1990-2001 period. Panel B plots the time series of the difference of annualized idiosyncratic variance between the closed and open portfolios. A firm is classified as open when G is less than or equal to five and as closed when G is greater than or equal to 14. Shaded bars represent differences that are significant at the five percent level.

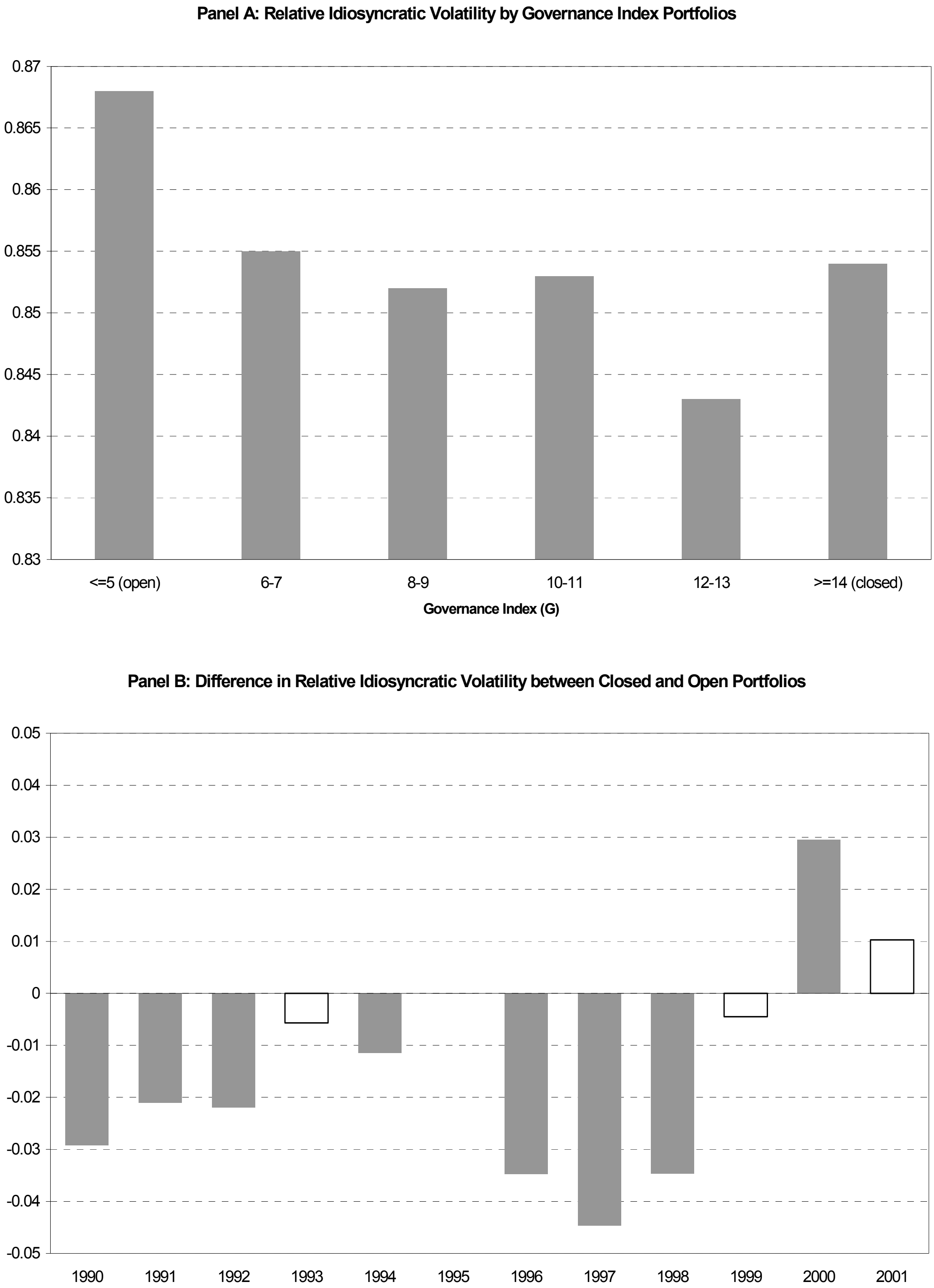

Figure 2. Relative Idiosyncratic Volatility by Government Index. Panel A plots averages of relative idiosyncratic variance by governance index (G) groups for the 1990-2001 period. Panel B plots the time series of the difference of relative idiosyncratic variance between the closed and open portfolios. A firm is classified as open when G is less than or equal to five and as closed when G is greater than or equal to 14. Shaded bars represent differences that are significant at the five percent level.