# Takeovers and the Cross-Section of Returns

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This paper considers the impact of the takeover likelihood on firm valuation. If firms are more likely to acquire when there is more free cash or lower required rates of return, the targets become more sensitive to shocks to cash flows or the price of risk. *Ceteris paribus*, firms exposed to takeovers have different rates of return than protected firms. Using takeover likelihood estimates, we create a "takeover factor," buying (selling) firms with a high (low) takeover likelihood, which generates "abnormal" returns. Several tests confirm that the takeover factor helps explaining cross-sectional differences in equity returns and is related to takeover activity.

This paper considers the impact of the takeover channel on valuation. While it is well known that target shareholders receive a large premium on a takeover, how expectations about takeover premiums affect firm valuation has not been investigated. One possible reason for this lack of interest may be the assumption that differences in takeover exposure are purely idiosyncratic, and hence do not affect a firm's cost of capital. In that case, the issue of incorporating the takeover channel into valuation is solved by simply adding the expected takeover premium to the expected cash flows. However, takeover activity, and hence a target's exposure might not be idiosyncratic.

In particular, Bruner (2004) and Rhodes-Kropf, Robinson, and Viswanathan (2005) show that takeover activity is time varying and related to the conditions in the equity market. Further, a systematic exposure to takeovers can have an important impact on firm valuations and returns, as the median bid premium—approximately 35%—and takeover activity—3,467 completed

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deals between 1980 and 1998—are both high (Andrade, Mitchell and Stafford, 2001).[1]

In this paper, we first provide a simple theoretical framework that uses an asset pricing model to value firms that differ in their takeover exposure. A central feature of the asset pricing model is time variation in the price of risk, which is assumed to be imperfectly correlated to changes in aggregate fundamentals (i.e., similar to Campbell and Vuolteenaho, 2004 (henceforth CV); and Lettau and Wachter, 2007).

In this framework, we consider two alternative motivations for acquisition activity.

The first motivation for acquisitions is driven through agency problems on the acquirer's part. These agency problems lead to empire building, which is exacerbated during times of positive cash-flow shocks (the "agency" view, with more acquisitions if fundamentals are good). This would explain the relation between takeover activity and market conditions and would cause firms exposed to takeovers to become more sensitive to shocks in aggregate fundamentals (i.e., cash-flow shocks). The second motivation for acquisitions is the valuation of potential synergies (the "synergy" view).[2] When the price of risk is low, the value of these synergies is high and firms tend to acquire, thereby increasing the sensitivity of potential targets to the changes in the price of risk (i.e., discount-rate shocks).

Within our model, we incorporate these two takeover motivations in separate scenarios. Both motivations imply that differences in the takeover likelihood lead to differences in exposure to state variables determining asset prices, and hence to differences in the expected rate of return. However, whether firms exposed to takeovers have a higher or a lower rate of return depends on the relative importance of the two acquisition motives. The "agency" view would unambiguously suggest that firms exposed to takeovers should have a higher rate of return: takeover premiums arrive when aggregate fundamentals are high, thus when investors least need the cash. The implications from the "synergy" view (i.e., of receiving the takeover premium when the price of risk is low or when future expected returns are low) depend on the importance of the investor's intertemporal hedging demands (see Merton, 1973). If such demands are important, investors strongly value receiving the takeover premiums at a



[1] There were 1,427 completed deals between 1980 and 1989 and 2,040 completed deals between 1990 and 1998. The median bid premium received by targets was 37.7% in the 1980s and 34.5% in the 1990s. Further, acquisition activity increased in 1999 and 2000 before dropping in 2001.

[2] This is similar in spirit to the *Q*-theory of investments (Abel, 1983, see also Jovanovic and Rosseau, 2002). Recently, other theories have been proposed to explain the time variation in takeover activity relying on misvaluation in capital markets (see Shleifer and Vishny, 2003; and Rhodes-Kropf and Viswanathan, 2004). Under certain conditions, to be discussed in Section 1, the use of such misvaluation theories to explain time-varying takeover activity does not affect the interpretation of our results.

time when future returns are low. In this case, the synergy view would suggest that firms exposed to takeovers should have a lower rate of return.[3]

Next, we document six empirical results to shed light on these implications. First, we construct a quintile-spread portfolio that buys firms with a high takeover vulnerability estimated using a logit regression, and sells firms with a low takeover vulnerability. This long-short portfolio is associated with annualized abnormal returns of 11.77% relative to the four-factor Fama-French (1992), and Carhart (1997) model in our time period from 1981 to 2004. These results are confirmed using 10-year rolling estimation windows for the logit estimation as well. These findings suggest that a higher exposure to takeovers leads to higher expected returns, supporting the agency view. Also, this would imply that the four-factor model does not fully account for state variables that are associated with time-varying risk premiums.

Second, the takeover-spread portfolio denoted the "takeover" factor, proxying for the risk due to stock-price sensitivity to state variables affecting time variation in risk premiums. We find that our proposed factor and the differences in the takeover likelihood across its quintile-spread portfolios seem to predict the real takeover activity.

Third, we verify that our takeover factor is indeed related to the takeover vulnerability rather than more general exposure to business cycles, by considering changes in a firm's takeover beta before and after the adoption of state antitakeover legislation in the state in which the firm is incorporated. As predicted by our model, takeover betas decrease after states adopt such legislation and firms experience an exogenous shock that decreases their exposure to takeovers.

Fourth, the takeover factor explains differences in the cross-section of equity returns. Our main results are for the cross-section of stocks, sorted into size and book-to-market portfolios, for which it is striking that the takeover factor can significantly improve the asset pricing model beyond the size and book-to-market factors.[4] In particular, adding the takeover factor to the four-factor model almost doubles the $R^2$ using the 100 size and book-to-market sorted portfolios and improves pricing performance as well. Further, this improvement in cross-sectional pricing is not limited to the extreme portfolios of high growth and/or small-size stocks, and is robust to using the rolling 10-year estimation windows, adding average portfolio characteristics, and using a different set of test portfolios and a different time period.

Fifth, we investigate the link between corporate governance and stock returns as documented in Cremers and Nair (2005, henceforth CN), and Gompers, Ishii, and Metrick (2003, henceforth GIM). While corporate governance and takeover

[3] It also follows, perhaps counter intuitively, that despite a potentially higher required rate of return, firms with a greater takeover exposure are also valued higher. This is due to the expected takeover premium, which is absent for a firm that is protected from takeovers.

[4] See Ferson, Sarkissian, and Simin (1999) on how a factor based on an anomaly can be expected to price a cross-section of equity returns sorted on the same dimension that created the anomaly.

activity are clearly related, many corporate governance issues are not directly related to takeovers, while takeovers can occur for reasons beyond governance (such as synergies). Here, we try to disentangle the return results in GIM and CN. GIM employ a governance (G) index they develop to show that a portfolio that buys firms with the highest level of shareholder rights and that sells firms with the lowest level of shareholder rights generates an annualized abnormal return of 8.5% in a sample from 1990 to 1999. CN investigate how different governance mechanisms interact and show that these abnormal returns exist (and are higher) only when the G index is complemented with the presence of a blockholder (or a high public pension fund ownership).[5] In this paper, we check if these abnormal returns decrease when the asset pricing model incorporates the takeover factor, noting that the takeover factor has a low correlation with the governance-spread portfolios. Specifically, we show that the abnormal returns associated with governance-spread portfolios (as used in GIM and CN) decrease significantly once we add the takeover factor to the asset pricing model that includes the Fama-French factors and the momentum factor. Thus, it appears that the asset pricing model employed in these earlier papers is incomplete and that their results are driven by corporate governance provisions that are takeover related.

Sixth and finally, using the two-beta model proposed by CV, we show that firms exposed to takeovers indeed have higher cash-flow betas, suggesting that takeover activity is indeed more likely to be related to changes in aggregate fundamentals rather than the price of risk, which is consistent with a higher expected return.

The central idea in this paper—that firms differing in takeover exposure also differ in their exposure to state variables that are important for asset prices—contributes to another area of active research. In particular, this paper contributes to the empirical asset pricing literature that uses factors other than the market factor to capture time variation in risk premiums. While an intertemporal capital asset pricing model was proposed as early as 1973 (Merton, 1973), empirical work to detect stochastic variation in investment opportunities, with the notable exception of Campbell (1991), has only been recent (e.g., Brennan, Wang and Xia, 2004).[6] This paper proposes to use the takeover likelihood as a proxy for a firm's exposure to these (unobservable) state variables. Thereby, we also investigate if the empirically successful Fama-French model completely accounts for such time variation in investment opportunities, which does not appear to be the case.

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[5] Bebchuk, Cohen, and Ferrell (2004) confirm the result in GIM using a narrower index using 6 (out of 24) provisions in the original index compiled by GIM.

