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Corporate Governance, Idiosyncratic Risk, and Information Flow

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ABSTRACT

We study the relationship of corporate governance policy and idiosyncratic risk. Firms with fewer antitakeover provisions display higher levels of idiosyncratic risk, trading activity, private information flow, and information about future earnings in stock prices. Trading interest by institutions, especially those active in merger arbitrage, strengthens the relationship of governance to idiosyncratic risk. Our results indicate that openness to the market for corporate control leads to more informative stock prices by encouraging collection of and trading on private information. Consistent with an information-flow interpretation, the component of volatility unrelated to governance is associated with the efficiency of corporate investment.

THE EFFECT OF CORPORATE GOVERNANCE on equity prices and the distribution of returns is an important issue in corporate finance. Gompers, Ishii, and Metrick (2003) and Cremers and Nair (2005) find that governance can directly influence equity prices. These and other authors generally posit that management constraints and incentives are the mechanisms through which governance influences prices. Any systematic effect on returns, however, also requires a link between governance provisions and investors' expectations or information. For instance, Gompers et al. (2003) argue that in the early 1990s, investors might not have fully appreciated the agency costs engendered by weak governance. This paper extends the current understanding by showing how governance provisions and informed trading interact to influence the incorporation of information into stock prices.

We test a trading link hypothesis, showing how specific aspects of governance that influence takeover vulnerability impact stock price informativeness. In particular, we focus on the specific path through the trading volume of arbitrage-oriented institutional investors. We reason that the absence of antitakeover provisions creates incentives to collect private information, which is a central determinant of idiosyncratic volatility. When trading activity is generated, it contributes to this idiosyncratic volatility and to other indications of

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private information flow. Our rationale follows Grossman and Stiglitz (1980), who predict that improving the cost–benefit trade-off on information collection leads to more informed trading and more informative pricing.

A priori, there are several reasons why fewer takeover restrictions could lead to more private information collection. First, fewer restrictions could imply a higher probability of a takeover (Ambrose and Megginson (1992)), providing traders more incentive to speculate. Larcker and Lys (1987) show that speculators in takeover situations are better informed about the likelihood of success, which suggests that, indeed, they have collected private information. Moreover, 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 collect and trade on private information in the pre-offer period. Second, fewer takeover restrictions could indicate that managers are not expecting a control offer (Comment and Schwert (1995)), implying that speculators may profit from correctly anticipating a higher probability of an offer. Third, fewer takeover restrictions could indicate that a firm's management or board would have limited bargaining power in the event of a control offer (Comment and Schwert (1995)), thereby attracting speculators who would prefer to quickly tender in response to an offer. Fourth, fundamental governance provisions that express openness to the market for control may be more common among firms that are open in other ways, including being open to sharing information with investors. Finally, strong investor protection, expressed by openness to takeovers, is associated with a reduced possibility of insiders (controlling shareholders and managers) expropriating outside investors. Thus, openness can directly encourage uninformed ownership and trading, thereby providing more cover for, and indirectly encouraging, privately informed trading. The final two possibilities suggest that governance provisions can affect information flow even when a merger is not imminent.

A stream of research establishes that volatility and information flow are closely associated. Ross (1989) shows that volatility is directly related to the rate of information arrival as an "important consequence of arbitrage-free economics." Strategic models and empirical evidence both establish that informed trade induces volatility (e.g., Glosten and Milgrom (1985) and French and Roll (1986)). Because private information is likely more common with respect to industries and firms than to the broad market, Roll (1988) focuses specifically on idiosyncratic volatility, providing evidence that idiosyncratic price changes mainly reflect private information being incorporated into stock prices by informed trading rather than public information. Thus, idiosyncratic volatility is a good candidate for a summary measure of information flow, especially for private information about firms.

Recent empirical evidence also supports this informational interpretation of idiosyncratic volatility. High levels of idiosyncratic volatility are associated with more efficient capital allocation (Durnev, Morck, and Yeung (2004)). Stock prices with high levels of idiosyncratic volatility contain more information about future earnings (Durnev et al. (2003)). 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 and opaque accounting induce low idiosyncratic volatility. Though this interpretation is not without controversy, our findings below support it.[1] To take into account the fact that limits to arbitrage, pricing errors, and noise also manifest in volatility, we verify our conclusions using other measures of information flow.

Our core empirical result is a strong negative relation between an index of a firm's antitakeover provisions, namely the Investor Responsibility Research Center (IRRC) governance index used by Gompers et al. (2003), and the firm-specific information impounded into stock prices as measured by idiosyncratic volatility.[2] We corroborate our explanation's focus on antitakeover measures by showing that a subset of powerful antitakeover provisions suggested by Cremers and Nair (2005) is a particularly important empirical determinant of idiosyncratic volatility. Further supporting an informational interpretation of our core result, we also 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 et al. (2002).[3] Additionally, using the method of Durnev et al. (2003), we find that stock prices in industries with fewer antitakeover provisions contain more information about future earnings.

Substantiating our trading link explanation, we provide direct evidence on trading as a mechanism through which governance is related to idiosyncratic risk. Specifically, this relationship persists and is stronger for stocks that are intensely traded by institutional investors, particularly institutional investors recently involved in risk arbitrage around takeovers. Thus, at least one of the links between governance and volatility and between governance and information flow is via arbitrage institutions. To our knowledge, such a link has not been previously documented in the literature.

To avoid spurious correlations, such as the possibility that larger firms might be less volatile and have more antitakeover measures, our study controls for a large set of volatility covariates suggested in the literature (see, e.g., Wei and Zhang (2006)). In view of the results of Bushee and Noe (2000), we also control for the transparency of firms' financial reporting and find that it is positively

[1] One concern is that restrictions on private information flow will not affect average volatility, since the information must eventually come out. A response explored by Durnev et al. (2004) is that the true stock value may be mean-reverting, in which case unexploited information depreciates.

[2] Core, Guay, and Rusticus (2006) find no link between governance and operating performance surprises, which raises questions about a governance–returns relation. Several papers consider a relation between idiosyncratic volatility and expected returns (e.g., Goyal and Santa-Clara (2003)). Added to our findings, these suggest the possibility of a governance to volatility to expected returns link.

[3] Idiosyncratic volatility, *PIN*, and the Llorente et al. (2002) measures each rely on different economic reasoning, yet they embody a common notion of stock price informativeness. Chen, Goldstein, and Jiang (2005) provide independent evidence that idiosyncratic volatility and *PIN* each reflect private information being incorporated via informed trading, in that both are significantly positively correlated with the sensitivity of real investment to stock price.

correlated with idiosyncratic volatility. However, this does not substitute for or replace the governance–risk relation. In a broader sense, 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.

To test whether takeover restrictions cause volatility—as opposed to the reverse—we study the change in idiosyncratic volatility following changes in the governance index. These tests focus directly on changes in idiosyncratic volatility that are likely attributable to the adoption or removal of takeover restrictions. If we are simply documenting self-selection (i.e., if firms for which there is little private information collection are more likely to adopt antitakeover provisions), idiosyncratic volatility should not shift around the change in the governance index. We find that idiosyncratic volatility is lower after a firm adopts takeover restrictions and higher after a firm eliminates them.

In a concluding analysis, we strengthen the interpretation of idiosyncratic risk as a measure of stock price informativeness by incorporating the governance–risk link into an examination of the relation between idiosyncratic risk and the quality of corporate investment decisions. Durnev et al. (2004) show that corporate investment decision-making quality increases in idiosyncratic volatility. Since good capital budgeting is one expression of good governance, the apparent relation of idiosyncratic risk to investment decision-making quality may be a statistical stand-in for an underlying economic relation between governance provisions and investing quality. Takeover restrictions might entrench current management, providing safety for poor investment decisions.

We find that antitakeover provisions are associated with a bias toward more investment for firms that otherwise seem to be underinvesting. Moreover, we can separate the level of volatility that is "expected" given the governance structure from the remaining "unexpected" volatility, and therefore consider how investment decision-making quality relates to each. While we find some evidence that expected volatility influences investing quality, there is a stronger relation between unexpected idiosyncratic volatility and investing quality, supporting the information-flow interpretation of volatility. Our results imply that information flow dominates the effects of governance provisions for investment decision-making quality, though both are important.

The paper proceeds as follows. We begin in Section I by describing our data. Section II presents our primary evidence on the relation between idiosyncratic volatility and antitakeover governance provisions. Section III provides supporting evidence and additional interpretation with respect to endogeneity concerns, alternative measures of private information flow, and the trading link hypothesis. Section IV considers the importance of accounting transparency for information flow as well as conducts other robustness tests. Section V examines the relationships among governance, idiosyncratic volatility, and corporate investment decision-making quality. Section VI concludes.

I. Data and Measures

We draw the data for our study from the IRRC database, the Center for Research in Stock Prices (CRSP) database, S&P COMPUSTAT, Thomson Financial's SDC Platinum M&A database, and Thomson's institutional ownership database compiled from S.E.C. 13F filings. Our initial sample includes all companies in the IRRC database for the period from 1990 to 2001, omitting financial firms and utilities (SIC 6000–6999 and 4900–4999). The resulting number of sample firms varies over time; however, on average there are 1,248 firms, with a minimum of 1,027 in 1992 and a maximum of 1,526 in 1998.

For all sample firms, we construct measures of idiosyncratic volatility, governance, and control variables. The average number of firms with both a governance index and idiosyncratic volatility is 1,140; the minimum is 943 in 1992 and the maximum is 1,514 in 1998.

A. Idiosyncratic Volatility and Measures of Information Flow

We study idiosyncratic volatility for each stock, estimated for each month using daily return data. Our measure of idiosyncratic volatility is based on a regression projection of stock returns on (alternatively) the returns of the market index, an industry index, and/or other factors. Consider first the case of the market model. For stock i,

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

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

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

where $\sigma_i^2 \equiv \mathrm{Var}(r_{id})$. We use sums of squares of daily returns in each month t to estimate monthly return variances, and sums of cross-products to estimate return covariances. In addition to developing our results for market-model idiosyncratic risk, we examine the robustness of our results using two other models, the Fama and French (1992) three-factor model and an industry factor model. Estimation of idiosyncratic volatility for these multifactor models is analogous to that for the market model.

From idiosyncratic volatility, we compute each stock's relative idiosyncratic volatility, as the ratio of idiosyncratic volatility to total volatility, $\frac{\sigma_{ie,t}^2}{\sigma_{it}^2}$, for each month t . 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 $1 - R_{it}^2$:

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

Our dependent variable Ψ_{it} measures idiosyncratic volatility relative to market-wide variation, or in other words, lack of synchronicity with the market.[4] One reason for scaling idiosyncratic volatility by the total variation in returns is that firms in some industries are more subject to economy-wide shocks than others, and firm-specific events may be correspondingly more intense. Additionally, this scaling allows for comparability to other studies, such as Durnev et al. (2004).

Panel A of Table I briefly describes the volatility variables. Panel A of Table II 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 to December 2001). For these panels, we estimate volatility for each sample month t, yielding observations for 161,691 firm-months. The mean idiosyncratic variance (annualized) is 0.194, which corresponds to an annualized standard deviation of 44%. Idiosyncratic volatility represents more than 85% of total individual stock volatility, on average.

Our focus on idiosyncratic volatility is motivated by information flow. To provide further evidence on this interpretation of idiosyncratic volatility, and to more completely test our information-flow hypotheses, we also investigate several alternative measures of information flow. These include share turnover (*TURN*), the probability of information-based trading (*PIN*) measure of Easley et al. (2002), the private information trading (*PRIVATE*) measure of Llorente et al. (2002), and the future earnings response coefficient (*FERC*) and future earnings incremental explanatory power (*FINC*) measures of Durnev et al. (2003). Panels B and C of Table I provide brief descriptions of these variables, and Panels B and C of Table II provide descriptive statistics.

B. Corporate Governance Index

A key independent variable in this paper is the IRRC governance index, which we denote as G, as used in Gompers et al.(2003).[5] We construct the index G for each sample firm for the years 1990, 1993, 1995, 1998, and 2000 from observations on a set of antitakeover-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 cumulating the indicator variables for each of the 24 nonoverlapping provisions for each firm. Larger values of the governance index G indicate that a firm is more insulated from takeovers and, in the judgment of

[4] Absolute idiosyncratic volatility, $\sigma_{ie,t}^2$, is an alternative dependent variable. Results (not tabulated here) using this alternative specification with an additional control for systematic volatility are broadly consistent with those we report below. Moreover, alternative transformations of variance, such as the logarithm of variance and standard deviation, also lead to similar results.

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

Table I
Definitions of Variables

Variables		Definition
		Panel A: Idiosyncratic Volatility Variables (firm-level)
Idiosyncratic volatility	σ_e^2	Monthly idiosyncratic variance (multiplied by 12 to annualize) estimated from the market model
Relative idiosyncratic volatility	σ_e^2/σ^2	Monthly relative idiosyncratic volatility given by the ratio of idiosyncratic variance by total variance
Logistic relative idiosyncratic volatility	Ψ	Monthly logistic transformed relative idiosyncratic volatility estimated from the market model
		Panel B: Alternative Information Flow Variables (firm-level)
Turnover	*TURN*	Monthly share volume divided by shares outstanding
Probability of information-based trading	*PIN*	Annual probability of information-based trading of Easley et al. (2002)
Amount of private information trading	*PRIVATE*	Annual amount of private information trading of Llorente et al. (2002) given for each firm-year by the b_{i2^a} estimate of the time-series regression: $r_{id} = b^a{}_{i0} + b^a{}_{i1} r_{i,d-1} + b^a{}_{i2} r_{i,d-1} V_{i,d-1} + \epsilon^a_{id}$, where r_{id} is daily stock return and V_{id} is log daily turnover detrended by subtracting a 200 trading day moving average
		Panel C: Future Earnings Response Variables (two-digit SIC industries)
Future earnings response coefficient	*FERC*	Sum of the coefficients on future changes in earnings $\sum_{\tau=1}^{3} b^b_{2,\tau}$ of the annual regression on each two-digit SIC industry with at least 10 firms: $r_{it} = b^b_0 + b^b_1 \Delta E_{it} + \sum_{\tau=1}^{3} b^b_{2,\tau} \Delta E_{i,t+\tau} + \sum_{\tau=1}^{3} b^b_{3,\tau} r_{i,t+\tau} + \epsilon^b_{it}$, where r_{it} is annual stock return calculated from fiscal year-end share price plus dividends adjusted by stock splits and distributions (COMPUSTAT annual items #199/#27 plus #26/#27), and ΔE_{it} is annual change in earnings before interest, taxes, depreciation, and amortization (annual item 13) scaled by previous fiscal year-end market capitalization (annual items #199 times #25)
Future earnings incremental explanatory power	*FINC*	Increase in the coefficient of determination (R^2) of the annual regression on each two-digit SIC industry with at least 10 firms: $r_{it} = b^b_0 + b^b_1 \Delta E_{it} + \sum_{\tau=1}^{3} b^b_{2,\tau} \Delta E_{i,t+\tau} + \sum_{\tau=1}^{3} b^b_{3,\tau} r_{i,t+\tau} + \epsilon^b_{it}$, relative to the base regression: $r_{it} = b^c_0 + b^c_1 \Delta E_{it} + \epsilon^c_{it}$

(*continued*)

Table I—*Continued*

Variables		Definition
		Panel D: Corporate Governance Variables (firm-level)
Governance index	*G*	IRRC–Gompers et al. (2003) governance index, which is based on 24 antitakeover provisions
Antitakeover index	*ATI*	Cremers and Nair (2005) antitakeover provisions index, which incorporates three antitakeover provisions
		Panel E: Control Variables (firm-level)
Return-on-equity	*ROE*	Quarterly return-on-equity (multiplied by four to annualize) given by most recent quarterly earnings before extraordinary items (COMPUSTAT quarterly item #8) divided by the book value of equity (quarterly item #60)
Volatility of return-on-equity	*VROE*	Sample variance of quarterly *ROE*s over the last 3 years
Leverage	*LEV*	Quarterly leverage defined as the ratio of long-term debt (COMPUSTAT quarterly item #51) to total assets (quarterly item #44)
Market-to-book	*M/B*	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)
Market capitalization	*SIZE*	Monthly log market capitalization (CRSP)
Dividend dummy	*DD*	Quarterly dividend dummy, which equals one if the firms pays dividends, and zero otherwise (dividends given by COMPUSTAT quarterly item #20)
Firm age	*AGE*	Monthly log age defined as the number of months (divided by 12) since the stock inclusion in the CRSP database
Diversification dummy	*DIVER*	Quarterly dummy variable that equals one when a firm operates in multisegments and zero otherwise
		Panel F: Institutional Trading Activity Variables (firm-level)
Institutional trading	*INST*	Quarterly absolute change in the number of shares held by institutions, as a fraction of annual trading volume
Arbitrage institutional trading	*INSTA*	Quarterly absolute change in the number of shares held by merger arbitrage-active institutions, as a fraction of annual trading volume
		Panel G: Accounting Transparency Variables (firm-level)
Earnings quality, first version	*EQ*2	Annual absolute value of firm-specific residuals from a two-digit SIC annual industry regression of total accruals (change in (Δ) current assets (COMPUSTAT annual item #4) minus the Δcurrent liabilities (annual item #5) plus the Δdebt in current liabilities (annual item #34) minus the Δcash (annual item #2), and minus depreciation and amortization (annual item #14)) on (the reciprocal of) total assets (annual item #6), revenue (annual item #12) growth, and fixed assets (annual item #7); variables scaled by total assets
Earnings quality, alternative version	*EQ*5	Annual absolute value of firm-specific residuals from a two-digit SIC annual industry regression of total accruals on lagged, contemporaneous, and leading cash flow from operations (earnings before extraordinary items (COMPUSTAT annual item #18) minus total current accruals (total accruals plus depreciation and amortization)); variables scaled by total assets

Table II

Descriptive Statistics

This table reports mean, median, standard deviation, maximum, minimum, and number of observations of variables. All variables are as defined in Table I. 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% levels.

		Mean	Median	SD	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: 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 C: Future Earnings Response Variables (two-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 D: Corporate Governance Variables (firm-level)							
Governance index	*G*	8.971	9.000	2.833	19.000	1.000	6043
Antitakeover index	*ATI*	1.830	2.000	0.944	3.000	0.000	6043
Panel E: 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
Panel F: 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 G: Accounting Transparency Variables (firm-level)							
Earnings quality, first version	*EQ*2	0.053	0.034	0.060	0.433	0.002	10004
Earnings quality, alternative version	*EQ*5	0.094	0.064	0.093	0.567	0.005	9920

Gompers et al. (2003), is less shareholder-friendly. When we need to specify a governance index for a particular month t, we use the most recently announced level. Panel D of Table I briefly describes the governance variables. Panel D of Table II presents summary statistics for the governance variables as a panel of 6,043 distinct observations. The median G is 9.0 and the standard deviation is 2.8.