[6] Brennan, Wang, and Xia (2004) note that

> However, despite this evidence of time variation in investment opportunities, and despite the lack of empirical success of the classic single period CAPM and its consumption variant, there has been little effort to test models based on Merton's classic framework.

Our results also imply that the benefits of corporate governance should not be inferred from the abnormal returns (relative to the Fama-French model) that GIM and CN document. It might indeed be true that better governance is beneficial, as suggested by the association between better governance with higher valuations and better operating performance (see GIM and CN). However, the results in this paper point out that the abnormal returns accruing to stronger governance are consistent with those firms having higher systematic risk, which is not fully captured by the Fama-French asset pricing model. Therefore, using these abnormal returns to advocate the case of stronger corporate governance could be misleading.

In the next section, we present a simple theoretical framework to highlight the main idea in this paper. In Section 2, we estimate a logit model to form portfolios based on different levels of takeover vulnerabilities and investigate their returns, including their association with takeover activity in the economy. In Section 3, we confirm that our logit model and the resulting takeover factor indeed capture cross-sectional differences in takeover vulnerability. Section 4 investigates the ability of the takeover factor to explain differences in the cross-section of equity returns and whether the takeover factor is related to the cash flow and discount rate as indicated by the two-beta model proposed by CV. Section 5 concludes.

## 1. Takeovers and Asset Prices

We specify a parsimonious environment that allows us to focus on differences in valuation arising from differences in takeover vulnerability. We categorize firms into potential acquirers and potential targets. All potential targets have identical final cash flows of $X_T$ that, for simplicity, are realized without any uncertainty. At time $t + k < T$, an acquirer can attempt an acquisition that pays the target a premium of $\Delta$ over the stock price, where $\Delta$ is a stochastic variable. In the two motivations for takeovers developed below, the takeover premium $\Delta$ can be either driven by the cash available (in the "agency" view) or by price of risk (in the "synergy" view). The targets differ in the level of managerial entrenchment that changes the likelihood with which a takeover bid succeeds or would occur in the first place.[7] The parameter $\tau$ reflects the likelihood with which a takeover bid succeeds. A lower value of $\tau$ hence reflects greater managerial entrenchment in the target firm.[8]

To value potential targets, we appeal to a well-known existence theorem (Harrison and Kreps, 1979). This theorem states that, in the absence of arbitrage, there exists a stochastic discount factor or pricing kernel, $M_T$, so that the price

[7] Examples of managerial entrenchment devices include takeover defenses and leverage (Stulz, 1988, and Harris and Raviv, 1988).

[8] The managers can differ in their private benefits, based on whether or not they follow entrenchment strategies. That is, managers with higher private benefits are more likely to be entrenched.

at time $t$ for any traded asset paying $X_T$ at time $T > t$ equals

$$P_t = P_t(X_T) = E_t[M_T X_T], \tag{1}$$

where $E_t$ denotes the expectation conditional on information available at time $t$. The price of the potential targets at time $t$ is then

$$E_t[P_{t+k} + \tau\Delta]\ E_t[M_{t+k}] + \mathrm{cov}_t(P_{t+k}, M_{t+k}) + \tau\ \mathrm{cov}_t(\Delta, M_{t+k}), \tag{2}$$

where $P_{t+k}$ is the present value at time $t + k$ of receiving $X_T$ at time $T$.

The covariance between the stochastic discount factor and the expected premium in the above expression leads to differences in expected returns between firms that have a different takeover exposure ($\tau$). The rest of the framework presents two potential reasons as to why this last covariance term might be different from 0. To do so, we first present a reduced-form linear characterization of the stochastic discount factor that depends on two parameters. We then present the two motivations for takeover activity that generate a link between takeovers and these asset pricing parameters.

### 1.1 Asset pricing

The asset pricing model we employ has the important feature that the price of risk varies, implying that at some times investors require a greater return per unit of risk than at others. This assumption is substantiated by a large and growing body of empirical work on the predictability of expected excess returns on aggregate stock market index (e.g., Shiller, 1984; Campbell and Shiller, 1988; Fama and French, 1988, 1989; Campbell, 1991; Hodrick, 1992; Lamont, 1998; and Lettau and Ludvigson, 2001). To capture this time-varying risk premium, we introduce a state variable, $z_t$, which follows the process

$$z_{t+1} = z_t + \sigma_z\, \varepsilon_{z,\,t+1}, \tag{3}$$

where $\varepsilon_z$ is a shock to the price of risk, distributed normally with zero mean and unit standard deviation. We do not take a stand on the source of this state variable and, consequently, do not take a stand on the relative merits between the various models that generate such time-varying risk premiums.

We assume that the shocks to $z$ are not perfectly linked to any variation in aggregate fundamentals. This makes our model similar to, among others, the model used in CV and Lettau and Wachter (2005). For simplicity, we assume that the shocks to $z$ are independent of the variation in aggregate fundamentals. The aggregate fundamentals are modeled as follows. We denote the log of aggregate payout to stockholders in the economy at time $t$ by $d_t$ and use a simple model of payout growth that follows the process:[9]

$$d_{t+1} = d_t + \sigma_d\, \varepsilon_{d,t+1}, \tag{4}$$

[9] This can be viewed as a simplified version of the dividend growth model used, for example, by Campbell (1991), Bansal and Yaron (2004), and Lettau and Wachter (2005).

where $\varepsilon_d$ is a shock to the payout growth and is distributed normally with zero mean and unit standard deviation.

The discount factor captures these two mentioned sources of variation through factors that are related to time-varying risk and to aggregate fundamentals. Since a stochastic discount factor can be linearly approximated by a Taylor expansion, we can express the price of a security that pays $X_T$ at time $T$ as

$$P_t(X_T) = E_t(M)E_t(X_T) + b\ \mathrm{cov}_t(-Z_T, X_T) + c\ \mathrm{cov}_t(D_T, X_T), \quad (5)$$

where $Z$ is a factor capturing shocks in the price of risk and $D$ is a factor capturing dividend or cash-flow shocks.[10]

Stocks whose payouts $X$ are positively correlated with aggregate cash-flow shocks $D$ pay off when aggregate fundamentals are high. Because these stocks distribute cash when investors least need it, investors will demand to receive a higher return on these stocks. Therefore, the parameter "$c$" should be negative. Whether parameter "$b$" is positive or negative depends on the importance of intertemporal hedging demands. In the absence of any intertemporal hedging concerns, investors demand a higher return on stocks that pay off when current valuations are high. Thus, investors demand a higher return on stocks whose returns covary negatively with the price of risk, implying that "$b$" should be negative as well. However, if intertemporal hedging concerns are important, such stocks also provide hedging benefits, by paying off when future expected returns will be low. This would lead to lower expected returns and a less-negative (or even positive) value of $b$ (see also CV).

## 1.2 Takeover activity

We consider two alternative motivations driving acquisition activity and investigate their implications for expected returns.[11]

**1.2.1 Agency problems.** How do returns to takeover targets vary if acquisitions are driven by agency problems that emanate from the separation of ownership and control? In the spirit of Jensen (1986) and, more recently, Dow, Gorton, and Krishnamurthy (2005), we characterize the agency problem by the assumption that managers of acquiring firms do not pay out cash directly to shareholders but instead use it to invest in acquisitions and other projects. These managers thus have "empire building" tendencies, which are easier to

[10] For an illustration of the linearization of the stochastic discount factor, consider the Campbell-Cochrane (1999) model. Although variation in aggregate fundamentals and the price of risk are closely linked in Campbell and Cochrane (1999), the discount factor—given by $M_{t,t+k} = \{(S_{t+k}C_{t+k})/(S_tC_t)\}^{-\gamma}$, where $C$ denotes the consumption and $S$ denotes the consumption surplus ratio—is approximately equal to $M_{t,t+k} = 1 - \gamma\ (S_{t+k} - S_t)/S_t - \gamma\ (C_{t+k} - C_t)/C_t$.

[11] To the extent that takeovers only occur if the premium is above a threshold level, aggregate merger activity will be related to stock market conditions. However, in our parsimonious model, we allow takeovers to occur regardless of the premium and focus instead on how the premium varies over time.

pursue when the financial constraints the firm faces are lower (i.e., when the amount of cash in the firm increases).[12] As a result, the cost of acquiring is a decreasing function of the firm's free cash flow.

The managers of potential targets, on the other hand, pay out cash directly to shareholders. Thus, the channel through which shocks to a firm's cash flows are transferred as shocks to the aggregate payout (dividends versus takeover premiums) depends on the fraction of acquirers in the economy.

Having already characterized the payout growth process, the cash held by acquirers at time $t + 1$ is then

$$c_{t+1} = a\ \sigma_d\ \varepsilon_{d,t+1}, \tag{6}$$

where $a$ denotes the fraction of firms in the economy that are acquirers.