For robustness, we double-check our results using the raw index G against 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 as 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. We also conduct tests using the Cremers and Nair (2005) antitakeover index, which incorporates only the three provisions in the IRRC–Gompers et al. (2003) index that are thought to be most effective in deterring takeover activity and/or increasing target bargaining power. The Cremers and Nair (2005) index varies from zero to three, with one point being accorded for blank-check preferred stock authorization, one point for a classified (staggered elections) board structure, and one point for limitations on the ability of shareholders to call special meetings or act by written consent. We denote this index as ATI, for antitakeover index. We verify that these provisions are an empirical deterrent to the arrival of control offers for our sample firms. The median ATI is 2.0 and the standard deviation is 0.9. With our convention that larger values for G, GD, and ATI correspond to more antitakeover provisions, our indexes are inverse measures of a firm's openness to the market for corporate control.

Some additional control variables that we use in our tests are also briefly described in Panels E–G of Table I, with descriptive statistics provided in Table II. We winsorize extreme observations at the bottom and top 1% levels to avoid spurious inferences. We discuss these variables below as appropriate.

II. Governance and Idiosyncratic Volatility

In this section, we present graphical analysis, specify our basic empirical design, and provide panel regression evidence on the relation between governance and idiosyncratic volatility.

A. Graphical Analysis

Figure 1 presents a visual summary of annualized data for market-model idiosyncratic risk, σ_e^2. Following Gompers et al. (2003), 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 within each level of the governance index over the full sample period. A much higher level of idiosyncratic risk is associated with governance structures that

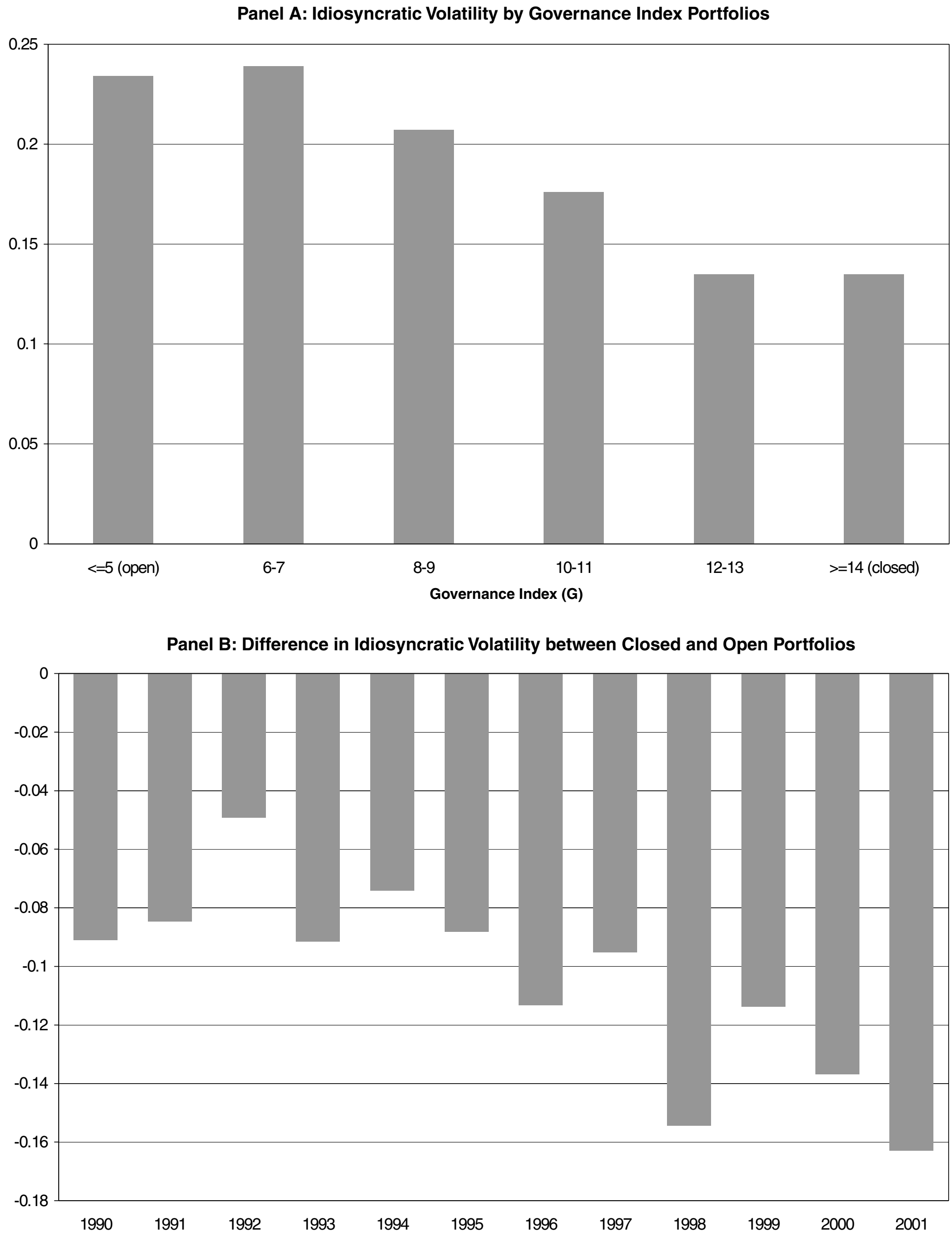

Figure 1. Idiosyncratic volatility by government index. Panel A plots averages of annualized idiosyncratic risk, σ_e^2, by governance index (G) groups for the period from January 1990 to December 2001. Panel B plots the time series of the difference of annualized idiosyncratic risk between the closed and open portfolios. A firm is classified as open when G is less than or equal to 5 and as closed when G is greater than or equal to 14. Shaded bars represent differences that are significant at the 5% level.

are very open to control offers. The open portfolio displays idiosyncratic variance of about 0.234, which corresponds to an annualized standard deviation of 48.4%. The corresponding figure for the closed portfolio is about 0.135, which corresponds to an annualized standard deviation of 36.7%. The difference is highly statistically significant (t-statistic = 27.74). The same comparison holds in every year, as shown in Panel B of Figure 1.

Similarly, Panel A of Figure 2 plots the average relative idiosyncratic volatility, σ_e^2/σ^2, according to the level of the governance index over the full sample period. Our core result is clear in the figure: Average relative idiosyncratic volatility for the open portfolio is greater than for the closed portfolio. The difference between the two extreme portfolios is highly statistically significant (t-statistic = 5.31). Moreover, all intermediate governance index portfolios present lower average relative idiosyncratic volatility than the open portfolio. The same comparison holds in every year, except at the end of the sample period, as shown in Panel B of Figure 2. The difference is statistically significant in 8 of 12 sample years.

B. Empirical Framework

The previous section establishes that firms with fewer antitakeover provisions have greater idiosyncratic volatility, on average. These differences could be driven by nongovernance factors that are only incidentally correlated with governance. In the remainder of this section, we establish that antitakeover aspects of governance are at the core of the relation. To do so, we begin by estimating the following regression equation:

$$\begin{aligned}\Psi_{it} = {} & c_0 + c_1 GOV_{i,t-1} + c_2 ROE_{i,t-1} + c_3 VROE_{i,t-1} + c_4 LEV_{i,t-1} + c_5 M/B_{i,t-1} \\ & + c_6 SIZE_{i,t-1} + c_7 DD_{i,t-1} + c_8 AGE_{i,t-1} + c_9 DIVER_{i,t-1} + \epsilon_{it},\end{aligned} \tag{4}$$

where i indexes firms, t is a monthly time index, and GOV is a particular measure of governance provisions. We include a number of control variables drawn from the literature on idiosyncratic risk in our empirical design. These include profitability (*ROE*), profits volatility (*VROE*), leverage (*LEV*), market-to-book ratio (*M/B*), equity capitalization (*SIZE*), a dividend payer dummy (*DD*), firm age (*AGE*), and an internal diversification dummy (*DIVER*). We measure variables for each firm-month where possible; when we need to match quarterly data with monthly data, we use the most recently observed quarterly figure. The quarterly earnings report date is used to determine when the information is available to investors (typically 2 months after the end of a fiscal quarter). A description of each control variable is in Panel E of Table I, with descriptive statistics provided in Table II.[6]

[6] 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 past 3 years. We also construct an alternative *VROE* variable as the residual standard deviation from an AR(1) process for firm-level quarterly *ROE*. This *VROE* measure uses all data available for each firm with sufficient consecutive observations and does not change over time. Panel regressions results (not tabulated here) using the alternative *VROE* are qualitatively the same as those we report.

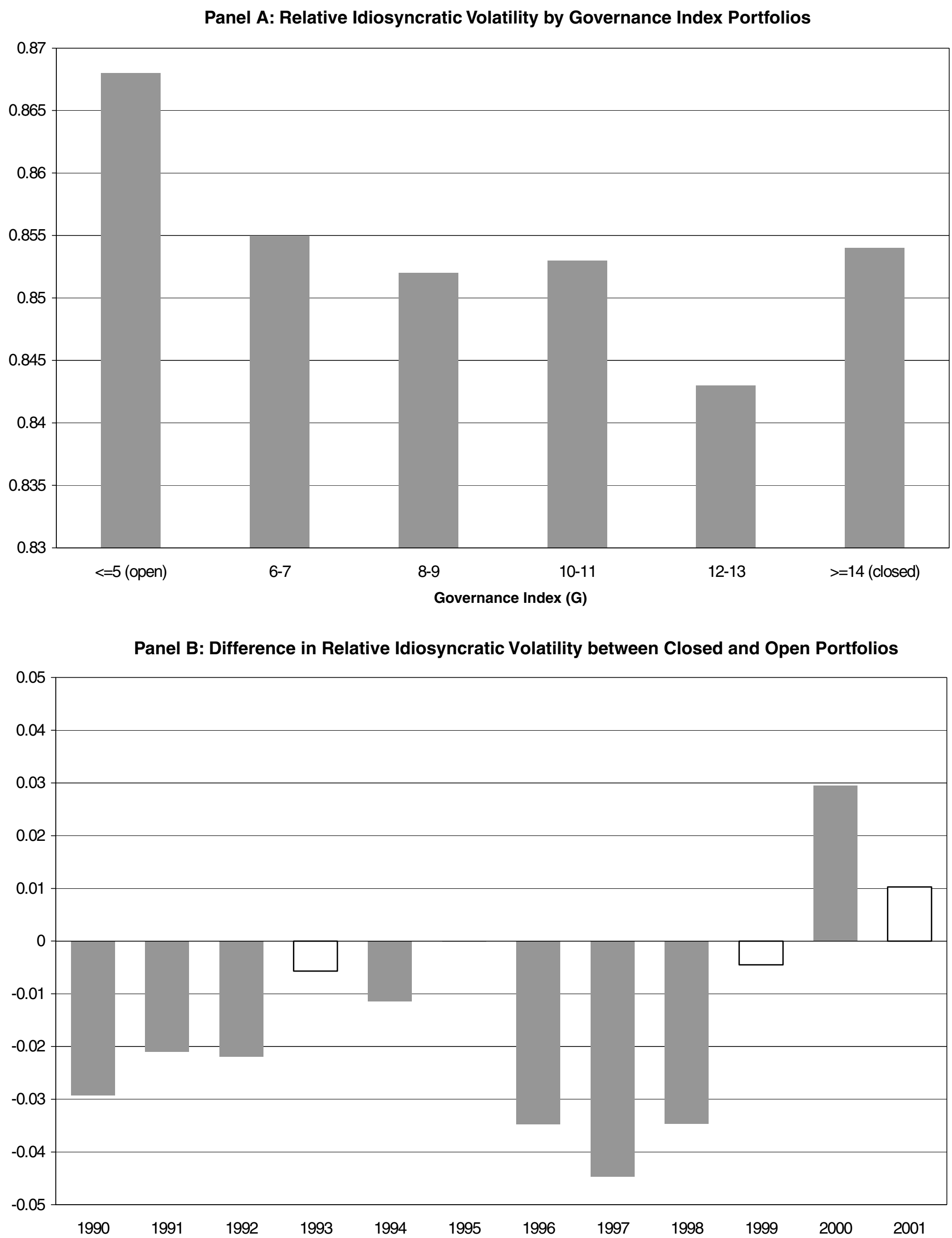

Figure 2. Relative idiosyncratic volatility by government index. Panel A plots averages of relative idiosyncratic risk, σ_e^2/σ^2, by governance index (G) groups for the period from January 1990 to December 2001. Panel B plots the time series of the difference of relative idiosyncratic risk between the closed and open portfolios. A firm is classified as open when G is less than or equal to 5 and as closed when G is greater than or equal to 14. Shaded bars represent differences that are significant at the 5% level.

In this section, we estimate equation (4) as a pooled cross-sectional time-series model. Regressions include two-digit SIC industry fixed effects to control for additional differences across industries. To verify robustness, we alternatively set *GOV* equal to each of the IRRC–Gompers et al. (2003) measures, *G* and *GD*, and to the Cremers and Nair (2005) antitakeover provision index, *ATI*. Lower values of *G*, *GD*, and *ATI* correspond to fewer antitakeover provisions. We are most 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.

Endogeneity is a well-known issue in governance regressions. As a first response, we always regress volatility on pre-determined 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. (2003) index, which lags by up to 3 years. In the case of other variables, we use the most recent observation. In a later section, we provide additional tests that focus on the time series of volatility changes.

C. Panel Regression Results

Table III presents estimates of the basic model in equation (4) in which logistic-transformed relative idiosyncratic volatility Ψ is the dependent variable. The table reports restricted versions of the basic model in which a governance index is the only regressor, as well as full versions with the complete set of control variables.

Columns (1), (3), and (5) display the restricted estimates. The consistent result is a significant negative relation between idiosyncratic volatility and a governance stance closed to takeovers. In column 1, for example, the regression coefficient on the *G* index is -0.0289 with a robust *t*-statistic of -14.70. Higher levels of the *G* index indicate less openness, so the relationship is clear. The same conclusion can be drawn from column (3), which uses *GD*: The estimated coefficient is -0.2173 with a *t*-statistic of -7.29. The negative relation between volatility and antitakeover provisions is confirmed when we use *ATI*, the antitakeover index. In column (5), the estimated *ATI* 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 are attenuated. Estimates are reported in columns (2), (4), and (6) of Table III. The estimated coefficient on the *G* index, for example, is -0.0129 with a *t* -statistic of -5.51. We conclude that antitakeover provisions are a strong statistical determinant of idiosyncratic volatility. This relation is also economically significant: Controlling for other firm characteristics, a one-point increase in the *G* index of the average firm reduces relative idiosyncratic volatility Ψ by 0.0129, or about 1.3% (see column (2)). More strikingly, the difference in relative idiosyncratic volatility between the closed and open portfolios is 26.9% (see column (4)).

In this section, we show a strong negative connection between antitakeover measures and idiosyncratic volatility. Durnev et al. (2004) and others

Table III

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. (2003) governance index; *GD*, which is zero if the governance index is less than or equal to 5 (open portfolio) and one if the index is greater than or equal to 14 (closed portfolio); and *ATI*, the antitakeover provisions index, which incorporates only three antitakeover provisions from the governance index. Firm-years with intermediate index values are not included when using *GD*. The regressors include profitability (*ROE*), profits volatility (*VROE*), leverage (*LEV*), market-to-book ratio (*M/B*), equity capitalization (*SIZE*), dividend-payer dummy (*DD*), firm age (*AGE*), and internal diversification dummy (*DIVER*). Refer to Table I for variable definitions. 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% levels. Robust *t*-statistics 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	160,456	119,541	28,216	21,315	160,456	119,541

recently argue that idiosyncratic volatility is an index of information intensity in general, and of the extent to which private information is revealed by trading in particular. This interpretation is consistent with older studies, as we have discussed above. Within that interpretation, our findings indicate that a governance stance that includes openness to takeovers results in more information flowing to the market via trading on private information.

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

In this section, we provide additional results to substantiate and interpret the negative relation between antitakeover provisions and idiosyncratic volatility. The first subsection presents evidence on idiosyncratic volatility changes following changes in takeover provisions, to address endogeneity concerns. The second subsection provides evidence that several direct measures of information flow have a similar relation to the governance index as does idiosyncratic risk. The final subsection provides direct evidence on a trading link between governance and volatility via arbitrage-active institutions.

A. Change in Idiosyncratic Volatility Following Governance Events

Endogeneity can be a serious concern with panel evidence of the type presented above. In particular, there is a possibility of reverse causality in that firms with lower (higher) idiosyncratic risk may be more likely to choose (avoid) takeover defenses. To address this concern, tests in this section focus on the change in idiosyncratic volatility around governance events, that is, changes in the governance index G. Specifically, we compare idiosyncratic volatility before and after the adoption or removal of takeover restrictions for a given firm. If we are simply documenting self-selection, we should not find a significant change in idiosyncratic volatility around changes in G.

To test whether there is a change in idiosyncratic volatility following an event, we regress the relative idiosyncratic volatility Ψ on a firm-specific indicator of event occurrence using a stacked data set of all events. We consider 1-year, 2-year, and 3-year event windows, with monthly data, following a change in G. We include only observations during the window just after a G change as compared to a window of equal length just before the change.[7] For example, using a 1-year window, we compare idiosyncratic volatility in the 1-year period before the change in G with the 1-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, by construction, the 1-year window does not have overlapping data. In contrast, the 2-year and 3-year windows may have overlapping events. In these cases, we exclude subsequent events with overlapping data. Results remain the same when we include overlapping events.

For our event indicator, we construct a post-change dummy variable that takes the value of one for the years after the G change, and zero for the

[7] By convention, we place G-change events at the beginning of each calendar year. This is arbitrary, and is one of the reasons for our use of alternate window lengths.

years before. We denote this indicator variable as $I_{\triangle G^+}$ for increases in G and $I_{\triangle G^-}$ for decreases in G. Thus, for example, the regression coefficient on the dummy variable $I_{\triangle G^+}$ gives an estimate of the difference in idiosyncratic volatility between the period following the adoption of antitakeover provisions and the prior period. We alternatively use a regressor that equals the numeric change in G for the years after the firm enacts a G change, and zero for the years before. We denote this variable $\triangle G^+$ for increases in G and $\triangle G^-$ for decreases in G (for easier interpretation we take the absolute value of the negative changes in G). In contrast to the dummy variable that simply registers whether a change has occurred, this variable measures the actual magnitude of the G change.

There are more cases of firms removing takeover restrictions than adopting them during our sample period. Using a 1-year window, there are 1,709 G - decrease events, while there are only 543 G-increase events. The vast majority of events corresponds to a one-point change in G: 72% of the total number of events in which G increases and 85% of the total number of events in which G decreases.

Table IV presents ordinary least squares (OLS) estimates of changes in idiosyncratic volatility, and alternatively, estimates including two-digit SIC industry and year dummies.[8] OLS estimation using the event dummy variable corresponds to a difference-in-means test of the relative idiosyncratic volatility Ψ between the periods immediately before and after the change in G. The results conform to our expectation that idiosyncratic volatility significantly decreases following adoption of antitakeover provisions. For example, using a 1-year window (see column 1), we find that Ψ drops 7.16% for an increase in G (adoption of one or more antitakeover provisions) with a t-statistic of -2.77.[9] Results are consistent across event windows, ranging from -0.0716 to -0.0892 for an increase in G. Results including industry-year dummies 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.