Since acquisitions are easier when acquirers have more cash available, the premium the acquirer offers is a function of the cash on hand, and is denoted by $\Delta(c_{t+1})$. This directly relates the takeover premium to the aggregate cash flow shocks in the stochastic discount factor. Consequently, takeover vulnerability will affect the rate of equity return. Using the specification of the takeover premium in Equation (2), we get the following proposition:

**Proposition 1.** *Firms with a greater exposure to takeovers have a higher expected rate of return due to a higher exposure to factors related to aggregate fundamentals. At the same time, firms with a higher exposure to takeovers, ceteris paribus, have a higher value.*

*Proof.* The value of a potential target firm can be written as

$$E_t[P_{t+k}\, M_{t+k}] + \tau\ E_t[\Delta, M_{t+k}] = E_t[P_{t+k}]\ E_t[M_{t+k}] + \tau\ E_t[\Delta]\ E[M_{t+k}] \\ +\mathrm{cov}_t(P_{t+k}, M_{t+k}) + \tau\ \mathrm{cov}_t(\Delta, M_{t+k}),$$

where the value of a firm completely protected from takeovers equals $E_t[P_{t+k}\ \mathrm{M}_{t+k}]$. The takeover premium $\Delta$ is a function of the shock to the acquirer's cash only, so the covariance between $M_{t+k}$ and $\Delta$ is given by $\mathrm{cov}_t(D_{t+k},\Delta)$. Since the premium increases as shocks to cash increase, using Equations (4) and (6), this covariance term is positive. Thus, the firms expected return increases in takeover vulnerability, where the higher return is only due to a higher beta on the factor related to aggregate fundamentals. Finally, $\tau\ E_t[\Delta, M_{t+k}] > 0$, so that higher takeover exposure is associated with a higher value. □

**1.2.2 Synergies.** This section considers the potential to generate synergies as an alternative motivation for acquisitions. These synergies are captured through an increase in the target's cash flow, from $X_T$ to $X_T(1 + \psi)$, after the

[12] Viewed literally, this motivation would only explain cash deals. However, managers can also use a combination of stock and cash, where it can be easier for the manager to pursue individual private benefits when the cash component is higher. One could also incorporate stock deals in an alternative view whereby stock issuance today for acquisition purposes leads to stronger financial constraints in the future. A manager with cash in hand would be less concerned about this cost.

acquisition. Thus, $\psi X_T$ denotes the potential synergies that can be attained by the combination of the two firms and which is uncertain. The perceived synergy is shared between the target, who receives a takeover premium $\Delta$ and the acquirer.[13] Since a large body of evidence on share-price reactions around takeover announcements suggests that on an average, targets receive a positive premium while acquirer returns are insignificantly different from zero, we attribute all synergies to the target, so that $\Delta = P_{t+k}\ \psi$.[14]

In this setting, the present value of the expected synergies increases as the future cost of capital decreases. These increases allow an acquirer to pay a higher takeover premium. More generally, the takeover premium is a function of the future price of risk and is denoted by $\Delta(z_{t+k})$. As a result, once again the takeover premium is related to the stochastic discount factor, this time through shocks to the price of risk. Applying this to Equation (2), we get the following proposition:

**Proposition 2.** *Firms with a greater exposure to takeovers have a greater exposure to state-specific risk factors that affect time-varying risk premiums than similar firms that are protected from takeovers. If intertemporal hedging demands are important, then firms exposed to takeovers would have a lower rate of return.*

*Proof.* The value of the firm exposed to takeovers can be written as

$$E_t[P_{t+k}M_{t+k}] + \tau\ E_t[\Delta]\ E[M_{t+k}] + \tau\ \mathrm{cov}_t(\Delta,\ M_{t+k}),$$

where the value of the firm protected from takeovers equals $E_t[P_{t+k}\ M_{t+k}]$. As the takeover premium is a function of shocks to the price of risk only, the covariance between $M_{t+k}$ and $\Delta$ is given by $\mathrm{cov}_t(-Z_{t+k},\Delta)$. Because the takeover premium increases as the price of risk decreases, this covariance term is positive. Thus, for the firm exposed to takeovers, the exposure to $Z$ is given by $b[\mathrm{cov}_t(P_{t+k},-Z_{t+k}) + \tau\ \mathrm{cov}_t(-Z_{t+k},\Delta)]$, which is increasing in $\tau$. In the presence of intertemporal hedging demands, $b$ can be positive, and hence the rate of returns to firms exposed to takeovers can be lower than similar firms that are protected from takeovers.

### 1.3 Discussion

Propositions 1 and 2 illustrate that takeover vulnerability can affect the expected rate of return. If firms are more likely to acquire when they have free cash or when the required rate of return is low, takeover targets become more sensitive to aggregate cash-flow shocks or to the price of risk. In our model, this effect on expected returns arises because the takeover premium depends on the two state variables, the amount of cash available and the price of risk, which determine time variation in the risk premiums.

[13] The acquirer management might also receive private benefits from the acquisition, such as those attributed with empire building (Jensen, 1986).

[14] See Bruner (2004) for a comprehensive survey.

Takeover vulnerability can either increase or decrease the rate of return, depending on the motives that drive acquisition activity. First, if agency motives are more important, we would expect to find higher expected returns for firms with a greater takeover vulnerability. In this case, takeovers would be more likely if acquirers have more cash, and stocks whose payouts are positively correlated with aggregate cash flows have higher required rates of returns. Second, if synergy motives are more important and intertemporal hedging demands are sufficiently large, we could expect to find lower expected returns for firms that are more likely takeover targets. In this case, if the price of risk is lower or future expected returns are lower, synergies are more valuable and thus the takeover premium is higher. Large hedging demands imply that investors would be willing to accept lower rates of returns on stocks that pay out when future rates of returns are low.

Next, we turn to the data and use the four-factor asset pricing model proposed by Fama-French (1992), and Carhart (1997) to empirically explore the association between the takeover likelihood and rates of return.

## 2. Takeover-Spread Portfolios

We first investigate if firm-specific differences in takeover exposure are related to differences in their equity returns. To this end, we form portfolios based on the takeover vulnerability of each firm, and estimate abnormal returns relative to the four-factor model.

### 2.1 Takeover vulnerability

The likelihood that a firm will be acquired is estimated by a logit regression. Acquisitions are identified from the Securities Data Corporation's (SDC) database. We consider both all announced and completed takeovers, or 100% completed takeovers only, and include both friendly and hostile bids. The number of takeover targets in our sample with full firm-level information from Compustat between 1981 and 2004 equals 5,457 using all announced and completed takeovers, and equals 2,813 targets using 100% completed takeovers only. If we only consider a much smaller sample of firms covered by the Investor Responsibility Research Center (IRRC) database with governance information available (as explained further in more detail), the number of targets equals 799 for all announced and completed takeovers versus 412 for completed takeovers.

Our first set of tests concern the probability of a takeover occurring in the next year. In the logit model, the "target dummy" is the dependent variable, and takes the value 1 if a firm is a target in that year. The logit model incorporates a number of independent variables that have been used in the prior literature seeking to explain the probability of takeovers (see, for example, Hasbrouck, 1985; Palepu, 1986; and Ambrose and Megginson, 1992). These variables include an industry dummy that measures whether a takeover attempt occurred in the same industry in the year prior to the acquisition, the return on assets of the firm (ROA), firm leverage (book debt to assets ratio), cash (the cash and short-term investments to assets ratio), firm size (ln(Mktcap)), or the log of market equity), *Q* (market/book ratio of the firm value), and asset structure (PPE, measured by the property, plant, and equipment to assets ratio). All of these independent variables are measured at the end of the previous fiscal year.

In addition, we also include a variable to indicate the presence of a large external shareholder, as it has been argued that takeovers are more likely to occur as shareholder control increases (Shleifer and Vishny, 1986). We proxy external blockholders by those institutional shareholders that have more than a 5% ownership stake in the firm's outstanding shares. To construct this measure, we use data on institutional share holdings from Thompson/CDA Spectrum, which collects quarterly information from SEC 13f filings. We use a dummy variable, denoted by *BLOCK*, which takes the value 1 when an institutional blockholder exists at the end of the previous year and 0 otherwise.

Panel A of Table 1 presents the mean values of these variables for the entire Compustat universe over 1981–2004 for which there are no missing data, separating targets from all other firms. We also test whether the means of the target group are different. For the sample of all announced and completed takeovers, all variables except asset structure, cash, and size are significantly different for the target group. For the sample of completed takeovers only, all variables except asset structure, leverage, and ROA are significantly different.

We also consider a much smaller sample used in earlier governance studies documenting a link between governance and abnormal returns (e.g., GIM and CN). This allows us to investigate the abnormal returns associated with the governance-spread portfolios in Section 5. The data requirement is that the firm is included in the IRRC database. This limits the sample to firms in the S&P 500, mid-cap 400, and small-cap 600 indices between 1990 and 2003, and reduces the number of realized targets to 412 firms. The results from this model can be different from the previous model not only because of differences in the time period, but also because this sample consists of relatively much larger firms.

For this smaller sample, we introduce two additional independent variables that are not available before 1990. The first captures the amount of takeover protection a firm has and is denoted by *EXT*. *EXT* is a linear transformation of G index constructed by GIM, so that a higher value of *EXT* (=24 − G) indicates a greater takeover exposure or greater shareholder rights. We also use a variable capturing the complementary effect between takeover defenses and blockholdings identified in CN.

Panel B of Table 1 presents the mean values of these variables for this smaller IRRC universe over 1991–2004 for which there are no missing IRRC or Compustat data, again separating the targets from all the other firms. We also test whether the means of the target group are different. For the sample of all announced and completed takeovers, all variables except asset structure and blockholding are significantly different for the target group. For the

**Table 1**
**Descriptive statistics**

| | Using announced and completed takeovers | | | Using 100% completed takeovers | | |
|---|---|---|---|---|---|---|
| | Mean nontargets | Mean targets | *t*-stat difference | Mean nontargets | Mean targets | *t*-stat difference |
| Panel A: Sample for 1981—2004 | | | | | | |
| *Q* | 2.03 | 1.90 | 2.99 | 2.03 | 1.82 | 3.48 |
| PPE | 0.55 | 0.55 | 0.31 | 0.55 | 0.54 | 1.32 |
| ln(Cash) | 1.69 | 1.71 | 0.46 | 1.69 | 1.84 | 3.31 |
| BLOCK | 0.47 | 0.55 | 12.06 | 0.47 | 0.63 | 16.52 |
| ln(Mktcap) | 4.96 | 4.95 | 0.25 | 4.95 | 5.06 | 2.47 |
| Industry | 0.86 | 0.90 | 7.86 | 0.86 | 0.92 | 9.28 |
| Leverage | 0.26 | 0.28 | 3.02 | 0.27 | 0.26 | 1.36 |
| ROA | −0.06 | −0.09 | 2.06 | −0.06 | −0.05 | 0.65 |
| No. of observations | 78,295 | 5,457 | | 80,939 | 2,813 | |
| Panel B: Sample for 1991—2004 | | | | | | |
| *Q* | 2.14 | 2.00 | 1.55 | 2.14 | 1.75 | 3.16 |
| PPE | 0.57 | 0.57 | 0.06 | 0.57 | 0.60 | 1.25 |
| ln(Cash) | 3.55 | 3.30 | 3.65 | 3.54 | 3.41 | 1.37 |
| BLOCK | 0.78 | 0.78 | 0.17 | 0.78 | 0.85 | 3.76 |
| ln(Mktcap) | 7.07 | 6.72 | 5.86 | 7.05 | 6.88 | 2.17 |
| Industry | 0.88 | 0.90 | 1.96 | 0.87 | 0.94 | 4.24 |
| Leverage | 0.25 | 0.28 | 5.34 | 0.25 | 0.27 | 2.10 |
| ROA | 0.01 | −0.04 | 7.92 | 0.01 | −0.01 | 2.38 |
| EXT | 17.11 | 19.08 | 11.66 | 17.15 | 19.51 | 10.15 |
| EXT × BLOCK | 13.19 | 14.79 | 5.41 | 13.18 | 16.69 | 8.63 |
| No. of observations | 14,533 | 799 | | 14,920 | 412 | |

This table presents the descriptive statistics of the independent variables used in the logit model of Table 2 for the Compustat-based sample for the sample period 1981–2004 in panel A, and for the sample covered by the Investor Responsibility Research Center (IRRC) for 1991–2004 in panel B. *Q* is the ratio of market-to-book value of assets, where market assets are defined as total assets plus market value of common stock minus book common equity and differed taxes. PPE is property, plant, and equipment to assets ratio. Industry is equal to 1 if, based on the Fama-French 48-industry classifications, there was a takeover in a firm's industry in the prior year. ROA is the return on assets. Leverage is the book debt to asset ratio. Cash is cash and short-term investments to assets ratio. Firm size is proxied by ln(Mktcap), the natural log of market equity. All independent variables are measured at the end of the fiscal year previous to the takeover event. *BLOCK* is a dummy variable equal to 1 if (at least) one institutional investor holds more than 5% of the companies stock and 0 otherwise. *EXT* is (24 − G), where G is the governance index as defined by Gompers, Ishii, and Metrick (2003). We separate out firms that were takeover targets in a given year, and also distinguish between all announced and completed takeovers versus completed takeovers only. Finally, we provide the mean of each variable (averaged over all firm-years) for both the target and nontarget groups, and the *t*-statistic for testing whether those means across the two groups are different.

**Table 2**
**Takeover vulnerability: likelihood of being acquired**

| Variable | Takeover likelihood, 1981–2004 | | | | Takeover likelihood, 1991–2004 | | | |
|---|---|---|---|---|---|---|---|---|
| | Coefficient | *t*-stat | *p*-value | Sign. | Coefficient | *t*-stat | *p*-value | Sign. |
| Panel A: Using announced and completed takeovers | | | | | | | | |
| Q | −0.042 | 5.26 | 0.00% | *** | −0.083 | 3.34 | 0.10% | *** |
| PPE | 0.031 | 1.13 | 25.90% | | 0.113 | 0.90 | 37.00% | |
| ln(Cash) | 0.003 | 0.34 | 73.60% | | 0.023 | 0.78 | 43.80% | |
| BLOCK | 0.287 | 9.59 | 0.00% | *** | −1.043 | 2.83 | 0.50% | *** |
| ln(Mktcap) | −0.025 | 2.11 | 3.50% | ** | −0.072 | 1.92 | 5.50% | * |
| Industry | 0.137 | 2.87 | 0.40% | *** | −0.025 | 0.19 | 84.60% | |
| Leverage | 0.101 | 3.57 | 0.00% | *** | 0.729 | 4.35 | 0.00% | *** |
| ROA | −0.020 | 2.17 | 3.00% | ** | −0.527 | 4.45 | 0.00% | *** |
| EXT | | | | | 0.048 | 2.82 | 0.50% | *** |
| EXT*BLOCK | | | | | 0.053 | 2.81 | 0.50% | ** |
| Pseudo $R^2$ | 1.76% | | | | 4.95% | | | |
| Observations | 83,752 | | | | 15,332 | | | |
| Targets | 5,457 | | | | 799 | | | |
| Panel B: Using 100% completed takeovers | | | | | | | | |
| *Q* | −0.067 | 5.03 | 0.00% | *** | −0.256 | 5.55 | 0.00% | *** |
| PPE | 0.021 | 0.52 | 60.50% | | 0.301 | 1.76 | 7.80% | * |
| ln(Cash) | 0.009 | 0.66 | 51.10% | | 0.002 | 0.04 | 96.60% | |
| BLOCK | 0.559 | 13.25 | 0.00% | *** | −0.860 | 1.40 | 16.10% | |
| ln(Mktcap) | −0.034 | 1.99 | 4.60% | ** | 0.023 | 0.45 | 65.10% | |
| Industry | 0.353 | 4.77 | 0.00% | *** | 0.373 | 1.67 | 9.50% | * |
| Leverage | −0.010 | 0.10 | 91.70% | | 0.084 | 0.31 | 75.80% | |
| ROA | 0.014 | 0.22 | 82.70% | | −0.221 | 1.08 | 27.90% | |
| EXT | | | | | 0.080 | 2.75 | 0.60% | *** |
| EXT*BLOCK | | | | | 0.066 | 2.15 | 3.20% | ** |
| Pseudo $R^2$ | 3.13% | | | | 9.27% | | | |
| Observations | 83,752 | | | | 15,332 | | | |
| Targets | 2,813 | | | | 412 | | | |

This table presents results of the maximum likelihood estimates of the logit model for the Compustat-based sample for the sample period 1981–2004 and for the sample covered by the Investor Responsibility Research Center (IRRC) for 1991–2004. Panel A reports the results using all announced and completed takeovers. Panel B reports the results using 100% completed takeovers only. The dependent variable is a dummy (*target*) equal to 1 if the company is target of an acquisition (friendly or hostile or neutral). See Table 1 for a description of the variables. All Compustat variables are industry adjusted (*Q*, PPE, ln(cash), leverage, and ROA). The logit also includes year dummies, which are not reported. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

sample of completed takeovers only, all variables except asset structure and cash are significantly different.

In the logit specification, the probability of becoming a target in the next year is thus estimated by using values of the independent variables at the end of the previous year. Table 2 shows the results for the two samples (Compustat sample in the time period 1981–2004, and IRRC sample for 1991–2004). All the Compustat variables are industry adjusted, and each logit specification also includes year dummies.