Additionally, there is strong evidence that idiosyncratic volatility increases after the G index declines. We find that Ψ increases by 20.14% following a G-decrease event with a t-statistic of 4.53 when using a 1-year event window. This is true for all event windows with the coefficient ranging from 0.2014 to 0.3110 for a G-decrease event using OLS. Results including industry and year dummies and those that consider the magnitude of the G change confirm that idiosyncratic volatility increases when a firm drops takeover restrictions.

Overall, we find that, indeed, idiosyncratic volatility is lower after a firm adopts takeover restrictions and higher after a firm removes takeover

[8] Results (not tabulated here) controlling for the changes in main control variables are similar to the results including industry and year dummies.

[9] The time-series effect is therefore even larger than the measured effect from pooled regressions presented earlier, and provides more evidence that the results are driven by causality.

Table IV

Change in Idiosyncratic Volatility Following Corporate Governance Events

This table reports estimates of an event study regression for Ψ, the logistic transformed relative idiosyncratic volatility, on changes in the IRRC–Gompers et al. (2003) governance index G. The event window includes, alternatively, the 1-, 2-, and 3-year period 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 post-event dummy variables. $I_{\Delta G^+}$ is a dummy variable that takes the value of one for the years that fall on and after 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 the 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 the 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 the 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. 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% levels. Robust t-statistics are in parentheses. Coefficients significant at the 5% level are in boldface.

	1-year		2-year		3-year	
	(1)	(2)	(3)	(4)	(5)	(6)
$I_{\Delta G^+}$	**−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)
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	1,709		1,333		1,197	
Industry dummies	No	Yes	No	Yes	No	Yes
Year dummies	No	Yes	No	Yes	No	Yes

restrictions. This event study evidence indicates that our earlier panel-based results are not likely driven by reverse causality.[10]

B. Governance and Private Information Trading

To substantiate our informational interpretation of the governance–volatility relationship, we next test for the relation between governance and several dependent variables that measure information flow more directly. The results

[10] Results (not tabulated here) extend our event study evidence on governance changes by considering takeover situations, instead of the changes in G, as the events. The results reveal an increase in idiosyncratic volatility during the takeover period, and support the hypothesis that the increase in idiosyncratic volatility in the takeover period mainly accrues for low G firms.

support the proposition that governance is a driver of information flow. To begin, we think of turnover as one alternative to idiosyncratic volatility in proxying for the intensity of private information flowing to a stock's market. Trading is theoretically linked to the quality or extent of private information (e.g., Blume, Easley, and O'Hara (1994)), and is thus a natural measure of private information flow. We investigate unsigned firm-level monthly turnover (*TURN*), a series formed by dividing monthly share volume by the number of shares outstanding. Additionally, we draw some direct information flow measures from the literature. Recent research provides several targeted private information flow indexes (in particular, *PIN* and *PRIVATE*, both discussed earlier) and some indexes of future earnings information contained in stock prices (such as *FINC* and *FERC*, also discussed earlier), which we also investigate in this section.[11] We estimate the following regression equation:

$$\begin{aligned} INF_{it} = c_0 &+ c_1 GOV_{i,t-1} + c_2 ROE_{i,t-1} + c_3 VROE_{i,t-1} + c_4 LEV_{i,t-1} \\ &+ c_5 M/B_{i,t-1} + c_6 SIZE_{i,t-1} + c_7 DD_{i,t-1} + c_8 AGE_{i,t-1} \\ &+ c_9 DIVER_{i,t-1} + \epsilon_{it}, \end{aligned} \tag{5}$$

where $GOV = \{G \vee GD\}$ and other regressors are the same as in equation (4) for idiosyncratic volatility. The variable *INF* refers to one of the measures just discussed, *TURN*, *PIN*, or *PRIVATE*; *TURN* is measured monthly, so in this case t is a monthly time index, and both *PIN* and *PRIVATE* are measured annually, so in these cases t refers to an annual index. We choose to include the same control variables as before because our goal is not to fully explore the cross-section of trading activity, but rather to control for influences on the extent of trading on private information. Nevertheless, our control regressors cover several categories of potential cross-sectional determinants of trading activity.

Columns (1) and (2) of Table V report results for the turnover regressions. The coefficient on the takeover restrictions index G in column 1 is negative and significant. Using the alternative dummy variable version GD in column (2), we also report a negative and significant coefficient. Thus, the evidence suggests that trading activity is higher in stocks of firms that are open to control offers. Coefficients on control variables are mostly consistent with expectations. Columns (3) and (4) of Table V present estimates of equation (5) using the annual probability of information trading (*PIN*) measure of Easley et al. (2002). We find that *PIN* is also negatively related to the governance index, which supports our hypothesis that open firms are more subject to private information trading. The coefficient of the takeover restriction index G in column (3) is negative and significant. Using the alternative dummy variable version GD in column (4), we also report a negative and significant coefficient. Columns (5) and (6) present estimates using the annual amount of private information trading measure (*PRIVATE*) of Llorente et al. (2002). We find that *PRIVATE* is also

[11] We thank Soeren Hvidkjaer for making data on *PIN* available on his website.

Table V

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

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

where *INF* 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). *INF* regressions are estimated on an annual basis (i.e., subscript t refers to years) except the one for *TURN*, which is estimated on an monthly basis, and include two-digit SIC industry fixed effects. *GOV* is, alternatively: *G*, the IRRC–Gompers et al. (2003) governance index; and *GD*, which is zero if the governance index is less than or equal to 5 (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*. The regressors include profitability (*ROE*), profits volatility (*VROE*), leverage (*LEV*), market-to-book ratio (*M/B*), equity capitalization (*SIZE*), dividend-payer dummy (*DD*), firm age (*AGE*), and internal diversification dummy (*DIVER*). Refer to Table I for variable definitions. Columns (7)–(8) report estimates of coefficients of the time-series cross-sectional industry-level regression

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

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% levels. Robust t-statistics are in parentheses. Coefficients significant at the 5% level are in boldface.

	TURN		*PIN*		*PRIVATE*		*FERC*	*FINC*
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
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	119,358	20,958	8,155	1,316	10,347	1,847	160	160

negatively related with the governance index. The coefficient of the takeover restrictions index G in column 5 is negative and significant, as is that for the alternative dummy variable version GD in column (6).

Next we test whether firms with fewer antitakeover provisions have stock prices that contain more information about future earnings. We estimate the following regression equation using a panel of annual observations aggregated up to the two-digit SIC industry level:

$$\begin{aligned} FER_{kt} = {} & c_0 + c_1 G_{k,t-1} + c_2 ROE_{k,t-1} + c_3 VROE_{k,t-1} + c_4 LEV_{k,t-1} \\ & + c_5 M/B_{k,t-1} + c_6 SIZE_{k,t-1} + c_7 DD_{k,t-1} + c_8 AGE_{k,t-1} \\ & + c_9 DIVER_{k,t-1} + \epsilon_{kt}, \end{aligned} \tag{6}$$

where $FER_{kt} = \{FERC_{kt} \vee FINC_{kt}\}$ for industry k in year t, $FERC$ is the annual future earnings response coefficient, and $FINC$ is the annual future earnings incremental explanatory power, following Durnev et al. (2003). Other regressors are two-digit SIC industry averages of the same regressors in equation (4) for idiosyncratic volatility. The dependent variables measure the amount of information about future earnings that is incorporated in current stock returns.

Columns (7) and (8) of Table V display estimates of these regressions including one-digit SIC industry fixed effects. We find that the governance index G is negatively associated with future earnings response measures. The G coefficients are -2.1178 on the $FERC$ regression and -0.1020 on the $FINC$ regression, with both coefficients statistically significant at conventional levels.

The results in this section provide further support to our earlier findings by showing how corporate governance is related to additional measures of information flow intensity. Several of the measures focus on private information in particular, supporting the proposition that private information flow is facilitated by a takeover–open governance stance.

C. Governance and Institutions: The Trading-Link Hypothesis

Institutional trading is an important channel through which information is incorporated into stock prices. Piotroski and Roulstone (2004) find that institutional trading is positively associated with idiosyncratic volatility. Hartzell and Starks (2003) find that institutional investors contribute to private information collection and trading.[12]

We introduce institutional trading as an additional control in our basic model of equation (4). This serves the purpose of verifying the robustness of the relation between governance and idiosyncratic volatility after controlling for the

[12] Other channels through which information is incorporated into stock prices include the activities of analysts and insiders. The evidence on the role of analysts is mixed, though recent evidence shows that idiosyncratic volatility is negatively related to analyst coverage (e.g., Piotroski and Roulstone (2004)). Evidence on the role of insiders shows a positive relation between insider trading and idiosyncratic volatility (e.g., Piotroski and Roulstone (2004)), but Gompers et al. (2003) find that insider trading is uncorrelated with antitakeover restrictions.

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. We use both a broad measure of institutional trading and also one focused on the trading of merger arbitrage-active institutions.