The logit estimation using announced and completed takeovers has more significant variables than using completed takeover only, potentially benefiting from more information from a larger set of target firms. Consistent with the prior literature, the generally statistically significant variables are *BLOCK*, the industry dummy variable intended to capture the clustering of takeover activity within industry and time, market-to-book ($Q$), and firm size (ln(Mktcap)). The positive coefficient on leverage is a bit puzzling, indicating that higher leverage increases the likelihood of being acquired, but this is only significant using all announced and completed takeovers, *albeit* in both the Compustat and the IRRC samples. However, it is consistent with the higher leverage of targets than nontargets in Table 1. Further, underperforming firms tend to be targets, as evidenced by the generally negative coefficient on ROA, which also is only significant using all announced and completed takeovers, *albeit* in both the Compustat and the IRRC samples. Finally, the two additional variables in the IRRC sample for 1991–2004 are both significant with the expected signs. Fewer takeover defenses (higher *EXT*) positively predict takeovers. The complementary effect (interacting *EXT* with institutional blockholding) indicates that takeover defenses are about twice as important in the presence of an institutional blockholder.

In the next section, we use these estimated coefficients to sort firms into portfolios based on the likelihood of being a takeover target. In a crucial robustness test, we then also estimate logit models using 10-year rolling estimation windows to remove any "look-ahead bias," where the takeover vulnerability estimates only rely on the past information.

Finally, the overall fit of the logit models is modest but similar to the previous literature (e.g., Ambrose and Megginson, 1992). For example, the pseudo $R^2$ equals 3.13% and 9.27% for the 1981–2004 and 1991–2004 samples, respectively (using completed takeovers, see panel B of Table 2). As our focus is primarily on the extent to which firms fall into either of the extreme groups of lowest versus highest estimated takeover likelihood, another way to think about the fit is to compare the percentage of actual targets falling into these extreme groups. Using quintile portfolios, the percentage of targets in the first and fifth takeover likelihood groups equals 15% and 25%, respectively, for the 1981–2004 sample, and equals 12% and 36%, respectively, for the 1991–2004 sample (again using completed takeovers). The differences between these percentages and their individual differences from 20% are clearly statistically significant.

### 2.2 Returns to portfolios based on takeover vulnerability

We sort firms into quintile and decile portfolios based on their takeover vulnerability, which is estimated in the different logit regressions. From the preceding section, we can see that firms with an institutional blockholder, low $Q$, low market capitalization, operating in an industry where a takeover occurred the previous year, higher leverage, and lower operating performance will tend to appear in the portfolio that has the highest exposure to takeovers. It is important to note that any one of the firm-specific characteristics alone does not dictate the portfolio that a firm is assigned to.[15] We focus on the equal-weighted returns

[15] Let us, for the sake of illustration, focus on market capitalization. A low market capitalization firm might have a high ROA, high $Q$, lack a blockholder, low fixed assets, and operate in an industry that hasn't recently witnessed

for the remainder of the paper in an attempt to reduce the noise inherent in predicting takeover targets.[16]

We investigate the returns of each of the quintile and decile takeover-vulnerability-sorted portfolios, as well as the returns to long-short portfolios that buy firms with the highest takeover vulnerability and shorts firms with the lowest takeover vulnerability. For additional robustness, we also investigate the returns to a takeover-spread portfolio that is formed based on decile, rather than quintile, classifications. The returns to these two sets of portfolios are adjusted for four factors capturing risk or style effects: the market factor, the size, and book-to-market factors proposed by Fama and French (1993), as well as the Carhart (1997) momentum factor.

The theoretical framework suggests two possibilities. If the factors in the four-factor model correctly capture the risk associated with time variation in the aggregate fundamentals and discount rates, we would not expect to find a significant abnormal return to the takeover-spread portfolio. In that case, a portfolio of firms more likely to be taken over would only have different betas. If, however, the four-factor model does not account for all such factors, we should find a significant abnormal return to the takeover-spread portfolio.[17]

Table 3 presents the mean returns and alphas of the quintile portfolios and the long-short portfolios based on both quintile and decile sorts, using the logit for announced and completed takeovers in panel A and for completed takeovers in panel B. We show results for three separate samples in each panel. The first two samples are 1981–2004 for all Compustat firms and 1991–2004 for all IRRC firms, and use the logit results from Table 2. These logit estimations use information for the whole period of 1981–2004, the same period used for sorting stocks into portfolios and calculating abnormal returns. Therefore, a vital robustness check is to confirm our results using only the past information. Otherwise, it could be possible that a "look-ahead bias" is responsible for these results (e.g., Butler, Grullon, and Weston, 2005). These real out-of-sample results are the third sample using rolling 10-year logit estimation windows, so that we can calculate alphas over 1991–2004 using all Compustat firms.[18]

an acquisition. Such a firm will not appear in the portfolio with the highest exposure to takeovers. Similarly, a firm with high market capitalization might appear in the portfolio with the highest takeover exposure if the firm has a blockholder, low ROA and low $Q$, high fixed assets, and is in an industry that has recently witnessed an acquisition.

16 The value-weighted results give similar, but weaker results, which in some CSRs (see Section 3) are not significant.

17 Since the market captures both the shocks to aggregate fundamentals and to discount rates (CV), it is reasonable to expect abnormal returns relative to a market model even when higher shocks to aggregate fundamentals are the only relevant channel.

18 The number of years in the rolling logit estimations is chosen to balance two effects. Utilizing only recent information, and hence using short windows reduces the number of realized targets. This lack of observations makes it difficult to arrive at any robust estimation. On the other hand, increasing the estimation window leaves us with fewer years to conduct our analysis. For example, if we consider a 20-year rolling logit regression, we are left with only 4 years (2001–2004) for which we can compute abnormal returns and perform cross-sectional tests. To balance these counteracting concerns, we choose 10 years as the time period in each logit. This allows us to focus our analysis on the post-1990 period.

**Table 3**
**Abnormal returns associated with takeover vulnerability**

| Takeover likelihood | 1981–2004 Mean | 1981–2004 Alpha | 1981–2004 $t$-stat | 1991–2004 Mean | 1991–2004 Alpha | 1991–2004 $t$-stat | Rolling estimation windows 1991–2004 Mean | Rolling estimation windows 1991–2004 Alpha | Rolling estimation windows 1991–2004 $t$-stat |
|---|---|---|---|---|---|---|---|---|---|
| Panel A: Using announced and completed takeovers | | | | | | | | | |
| 1 | 1.61% | −3.80% | 2.80 | 11.53% | −1.05% | 0.69 | 9.46% | −1.78% | 0.65 |
| 2 | 11.04% | 2.09% | 1.69 | 13.34% | 0.24% | 0.14 | 15.68% | 5.78% | 2.97 |
| 3 | 11.13% | 5.49% | 3.36 | 17.10% | 3.65% | 1.92 | 17.39% | 7.30% | 3.91 |
| 4 | 9.39% | 5.79% | 3.40 | 20.53% | 6.96% | 3.67 | 18.20% | 6.00% | 4.09 |
| 5 | 13.85% | 7.97% | 5.42 | 26.34% | 11.06% | 4.34 | 18.41% | 7.95% | 3.46 |
| 5–1 | 12.24% | 11.77% | 7.18 | 14.81% | 12.11% | 4.14 | 8.95% | 9.72% | 2.74 |
| 10–1 | 20.74% | 21.67% | 10.00 | 17.17% | 13.18% | 3.16 | 13.50% | 15.32% | 3.34 |
| Panel B: Using 100% completed takeovers | | | | | | | | | |
| 1 | 1.71% | −3.58% | 2.39 | 12.87% | 0.28% | 0.18 | 8.57% | −0.62% | 0.23 |
| 2 | 9.26% | 0.14% | 0.11 | 13.87% | 0.35% | 0.19 | 14.71% | 1.67% | 1.10 |
| 3 | 11.55% | 5.26% | 3.09 | 17.48% | 4.19% | 2.50 | 15.01% | 3.87% | 2.10 |
| 4 | 9.59% | 6.63% | 3.45 | 20.22% | 6.24% | 3.12 | 21.75% | 12.84% | 4.83 |
| 5 | 14.88% | 9.14% | 6.54 | 24.11% | 9.55% | 4.54 | 18.98% | 7.52% | 3.74 |
| 5–1 | 13.17% | 12.72% | 7.68 | 11.24% | 9.27% | 3.89 | 10.41% | 8.14% | 2.76 |
| 10–1 | 22.16% | 23.76% | 11.37 | 12.37% | 9.14% | 2.97 | 15.60% | 13.98% | 3.51 |

We report the annualized mean, the annualized abnormal return (alpha), and the corresponding $t$-statistic of five equally weighted portfolios that are sorted according to their takeover vulnerabilities using the coefficients estimated in Table 1, for all announced and completed takeovers in panel A, and for all 100% completed takeovers in panel B. We also report the annualized mean and alpha and the corresponding $t$-statistic of an equally weighted portfolio that buys firms in the highest takeover likelihood category and shorts firms in the lowest category based on quintile (5–1) and decile (10–1) sorts. The alphas are relative to the four-factor Fama-French (1992) and Carhart (1997) models. We report the results for three separate samples: the entire Compustat sample for the years 1981–2004; the Investor Research Responsibility Center sample between years 1991 and 2004; and the entire Compustat sample for the years 1991–2004. The first two employ the takeover likelihoods from the respective logit models from Table 2. The third sample uses rolling 10-year estimation windows to estimate the logit.