Specifically, we estimate the following regression equation:

$$\begin{aligned}\Psi_{it} = {} & c_0 + c_1 G_{i,t-1} + c_2 ROE_{i,t-1} + c_3 VROE_{i,t-1} + c_4 LEV_{i,t-1} \\ & + c_5 M/B_{i,t-1} + c_6 SIZE_{i,t-1} + c_7 DD_{i,t-1} + c_8 AGE_{i,t-1} \\ & + c_9 DIVER_{i,t-1} + c_{10} INS_{i,t-1} + c_{11} INS_{i,t-1} \times G_{i,t-1} + \epsilon_{it},\end{aligned} \tag{7}$$

where all variables are as previously defined, except $INS = \{INST \vee INSTA\}$, $INST$ is the quarterly absolute change in the number of a firm's shares held by institutional investors as a fraction of the stock's annual trading volume, and $INSTA$ is defined as the quarterly absolute change in the number of firm's shares held by merger arbitrage-active institutions 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 firm for which an acquisition offer is made during the quarter, according 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 a quarter, the institution either began the quarter owning at least 1% 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, then it increased its holdings by at least 0.5% of the shares outstanding. Second, the institution bought at least 3% of shares outstanding for at least one merger-situation firm. We experiment 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 merger and acquisition situations. The description of these regressors is summarized in Panel F of Table I, and descriptive statistics are provided in Table II.

Columns (1) and (2) of Table VI report estimates of equation (7) using $INST$, the broad measure of institutional trading, with and without the interaction regressor $INST \times G$, respectively. The estimate of the governance coefficient c_1 is strongly significantly negative in both cases. Institutional trading is associated with more idiosyncratic volatility ($c_{10} > 0$). The interaction regressor exerts a significantly negative effect when present ($c_{11} < 0$). Institutional trading therefore adds to the negative statistical effect of the governance index on volatility, in that institutional trading accelerates the incorporation of

Table VI

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. (2003) governance index. The regressors include profitability (*ROE*), profits volatility (*VROE*), leverage (*LEV*), market-to-book ratio (*M/B*), equity capitalization (*SIZE*), dividend-payer dummy (*DD*), firm age (*AGE*), and internal diversification dummy (*DIVER*). *INS* alternatively refers to: *INST*, the quarterly absolute change in the number of firm's shares held by institutions as a fraction of annual trading volume; and *INSTA*, the quarterly absolute change in the number of firm's shares held by takeover arbitrage institutions, as a fraction of annual trading volume. Refer to Table I for variable definitions. 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% levels. Robust *t*-statistics 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 × *G*		**−0.0438**		
		(−2.38)		
INSTA			**0.4360**	**0.9817**
			(5.67)	(3.89)
INSTA × *G*				**−0.0595**
				(−2.30)
R^2	7.78%	7.79%	7.73%	7.74%
N	101,169	101,169	99,075	99,075

firm-specific information into stock prices and, consequently, increases idiosyncratic volatility.

Columns (3) and (4) of Table VI report analogous results, but using *INSTA*, the targeted measure of arbitrage-active institutional trading. The relation between governance and volatility remains strong after controlling for this focused type of institutional trading. The results support our trading-link hypothesis in that the governance–idiosyncratic risk relation is stronger in the presence of high levels of arbitrage-active institutional trading at the firm level. The coefficient on the interaction variable, $INSTA \times G$, 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.

The trading link nature of a negative relation between antitakeover provisions and idiosyncratic risk is consistent with three sets of recent results. First, according to evidence in Chakravarty (2001) and Hartzell and Starks (2003), institutional investors actively collect and trade on private information, and Piotroski and Roulstone (2004) specifically show that institutional trading is positively associated with idiosyncratic volatility. Second, there is a connection both 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) document a link from disclosure quality (a different aspect of information flow) to ownership by "transient" institutions and then to total volatility. 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. Our results extend the current understanding by showing that governance and institutional trading exert mutually reinforcing, and therefore more powerful, influences on information flow. The stronger a firm's takeover defenses, the more impeded is the flow of information. The effect is accentuated when arbitrage-oriented institutions avoid the stock.

IV. Robustness

In this section, we show that our primary findings are robust to controls for accounting transparency, and also to variations in idiosyncratic volatility measurement and other aspects of our methodology. We also discuss the connection between our results and those of Jin and Myers (2005).

A. Controlling for Accounting Transparency

Accounting disclosure is a central element of information flow. Whether more transparent disclosures encourage the collection of private information or crowd it out depends on the balance of effects on the benefits and costs of acquiring private information. In our context, if intense trading on private information underlies idiosyncratic volatility, then we could observe less volatility for high-transparency stocks, since more information flows via lower-frequency accounting releases (as in Kim and Verrecchia (2001)), or more volatility as additional information collection is encouraged (as in Kim and Verrecchia (1991)).

This reasoning leads to two 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 the transparency level is directly taken into account. Therefore, we test whether G retains its negative coefficient in the idiosyncratic volatility regression once a measure of transparency is also included as a control. Second, if, as we hypothesize, the idiosyncratic volatility measure 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 either more idiosyncratic volatility than otherwise (an information encouragement effect) or less (a crowding out effect).

Our main result in this subsection is that governance remains a significant determinant of idiosyncratic volatility even after controlling for transparency. Also, firms with high levels of transparency display high levels of idiosyncratic risk, controlling for the level of governance openness—consistent with an encouragement effect.

We now describe our tests and results in more detail. To measure transparency, we follow the accounting literature (e.g., Francis et al. (2005)) and focus 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. Given the panel nature of our study, we require measures that can vary over time. Therefore, we modify the procedures of Francis et al. (2005), who work with time-series regressions, and benchmark normal accruals with annual cross-sectional industry regressions. Using the same naming convention as in Francis et al. (2005), our first measure is called $EQ2$. Nearly the same as the Teoh, Welch, and Wong (1998) measure of earnings quality, $EQ2$ is defined as the absolute value of firm-specific residuals from an industry regression of total accruals on (the reciprocal of) total assets, revenue growth, and fixed assets.[13] 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)). Both $EQ2$ and $EQ5$ are *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

[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 uses cash revenues (i.e., income statement revenues adjusted for the change in accounts receivable). The same adjustment is employed for the alternative measures. See Francis et al. (2005) for details.

observed annual figure. Panel G of Table I provides a brief description of these measures and our source data and Table II provides descriptive statistics.

In Table VII we present idiosyncratic volatility regression results controlling for accounting transparency. The table is set up similar to earlier tables, and reports variations on the same basic model in equation (4). The key difference is that an accruals-based measure of accounting quality EQ, where $EQ = \{EQ2 \vee EQ5\}$, is included as a regressor to explain relative idiosyncratic volatility Ψ. We employ $EQ2$ in columns (1) and (2), and $EQ5$ 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 EQ coefficient are negative and significant in every column (except in column (4)), implying that 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. Less accounting information apparently disproportionately inhibits efforts to collect more private information. This evidence is consistent with theoretical suggestions that high-quality disclosure can encourage the collection of private information, leading to more idiosyncratic volatility.

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 argue that the net benefit of hiding bad news from investors (which smooths returns but requires that insiders absorb bad-news costs) is smaller in well-governed firms, because insiders have limited opportunity to expropriate the proceeds of good news. Transparency prevents even poorly governed firms from hiding bad news, allowing for unimpeded volatility. Jin and Myers (2005) provide cross-country evidence that a low transparency level and poor investor protection result in low levels of relative idiosyncratic volatility (high stock return synchronization). Our firm-level results confirm 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.

Importantly, our core results and trading link rationale complement Jin and Myers (2005). We show that a particular aspect of firm-level governance, namely antitakeover provisions, is linked with the incorporation of firm-specific information into stock prices. Our thinking is that the strong investor protection implied by openness to takeovers is associated with a lower possibility of insiders (controlling shareholders and managers) expropriating outside investors. For such firms, ownership by outside investors is therefore encouraged. More such investors means more noise trading, providing more cover and profits for those willing to invest in private information collection. Thus, fewer antitakeover provisions can promote private information collection and trading by outside investors.

Table VII

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. (2003) governance index; and *GD*, which is zero if the governance index is less than or equal to 5 (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*. The regressors include profitability (*ROE*), profits volatility (*VROE*), leverage (*LEV*), market-to-book ratio (*M/B*), equity capitalization (*SIZE*), dividend-payer dummy (*DD*), firm age (*AGE*), and internal diversification dummy (*DIVER*). *EQ* is alternatively: *EQ*2, 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 *EQ*5, 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. Refer to Table I for variable definitions. 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% levels. Robust *t*-statistics 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)
*EQ*2	**−0.4632**	**−0.7336**		
	(−3.62)	(−2.29)		
*EQ*5			**−0.3207**	−0.0933
			(−3.88)	(−0.48)
R^2	7.94%	7.67%	7.73%	7.88%
N	93,118	16,676	91,964	16,719

B. Additional Robustness Checks

This subsection discusses several additional robustness checks. The robustness results appear in Table VIII. These results verify that our findings are not due to the particular model of returns used to estimate idiosyncratic volatility, to the governance openness formulations used, or to the estimation methodology used.

In particular, we estimate the basic model in equation (4) using a dependent variable based on an industry-factor model of idiosyncratic volatility, on the Fama and French (1992) three-factor model of idiosyncratic volatility, and on the French, Schwert, and Stambaugh (1987) volatility estimator, in which the additional terms adjust for biases that result from autocorrelation and cross-autocorrelations of daily returns. We use quarterly rather than monthly idiosyncratic volatility estimates, so that the dependent variable has a frequency consistent with most of our control variables. We estimate a version using only the earliest observation (1990) on the governance index present in the data set to further address endogeneity concerns. This also helps ensure that our results are not driven by multiple observations on the same firms. We also estimate our basic model using the Fama and MacBeth (1973) regression approach to be sure the results are not driven by errors-in-variables and autocorrelation. Finally, we estimate the model using a difference-in-differences approach, by including both year and firm fixed effects (Bertrand and Mullainathan (2003)), as an additional robustness check for endogeneity.