As this sample based on out-of-sample rolling regressions only starts in 1991, several other tests done in the subsequent sections in this paper could only be conducted with the first (or the first two) of these three samples. Accordingly, we cannot rule out the possibility that these other results are (at least partly) driven by a "look-ahead bias," as data limitations prevent us from doing the out-of-sample robustness check.

We find that both the mean returns and the abnormal returns are generally increasing with the likelihood of takeovers. We first consider the results using the logit for announced and completed takeovers. An equal-weighted portfolio that buys firms with a high takeover vulnerability (quintile 5) and shorts firms with a low takeover vulnerability (quintile 1) generates a highly significant annualized abnormal return of 11.77% between 1981 and 2004, with a $t$-stat of 7.18. Using decile classifications, the abnormal returns to such a takeover-spread portfolio is even more striking and equals 21.67% with a $t$-stat of 10.0.[19]

[19] To shed light on the source of these abnormal returns, we remove from our samples all firms that were actual targets, and recompute abnormal returns accruing to the different portfolios. Our results remain consistent and of (an arguably surprisingly) similar magnitude. Therefore, these abnormal returns are not caused by the announcement returns to realized targets (not tabulated to save space).

The corresponding numbers for the value-weighted portfolio are, as expected, lower and equal to 2.90% ($t$-stat = 1.64) for quintile sorts and 7.76% ($t$-stat = 3.49) for the decile classifications (not tabulated).

The results using the logit for completed takeovers are very similar. As it turns out, the returns of that quintile takeover-spread portfolios have a correlation of 95% with the corresponding takeover-spread portfolio based on the logit for announced and completed takeovers. Therefore, in the remainder of the paper, we only report the results using the logit for announced and completed takeovers.

The results for the sample between 1991 and 2004 using the logit model that includes takeover defenses as an additional independent variable (*EXT*) are also similar. Again, we find that abnormal returns increase with takeover vulnerability. The takeover-spread portfolio generates an annualized abnormal return of 12.11% ($t$-stat = 4.14) for the quintile classification and 13.18% ($t$-stat = 3.16) for the decile classification.

Finally, the results are robust to using the rolling 10-year logit estimation windows. In this case, where only previous information is used when calculating takeover probabilities, the takeover-spread portfolio generates an annualized alpha of 9.72% ($t$-stat = 2.74) for the quintile classification and 15.32% ($t$-stat = 3.34) for the decile classification. However, the estimates based on the out-of-sample rolling regressions are much noisier, as indicated by the considerably higher standard deviations. This is likely due to the smaller sample size and the removal of any "look-ahead bias."

The results in this section are consistent with the notion that takeover vulnerability strongly affects the rate of return. In support of Proposition 1, we find that a greater takeover vulnerability is associated with a higher rate of return. The proposition also states that takeover vulnerability increases firm values as well. Direct evidence is provided in GIM and CN linking better takeover governance with higher $Q$ ratios.[20] Further, the above results also appear to support the "agency costs" acquisition motive that makes takeover targets more sensitive to aggregate fundamentals rather than to discount-rate shocks (Proposition 2). The four-factor model does not seem to capture this risk completely.

## 3. The "Takeover" Factor and Takeover Betas

The "takeover" factor is intended to mimic the state variables related to time-varying risk premiums, and is constructed as the equally weighted long-short portfolio that buys firms in the highest quintile and sells firms in the lowest quintile of takeover vulnerability. Bruner (2004) and Rhodes-Kropf and

[20] The coefficient on $Q$ in the takeover logit regressions is negative, which is apparently incompatible with takeover targets having higher firm values, suggesting that firms with lower $Q$ are more likely to be taken over. However, Proposition 1 states that, *ceteris paribus*, takeover targets should have a higher valuation. $Q$ is affected by several factors, some of which are potentially unrelated to takeovers and consequently to check whether our result is true, one needs to control for other factors and then check if takeover defenses hurt firm value. This is exactly what GIM and CN have done.

Viswanathan (2004) confirm that takeover activity is time varying and indeed related to the conditions in the equity market, so that exposure to takeovers could have an important impact on returns.

In this section, we use three tests to confirm that our logit model and the resulting takeover factor indeed capture cross-sectional differences in takeover vulnerability. First, the model and the takeover factor can predict real takeover activity. Second, cross-sectional changes in the takeover likelihood are directly related to changes in the corresponding takeover betas. We show this by considering the adoption of state antitakeover legislation, which makes takeovers of firms incorporated in the affected state more difficult, so that their takeover betas decrease subsequently. Third, the takeover factor can explain the previously documented abnormal returns accruing to governance-based spared portfolios (see GIM and CN).

### 3.1 Predicting takeover activity

Figure 1 plots the annual return to the takeover-spread portfolio together with the average takeover activity and the (scaled) difference in the average takeover likelihood of the firms in the two extreme quintile portfolios for the full sample of 1981–2004.[21] Takeover activity is measured each year as the (normalized) average deal value, taking into account all announced and completed takeovers. The average takeover likelihood of the firms in the first and fifth quintile takeover-likelihood-sorted portfolio equals 1.75% and 4.04%, respectively.[22]

As the figure indicates, the takeover factor indeed appears to predict takeover activity and thus seems related to the real takeover activity in the economy. Similarly, the difference in the takeover likelihood across the two extreme quintile portfolios seems to be also leading the actual takeover activity. More formally, the correlation between the lagged annual returns of the takeover factor and takeover activity equals 41%, and the correlation between the lagged takeover likelihood difference and takeover activity equals 65% (see panel A of Table 4). Regressions of takeover activity on the lagged takeover factor returns or the lagged likelihood differences give significance in both cases (see panels B and C of Table 4, respectively). The lagged takeover likelihood remains significant even after controlling for lagged takeover activity. Even though we only have 24 annual observations, this provides some support that the takeover factor is indeed picking up takeover vulnerability rather than some more general business cycle factor.

### 3.2 State antitakeover laws and firm-level takeover betas

In this subsection, we use state adoptions of antitakeover legislation as events that represent exogenous shocks to takeover vulnerability: only the firms

[21] The logit estimation gives the takeover likelihood for each firm for each year, which is averaged for all firms in the highest and lowest quintiles of takeover likelihood, and their difference across these quintiles is scaled to have the same standard deviation as the takeover activity to facilitate comparisons.

[22] These are averaged across time and across all firms in the respective portfolios, without any scaling.

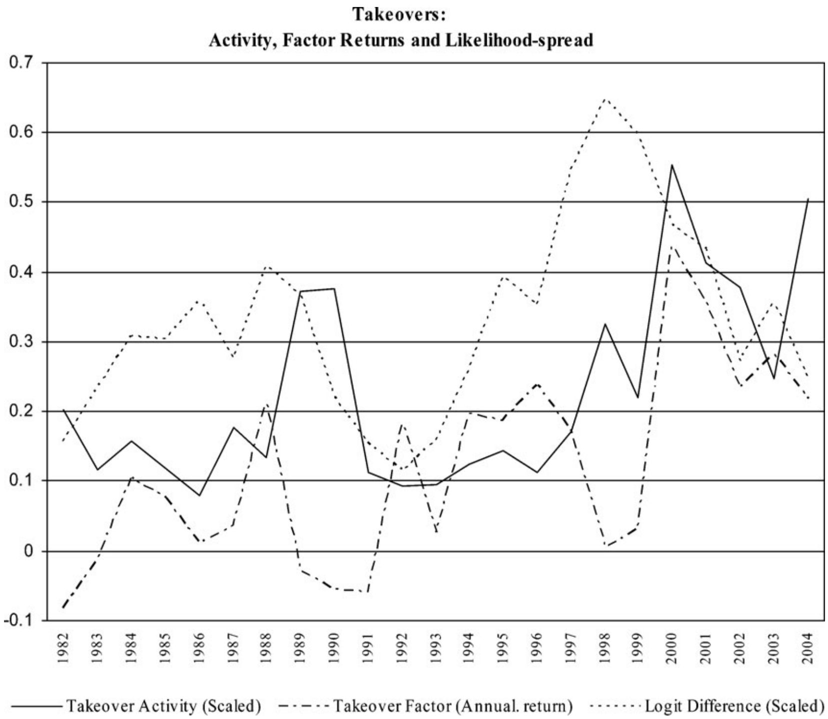


**Figure 1**
**Takeovers 1981–2004: activity, factor returns, and likelihood**
This figure plots the average takeover activity, takeover factor returns, and the takeover likelihood spread for each year over 1981–2004. The takeover activity is measured as the average deal value of all completed takeovers in SDC, the takeover factor is the equally weighted long-short portfolio based on quintiles from panel A of Table 3, and logit is calculated as the difference in the average takeover likelihood of the two extreme quintile portfolios that make up the takeover factor, based on the logit coefficients as reported in panel A of Table 2.

incorporated in the affected states should be affected.[23] Over the course of the 1980s, most (but not all) states passed “second-generation” antitakeover laws (SGAT) that made takeovers of firms incorporated in the affected states more difficult. Specifically, if exposure to the takeover factor indeed is related to the actual takeover vulnerability, a firm’s takeover beta should decrease after the state in which the firm is incorporated adopts a SGAT law. Bertrand and Mullainathan (2003) provide more discussion and detailed lists of these events. We adopt the methodology of Cheng, Nagar, and Rajan (2005) in focusing on the impact of the first SGAT law passed in each state.