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

V. Governance and Corporate Investment Decisions

This section extends our study of governance and idiosyncratic risk to consider their joint 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 relative idiosyncratic volatilities tend to exhibit marginal Tobin's q (hereinafter, $\dot{q}$) closer to the value of unity it would display in a frictionless full-information value-maximizing 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, it is natural to expect a negative relationship between idiosyncratic volatility and $|\dot{q} - 1|$. 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. Antitakeover protections, such as those measured by the G index, are thought to free management to overinvest (Jensen (1986)), so it makes sense to consider how governance relates to such judgments. Frictions that prevent immediate full investment in favorable opportunities would also induce $\dot{q} > 1$; such frictions are likely to vary

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. (2003) governance index; and *GD*, which is zero if the governance index is less than or equal to 5 (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*. The regressors include profitability (*ROE*), profits volatility (*VROE*), leverage (*LEV*), market-to-book ratio (*M/B*), equity capitalization (*SIZE*), dividend-payer dummy (*DD*), firm age (*AGE*), and internal diversification dummy (*DIVER*). Refer to Table I for variable definitions. Regressions include two-digit SIC industry fixed effects. Financial and utilities industries are omitted (SIC 6000–6999 and 4900–4999). All variables are winsorized at the bottom and top 1% levels. Columns (1) and (2) report results using idiosyncratic volatility estimates from an industry model of returns according to the Fama and French (1997) 48 industry SIC classification scheme. Columns (3) and (4) report results using idiosyncratic volatility estimates from Fama and French (1992) 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 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 using the Fama and MacBeth (1973) estimation approach with estimates given by monthly cross-sectional firm-level regressions. Column (13) reports results including firm fixed effects and year fixed effects (difference-in-differences). Robust *t*-statistics are in parentheses. Coefficients significant at the 5% level are in boldface.

	Industry Model		Fama–French Model		Autocorrelation Correction		Quarterly		First *G*		Fama–MacBeth		Difference-in-differences
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)
G	**−0.0067**		**−0.0024**		**−0.0054**		**−0.0192**		**−0.0102**		**−0.0136**		**−0.0412**
	(−4.58)		(−2.01)		(−2.09)		(−5.39)		(−3.95)		(−6.00)		(−5.08)
GD		**−0.0727**		**−0.0692**		**−0.2124**		**−0.4058**		**−0.1717**		**−0.2253**	
		(−3.08)		(−3.67)		(−5.20)		(−7.14)		(−4.11)		(−6.95)	
ROE	0.0174	−0.0181	0.0004	0.0010	0.0336	−0.0460	0.0290	−0.1216	−0.0080	−0.1416	−0.0334	−0.1350	−0.0469
	(0.96)	(−0.40)	(0.03)	(0.03)	(1.06)	(−0.56)	(0.69)	(−0.93)	(−0.22)	(−1.54)	(−0.65)	(−1.49)	(−1.53)

(*continued*)

Table VIII—*Continued*

	Industry Model		Fama–French Model		Autocorrelation Correction		Quarterly		First *G*		Fama–MacBeth		Difference-in-differences
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)
VROE	0.0019	0.0120	0.0059	0.0118	0.0079	0.0111	0.0097	0.0063	0.0115	0.0065	−0.0028	−0.2088	−0.0159
	(0.55)	(1.67)	(1.92)	(1.55)	(1.12)	(0.95)	(1.80)	(0.43)	(1.32)	(0.57)	(−0.22)	(−0.75)	(−1.94)
LEV	**0.0653**	**0.1887**	0.0223	0.0395	**0.2060**	**0.3833**	**0.3463**	**0.4749**	**0.2581**	**0.2904**	**0.2783**	**0.3633**	−0.0260
	(2.36)	(2.78)	(1.03)	(0.78)	(4.19)	(3.22)	(5.13)	(2.81)	(4.86)	(2.08)	(4.84)	(3.53)	(−0.31)
M/B	**−0.0198**	**−0.0770**	−0.0011	**−0.0350**	**−0.0761**	**−0.1024**	**−0.0787**	**−0.1343**	−0.0216	−0.0542	**−0.0584**	−0.0419	**−0.0841**
	(−2.63)	(−4.15)	(−0.19)	(−2.39)	(−5.72)	(−3.13)	(−4.37)	(−2.86)	(−1.52)	(−1.46)	(−2.76)	(−1.13)	(−4.04)
SIZE	**−0.4079**	**−0.4124**	**−0.2025**	**−0.2000**	**−0.2177**	**−0.1998**	**−0.4989**	**−0.5006**	**−0.3832**	**−0.3841**	**−0.3951**	**−0.3981**	**−0.3604**
	(−120.78)	(−51.92)	(−77.80)	(−30.92)	(−37.00)	(−14.50)	(−60.99)	(−25.42)	(−65.22)	(−25.80)	(−18.19)	(−16.11)	(−20.59)
DD	**0.1406**	**0.2887**	**0.0769**	**0.1534**	**0.0602**	**0.2163**	**0.1471**	**0.3507**	**0.1102**	**0.2083**	**0.1971**	**0.2659**	**0.0794**
	(14.51)	(11.49)	(9.91)	(7.78)	(3.49)	(4.88)	(6.19)	(5.49)	(5.66)	(4.37)	(6.39)	(6.47)	(2.15)
AGE	**0.0510**	0.0209	**0.0104**	0.0002	**0.0411**	**0.0778**	**0.0610**	**0.1108**	**0.0770**	**0.0832**	0.0073	**0.1017**	**−0.1601**
	(8.03)	(1.43)	(2.10)	(0.02)	(3.70)	(3.02)	(4.02)	(3.08)	(5.59)	(2.53)	(0.67)	(3.66)	(−2.72)
DIVER	−0.0110	−0.0127	**−0.0533**	**−0.0592**	**−0.0787**	**−0.0802**	**−0.1134**	−0.0940	**−0.0947**	−0.0630	−0.0308	−0.0202	**−0.1680**
	(−1.31)	(−0.60)	(−8.04)	(−3.46)	(−5.28)	(−2.16)	(−5.64)	(−1.83)	(−6.11)	(−1.57)	(−1.49)	(−0.54)	(−7.62)
R^2	17.86%	19.14%	7.99%	8.82%	2.59%	2.81%	16.04%	17.78%	8.04%	8.53%			19.82%
N	119,341	21,256	119,609	21,318	115,103	20,335	40,625	7,259	90,617	16,154			119,541

according to the nature of real investment and business conditions, and can thus be proxied by industry and control variables.

We show in previous sections that corporate governance is a direct determiner of idiosyncratic volatility. This raises the 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 this section, we present tests designed to sort out these relationships.

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 for 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 cross-section. To show that this setup does not drive our results, we demonstrate that the primary findings also hold using a firm-specific $\dot{q}$.

Full details on the procedure we use to estimate $\dot{q}$ 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 subsample of firms i in each year t:

$$\Delta EV_{it} = \lambda_0 + \dot{q}\,\Delta NFA_{it} + \lambda_1 D_{it} + \lambda_2 EV_{i,t-1} + u_{it}, \tag{8}$$

where ΔEV_{it} is the 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 u_{it} is a residual term. All variables are scaled by the previous year's level of 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.

A. Industry-Level Results

Using all industry-years' estimates of $\dot{q}$, we form a panel data set on the absolute deviations of the marginal q from one, $|\dot{q} - 1|$. This quantity is the subject of analysis by estimating the following annual regression equation:

$$\begin{aligned} |\dot{q} - 1|_{kt} = {} & \alpha_0 + \alpha_1 \Psi_{k,t-1} + \alpha_2 G_{k,t-1} + \alpha_3 LEV_{k,t-1} \\ & + \alpha_4 M/B_{k,t-1} + \alpha_5 SIZE_{k,t-1} + \alpha_6 DIVER_{k,t-1} + \nu_{kt}, \end{aligned} \tag{9}$$

where regressors are two-digit SIC industry averages. In view of the fact that industry $\dot{q}$ is effectively weighted by asset value, we also use weighted averages for the control variables. Panel A of Table IX presents the results of estimating this panel regression where $|\dot{q} - 1|$ is the dependent variable.[14]

The results show that $|\dot{q} - 1|$ is negatively related to relative idiosyncratic volatility Ψ (column 1), consistent with the result in Durnev et al. (2004).

[14] We check that similar results hold using the log absolute deviation as the dependent variable, which should better conform to OLS regression assumptions. Our results are also robust to the inclusion of one-digit SIC industry and year fixed effects.