Our empirical test consists of two steps. In the first step we estimate, using daily returns, annual firm-level betas with respect to a five-factor model that includes the four factors of the Fama-French and Carhart model plus the takeover

[23] We thank Lubomir Litov for sharing his data on state takeover legislation (see John and Litov, 2005). We thank the anonymous referee for suggesting this test.

**Table 4**
**Takeovers: activity, factor returns, and likelihood differences**

| Correlation matrix | Activity | Factor | Logit difference | Activity (lagged) | Factor (lagged) |
|---|---|---|---|---|---|
| Panel A: Descriptive statistics | | | | | |
| Factor | 39% | | | | |
| Logit difference | 28% | 26% | | | |
| Lagged activity | 50% | 24% | 8% | | |
| Lagged factor | 41% | 34% | 20% | 32% | |
| Lagged logit difference | 65% | 27% | 76% | 38% | 22% |
| Variable | Coefficient | $t$-stat | | Coefficient | $t$-stat |
| Panel B: Predicting takeover activity using lagged takeover factor | | | | | |
| Constant | 0.07 | 4.97 | | 0.04 | 2.02 |
| Lagged factor | 1.77 | 2.11 | | 1.21 | 1.51 |
| Lagged activity | | | | 0.45 | 2.20 |
| $R^2$ | 9% | | | 17% | |
| Panel C: Predicting takeover activity using lagged logit difference | | | | | |
| Constant | 0.01 | 0.31 | | −0.01 | 0.25 |
| Lagged logit difference | 4.39 | 3.97 | | 3.63 | 3.25 |
| Lagged activity | | | | 0.32 | 1.78 |
| $R^2$ | 24% | | | 29% | |

Panel A reports the correlation matrix for these variables at the annual frequency: takeover activity (Activity), the takeover-factor returns (Factor), and the takeover likelihood differences associated with the long and short portfolios that make up the takeover factor (Logit). The takeover activity is measured as the average deal value of all completed takeovers in SDC, the takeover factor is the equally weighted long-short portfolio based on quintiles from panel A of Table 3, and logit is calculated as the difference in the average takeover likelihood of the two extreme quintile portfolios that make up the takeover factor, based on the logit coefficients as reported in panel A of Table 2. Lagged means lagged by a single year. Panels B and C report predictive regressions of takeover activity on the takeover factor and the logit difference, respectively.

factor. In the second step, we conduct pooled panel regressions of these annual takeover betas on a dummy indicating whether SGAT laws have been passed, controlling for firm- and year-fixed effects. As a result, we closely follow the approach of Bertrand and Mullainathan (2003), who advocate using the full cross-section of all states before and after passing SGAT laws. As they write, this approach

> "accounts for the fact that there are many takeover laws staggered over time. The staggered passage of the anti-takeover statutes also means that our control group is not restricted to states that never pass a law. . . . It implicitly takes as the control group all firms incorporated in states not passing a law at [that] time."

Panel A of Table 5 presents some descriptive statistics, showing that the average takeover beta equals 0.67% with a large standard deviation. Averaging across both time series and the cross-section, about half the firms are incorporated in a state that passed a SGAT law. The correlation between the takeover beta and the dummy that a SGAT law is adopted equals 4.97% and is significant at 1% (see panel B). Finally, panel C of Table 5 gives the results for the pooled panel regressions of annual firm takeover betas on the dummy of whether a

**Table 5**
**State antitakeover laws and firm-level takeover betas**

| | Mean | Standard deviation | | |
|---|---|---|---|---|
| Panel A: Descriptive statistics | | | | |
| Takeover beta | 0.67% | 93.39% | | |
| Dummy (law adopted) | 45.17% | 49.77% | | |
| Takeover probability | 3.01% | 1.66% | | |
| Takeover probability × Dummy | 1.50% | 2.03% | | |
| | Takeover beta | Dummy (law adopted) | Takeover probability | |
| Panel B: Correlation matrix | | | | |
| Takeover beta | | | | |
| Dummy (law adopted) | 4.97% | | | |
| Takeover probability | 10.66% | 16.49% | | |
| Takeover probability × Law adopted | 7.51% | 81.65% | 53.93% | |
| | (1) | (2) | (3) | (4) |
| Panel C: Takeover beta regressions | | | | |
| Dummy (law adopted) | −0.040 (2.17) | −0.043 (2.36) | | −0.016 (0.69) |
| Takeover probability | | 10.18 (22.19) | 10.77 (21.28) | 10.66 (20.08) |
| Takeover probability × Law adopted | | | −1.10 (2.92) | −0.90 (1.86) |
| $R^2$ | 0.36% | 2.56% | 2.63% | 2.64% |
| Year dummies | Yes | Yes | Yes | Yes |
| Firm-fixed effects | Yes | Yes | Yes | Yes |
| No. of observations | 78,266 | 78,266 | 78,266 | 78,266 |

Panel A presents descriptive statistics of the three variables of interest for the time period of 1980–2004. Takeover beta is the annual, firm-level takeover beta (i.e., beta on the takeover factor) in a time-series regression, estimated separately each year, of daily firm excess returns on the four-factor Fama-French model with the takeover factor added. Dummy (law adopted) equals 1 if the state in which a firm is incorporated has passed a first major antitakeover law. Takeover probability is the firm's takeover likelihood from the logit of panel A of Table 2. Panel B presents the correlation matrix across both time series and cross-sectional dimensions. Takeover probability × Law adopted is the interaction between the firm's takeover likelihood and Dummy (law adopted). Panel C presents the results from a pooled panel regression of firm-level takeover betas on the dummy indicating the state-level antitakeover law was passed, the firm's takeover likelihood, and the interaction of these. In each regression, we include both firm- and year-fixed effects, and cluster the robust standard errors by firm. *t*-statistics are provided below each coefficient in parentheses.

SGAT law has passed, the firm's *ex-ante* takeover probability according to the fitted values of the logit model of panel A of Table 2, and the interaction of the takeover probability and the dummy.[24] Each regression also includes year dummies and firm-fixed effects, and the robust standard errors are clustered by the firm.

In regressions (1) and (2), the coefficient on the dummy is negative and significant without and with controlling for the takeover probability, respectively. Regressions (3) and (4) add the interaction of the takeover probability with the dummy, and show that the decrease in takeover beta is stronger for firms that are more likely targets. Therefore, the takeover beta decreases after the state of a firm's incorporation passes a SGAT law, indicating that a firm's exposure to the takeover factor is indeed affected by exogenous shocks to its takeover vulnerability.

[24] This subsection uses the 1981–2004 sample. We cannot use the 10-year rolling estimation approach or the 1991–2004 sample, as all SGAT laws were passed in the 1980s.

### 3.3 The takeover factor and abnormal returns associated with governance

In this subsection, we examine the impact of the takeover factor on the findings in GIM and CN. These papers investigate the impact of corporate governance on firm value using valuation measures, accounting measures of profitability, and equity returns. With regards to equity returns, GIM compile a G index and document that firms with more shareholder rights (low G index) have higher abnormal returns relative to a Fama-French model. CN show that the positive abnormal return accruing to firms with low levels of protection exists only, and is larger, if the lack of takeover defenses is combined with a large external shareholder.

The results in the previous section show that the takeover factor has large abnormal returns. We investigate whether the abnormal returns documented in GIM and CN decrease once the asset pricing model includes the takeover factor. In doing so, we will be able to shed light on the source of the high abnormal returns initially documented in GIM. They speculate that these results could be due to, for example, investors learning about the importance of corporate governance over the time of their sample. Another possibility they discuss is some type of omitted-variable bias or model misspecification. A direct causal link between governance and returns is rejected by Core, Guay, and Rusticus (2006) who do not find evidence that the market is negatively surprised by the poor operating performance of weak governance firms. On the other hand, Cremers, Nair, and Wei (2007) present results indicating that the combination of fewer takeover defenses and the presence of institutional blockholders leads to higher credit spreads and higher expected returns for corporate bonds.