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 (two-digit SIC) 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 industry marginal Tobin's q relative to one (see the Appendix for full details). Regressors are two-digit SIC industry averages. Ψ is the logistic transformed relative idiosyncratic volatility. G is the IRRC–Gompers et al. (2003) governance index. The regressors include profitability (*ROE*), profits volatility (*VROE*), leverage (*LEV*), market-to-book ratio (*M/B*), equity capitalization (*SIZE*), dividend-payer dummy (*DD*), firm age (*AGE*), and internal diversification dummy (*DIVER*). 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. Panel B reports estimates of coefficients of the annual time-series cross-sectional firm-level regressions of firm i in each year t

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

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

where ΔEV_{it} is the change in enterprise value, ΔNFA_{it} is the change in net fixed assets, I_{it}^k refers to an indicator variable for firm membership in the k^th industry, and D_{it} 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. Robust t-statistics 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	**−0.2923**		**−0.2979**	−0.1065	0.3524	
	(2.17)		(2.16)	(0.51)	(1.35)	
G		−0.1012	−0.1039	**−0.1878**	−0.1552	
		(1.86)	(1.92)	(2.18)	(1.43)	
$\Psi^{predicted}$						1.8675
						(1.69)
$\Psi^{residual}$						**−0.2979**
						(2.16)
LEV	**−1.8163**	**−2.7243**	**−2.2528**	−1.9109	0.8788	**−2.2528**
	(2.25)	(3.27)	(2.65)	(1.35)	(0.48)	(2.65)
M/B	0.2341	0.1526	0.1952	−0.1507	−0.5886	0.1952
	(1.20)	(0.78)	(1.01)	(0.55)	(1.71)	(1.01)
SIZE	−0.1446	−0.0927	−0.1465	0.1154	**0.4622**	−0.1465
	(1.47)	(0.95)	(1.49)	(0.79	(2.52)	(0.149)
DIVER	−0.2520	−0.0593	−0.0928	−0.4887	−0.7099	−0.0928
	(0.85)	(0.22)	(0.35)	(1.12)	(1.34)	(0.35)
Constant	**4.3614**	**4.2127**	**5.3905**	2.6402	−3.4653	1.0513
	(3.09)	(2.91)	(3.48)	(1.13)	(1.24)	(0.49)
R^2 (pseudo-R^2 for tobits)	6.03%	5.52%	7.35%	13.0%	24.6%	7.35%
N	236	236	236	236	236	236

(*continued*)

Table IX—*Continued*

Panel B: Firm-Level Regressions			
	ΔEV	$\ln(u^2)$	$\ln(u^2)$
$\dot{q}^{overall}$	1.3254		
	(3.42)		
Ψ		**−0.1083**	
		(2.32)	
G		**−0.0835**	
		(7.73)	
$\Psi^{predicted}$			−0.0715
			(0.79)
$\Psi^{residual}$			**−0.1127**
			(2.10)
LEV		**−1.5545**	**−1.6150**
		(6.65)	(6.86)
M/B		**1.2139**	**1.2490**
		(20.22)	(20.72)
SIZE		**−0.1643**	**0.1904**
		(5.73)	(6.50)
DIVER		−0.0645	**−0.1228**
		(1.01)	(1.92)
λ_1	−2.140		
	(0.91)		
λ_2	0.0191		
	(0.12)		
Constant		0.2266	−0.2171
		(0.48)	(0.46)
R^2	2.29%	46.2%	46.0%
N	6,790	6,790	6,790

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). 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 results indicate that the positive relation between takeover protections and corporate investment efficiency is due to firms with $\dot{q} > 1$, that is, firms that might be underinvesting. The negative relation between $|\dot{q} - 1|$ and G is evident only for high-$\dot{q}$ observations. The implication is that antitakeover protections are associated with a smaller (but still positive) degree of underinvestment.

The importance of idiosyncratic volatility in reducing $\left|\dot{q} - 1\right|$ 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 the predicted value is significantly positive at the 10% level, and the coefficient on the residual is strongly 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 confirm 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, Morck, and Yeung (2004) explanation 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 confirm that our central conclusions 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 marginal 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

$$\Delta EV_{it} = \sum_{k=1}^{K} \lambda_0^k I_{it}^k + \dot{q}^{\text{overall}} \Delta NFA_{it} + \sum_{k=1}^{K} \dot{q}^k I_{it}^k \Delta NFA_{it} + \lambda_1 D_{it} + \lambda_2 EV_{i,t-1} + u_{it}, \tag{10}$$

$$\ln\left(u_{it}^2\right) = \alpha_0 + \alpha_1 \Psi_{it} + \alpha_2 G_{it} + \alpha_3 LEV_{it} + \alpha_4 M/B_{it} + \alpha_5 SIZE_{it} + \alpha_6 DIVER_{it} + \nu_{it}, \tag{11}$$

where k indexes industries and the variables are as previously defined, except I_{it}^k refers to an indicator variable for firm-i membership in the k^{th} industry in year t, and $\dot{q}^k$ captures the industry-specific deviation from the full-sample estimate of $\dot{q}^{overall}$. In this system, we posit that the squared residual term u^2 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, 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 in the industry-level analysis).

Panel B of Table IX presents OLS estimates of this system; to save space, we suppress coefficients on industry dummies and 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 that the estimate of $\dot{q}^{overall}$ is near unity (1.325), and that it is strongly statistically significant. Though not reported in the table, the null hypothesis that $\dot{q}^{overall} = 1.0$ is not rejected.

Columns (2) and (3) contain estimates of the determinants of the logarithm of the squared residual. In this specification, the 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 intraindustry 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 indicate a firm $\dot{q}$ that is more 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 a decomposition of idiosyncratic volatility similar to those used in the previous subsection. To compute these results, we again follow Almeida and Wolfenzon (2005) and decompose relative idiosyncratic volatility Ψ into a component linearly related to the governance index G and a residual component. As before, we refer to these components as the predicted value and the residual value of Ψ, denoted $\Psi^{predicted}$ and $\Psi^{residual}$, respectively. Coefficient estimates imply that the residual value portion of Ψ is associated with an attenuated level of firm-specific $\dot{q}$ deviation from the industry level. This result is consistent with the findings in the industry-level subsection.[15]

In summary, Table IX establishes two main results. First, idiosyncratic volatility is a driver of marginal Tobin's q, controlling for any additional influence of antitakeover governance provisions. This result strengthens the 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 ($\dot{q} > 1$). From the perspective of an investor, limited information flow allows for both corporate overinvestment and underinvestment in that value is delinked from corporate decisions. At the same time, takeover

[15] Note that the coefficients 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.

protections that discourage the flow of information also encourage managerial spending, and actually discourage underinvestment. Overall, it is the information flow rather than governance that exerts the dominant influence.

VI. Conclusion

We find that idiosyncratic risk is decreasing in a firm's degree of insulation from takeovers. 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. Further, openness to the market for control is linked to information flow in a way not captured by the openness of a firm's financial reporting.

In support of an informational interpretation of our primary 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 fewer antitakeover provisions incorporate more information about future earnings.

We also show that an institutional trading link is one mechanism for the relationship from governance to idiosyncratic risk. That is, the governance–risk relationship is more pronounced for firms that are subject to intense trading by institutional investors, and particularly for those that have recently been involved in risk arbitrage around mergers. Thus, at least one of the links between governance and idiosyncratic volatility and between governance and information flow occurs through arbitrageur institutions.

Finally, we establish the connections among corporate governance characteristics, idiosyncratic risk, and corporate investment decision making quality. Like others, we find a positive correlation between idiosyncratic risk and decision-making quality. Upon decomposing volatility into governance-related and nongovernance-related components, we find that it is mainly nongovernance-related idiosyncratic volatility that is associated with the quality of investment decision making. Our finding means that it is information flow more than governance that is important for this practical business outcome, which further substantiates an information-flow interpretation of idiosyncratic risk.

The finance literature emphasizes that governance influences management actions. Our contribution is to develop evidence on one specific way in which governance influences the actions of outside investors. Specifically, we establish a link between governance structures and stock price efficiency by showing that takeover restrictions impede information flow to stock prices. Moreover, openness to the market for corporate control and informed trading by institutions interact to influence the extent to which stock prices incorporate information in an accurate and timely fashion. We recommend further exploration of how different types of institutions interact with the market for corporate control to influence stock price efficiency and other economic outcomes.

Appendix

To obtain estimates of marginal Tobin's q, we work with a 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 $EV = NFA + GCV$. The change (Δ) over time in enterprise value is $\Delta EV = \Delta NFA + \Delta GCV$, provided that outside investors do not contribute additional capital nor withdraw any value; we add consideration of such complications below. Tobin's q relates the firm's market value to the replacement value of its physical assets. 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, that is, $\dot{q} \equiv \frac{\Delta EV}{\Delta NFA}$. This definition suggests that $\dot{q}$ can be estimated by regressing the change in enterprise value (ΔEV, a market value concept) on the change in physical assets (ΔNFA, a replacement value concept).

To implement this idea, one must allow for expected changes in NFA and EV, including depreciation, for disbursements of value to claimants via dividends and repurchases (D), and for disbursements' tax effects. In the disbursements measure, interest payments can be ignored if we treat debt as effectively perpetual. In this context, Durnev et al. (2004) propose an augmented regression equation to estimate $\dot{q}$. In our notation, their regression is

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

where i indexes firms and t indexes years. 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 firm. We scale regressors by the previous year's assets.

Durnev et al. (2004) obtain time-invariant estimates $\dot{q}$ by estimating this regression for each three-digit SIC industry using recursive procedures to adjust book values to replacement values. For our purposes, we require panel data on $\dot{q}$ over the 1990s. For comparison to our other results, we also require some information on $\dot{q}$ at the firm level. With these goals in mind, we modify the Durnev et al. (2004) procedures somewhat. The need for a panel precludes the use of recursions, so we rely more on book values and use data from statements of cash flows wherever possible, since these are inherently at current values, and factor in consideration inventory valuation methods. To obtain a larger 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.

Regressors for each firm-year are computed from COMPUSTAT annual data items as follows: $EV =$ Item 25 $\times$ Item 199 + Item 56 + Item 9 + Item 34 $-$ Item 4, to be consistent with the treatment of current assets in Durnev, Morck, and Yeung (2004); $\Delta NFA =$ Item 128 + Item 129 + ΔItem 240 $-$ Item 303, if Item 303 is missing then we add ΔItem 3 instead; and $D =$ Item 127 + Item 115.

We require 20 firms with complete data in each industry-year in order to estimate equation (A1). 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% in absolute value, and we eliminate industry-years with $\dot{q}$ estimates of more than six in absolute value.

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