We focus on the sample for which takeover defense information, as used in GIM and CN, is available and consequently estimate takeover vulnerabilities based on the corresponding logit (i.e., the 1991–2004 sample of panel A of Table 2).[25] Since the variables used to form the governance portfolios in GIM and CN are also used in the logit model, it is important to first underline the merits of the logit model employed. First, the logit model has many other characteristics beyond G index and blockholding that contribute to the logit estimation. Further and most crucially, the correlation between the returns of the "democracy-minus-dictatorship" (low minus high managerial protection) portfolio used by GIM and the takeover factor is quite low (6%). Therefore, there is no a priori empirical reason to suspect a strong connection between these two portfolios.

Following GIM, we use the "G index" they compile ($<0<G<24$), and first form a portfolio that buys firms with the lowest level of takeover protection ($G < 6$) and shorts firms with the highest level of takeover protection ($G > 13$). To characterize the lowest and the highest level, we use the same cutoff levels as GIM and the same terminology to call this the "democracy-minus-dictatorship" portfolio. First, we consider the same time period as GIM and replicate their

[25] We cannot use the 10-year rolling estimation approach as *EXT* is only available starting in 1990.

**Table 6**
**Abnormal returns associated with governance spread portfolios**

| | FF4 | FF4 + takeover |
|---|---|---|
| Panel A: Democracy—dictatorship long-short portfolios, 1991–1999 | | |
| VW alpha | 8.65% | 4.59% |
| | (2.97) | (1.36) |
| EW alpha | 4.70% | 2.59% |
| | (2.00) | (0.94) |
| Panel B: Democracy—dictatorship long-short portfolios, 1991–2004 | | |
| VW Alpha | 4.40% | 2.70% |
| | (1.65) | (0.95) |
| EW alpha | 3.63% | −0.52% |
| | (1.65) | (0.24) |
| Panel C: Democracy—dictatorship conditional on BLOCK, 1991–2004 | | |
| VW alpha | 6.72% | 3.54% |
| BLOCK = 4 | (1.86) | (0.82) |
| EW alpha | 4.68% | 3.23% |
| BLOCK = 4 | (1.83) | (0.86) |

We report the annualized mean, the annualized abnormal return (alpha), and the corresponding $t$-statistic of a (value-weighted, VW, and equal-weighted, EW) portfolio that buys firms in the highest category of governance (fewest takeover defenses or most shareholder rights) and shorts firms in the lowest category of governance. Governance is measured using the G-index, compiled by GIM, and by a combination of the G-index and institutional blockholding (*BLOCK*, see CN). The alphas are first computed relative to the four-factor model and then relative to a five-factor model that appends the four-factor model with a takeover-spread portfolio. The takeover-spread portfolio buys firms in the highest category and shorts firms in the lowest category of takeover vulnerability (see Table 3). $t$-statistics are provided below each alpha, in parentheses.

result of the abnormal returns to the democracy-minus-dictatorship portfolio between 1990 and 1999 (Table 6, panel A). Consistent with the findings of GIM, we find that the democracy-minus-dictatorship portfolio is associated with an annualized abnormal return of 8.65% ($t$-stat = 2.97) relative to an asset pricing model that uses market, size, book-to-market, and momentum factors.[26]

Next, we append the four-factor model with the takeover factor. The democracy-minus-dictatorship portfolio now generates a much lower abnormal return of 4.59% and is no longer significant ($t$-stat = 1.36, see panel A of Table 6). The equal-weighted version of such a portfolio is associated with an abnormal return of 2.59%, which is also insignificant at standard levels. This documented reduction in abnormal returns also follows when the time period considered is extended from 1999 to 2003—decreasing from 4.40% ($t$-stat = 1.65) to 2.70% ($t$-stat = 0.95) for the value-weighted case and from 3.62% ($t$-stat = 1.64) to −0.52% ($t$-stat = −0.24) for the equal-weighted portfolios. However, for the time period between 1991 and 2004, the abnormal returns of the democracy-minus-dictatorship portfolio, even without the takeover factor, are low.

One possible reason for a weakening of the GIM results on extending the time period from 1999 to 2003 is perhaps the reduction in takeover activity during

[26] The abnormal returns are not exactly identical (a difference of 0.20%) due to slight differences in the construction of the momentum factor.

this time period.[27] As suggested by our framework, lower takeover activity would imply a smaller difference in the returns between firms exposed to and firms protected from takeovers. Another reason is provided by CN. They find that takeover defenses and shareholder monitoring are complements in being associated with equity abnormal returns and accounting performance. Further, they document the complementary effect to be stronger in smaller firms. Using only takeover defenses, through G index, might be capturing only part of the true effect associated with governance.

Therefore, we verify the robustness of the pattern that abnormal returns associated with corporate governance decrease when the takeover-spread factor is included in the asset pricing model. To do so, we check the changes in abnormal returns associated with the existence of both low takeover defenses and high shareholder monitoring (see CN) when the takeover-spread portfolio is added to the asset pricing model. We first compute the abnormal returns to a portfolio that buys firms with few takeover defenses and high shareholder monitoring and shorts firms with many takeover defenses and low shareholder monitoring. To proxy for shareholder monitoring, we follow CN and use the presence of an institutional blockholder (*BLOCK*). Without the takeover factor, the abnormal return of this governance-spread portfolio from 1990 to 2004 is 6.72% (using *BLOCK*). On introducing the takeover-spread portfolio to the Fama-French model, however, the documented abnormal return to the complementary governance portfolio also decreases from 6.72% ($t$-stat = 1.86) to 3.54% ($t$-stat = 0.82).

This finding has important implications, suggesting that the documented abnormal returns associated with governance are (at least partly) due to the misspecification of the asset pricing model. As discussed, this sheds light on the interpretation of the findings in GIM and CN, and is consistent with the results in Core, Guay, and Rusticus (2006) and Cremers, Nair, and Wei (2007). While this interpretation cautions against the use of these takeover-related abnormal returns to advocate for stronger governance, it is also important to note that the other positive aspects of governance shown in these two papers, specifically with regards to improved fundamental accounting performance, is unaffected by this. Finally, these results provide further direct support that the takeover factor indeed captures cross-sectional differences in takeover vulnerability.

## 4. Cross-Sectional Pricing

### 4.1 Methodology

In cross-sectional regressions (CSRs), we investigate whether the takeover factor is priced in addition to the market, size (SMB), book-to-market (HML), and momentum factors that together form the empirically successful four-factor

[27] The reduction in alphas on extending the time period is also documented by CN.

**Table 7**
**Correlation of pricing factors with the takeover factor**

| | Market | SMB | HML | UMD |
|---|---|---|---|---|
| Panel A: Time-series correlation of the factors | | | | |
| SMB | 18.06% | | | |
| HML | −53.04% | −42.10% | | |
| UMD | −14.44% | −8.54% | 6.26% | |
| Takeover | −31.84% | −10.27% | 50.54% | −33.83% |
| Panel B: Correlation matrix of the multivariate betas | | | | |
| SMB | −24.49% | | | |
| HML | 37.41% | −24.13% | | |
| UMD | 3.60% | 10.15% | −9.87% | |
| Takeover | 19.86% | 30.01% | −0.63% | 73.38% |

Panel A provides the times correlation among the factors in the four-factor Fama-French (1992) and the Carhart (1997) model (the market, SMB or size, HML or book-to-market, and the UMD or momentum factors) with the takeover factor (based on quintile-sort on the takeover likelihood, buying firms with the low likelihood of being taken over, and shorting firms with the low likelihood of being taken over between 1981 and 2004). Panel B gives the correlation between the multivariate betas on these factors for the 100 size and book-to-market-sorted portfolios.

model (Fama and French, 1992, and Carhart, 1997). The main econometric approach we use is the two-stage CSR. In the first stage, the multivariate betas are estimated using ordinary least squares (OLS). The second stage is a single CSR of average excess returns on betas estimated with generalized least squares (GLS).[28] GLS in the second stage provides improved asymptotic efficiency (Shanken, 1992) and robustness to proxy misspecification (Kandel and Stambaugh, 1995). Following Shanken (1992), the second-stage standard errors are corrected for the bias induced by sampling errors in the first-stage betas. The two-stage CSRs test whether the takeover factor can explain differences in the cross-section of returns (i.e., whether there exists a positive and significant coefficient on the takeover betas in the second-stage regression).

In addition, we test our econometric specification using the Hansen and Jagannathan (1997) distance (HJ-dist) and the J-GMM tests (e.g., Cochrane, 2004). Hansen and Jagannathan demonstrate how to measure the distance between a true stochastic discount factor that prices all assets and the one implied by the asset pricing model. If the model is correct, the HJ-dist should not be significantly different from zero, using the statistical test developed in Jagannathan and Wang (1996).[29]

### 4.2 Results for the 100 book-to-market/size-sorted test portfolios

Table 7 presents the correlation matrix of the factors used to explain the cross-section of equity returns (panel A), as well as of the multivariate betas on

[28] Results are generally robust to using OLS in the second stage.

[29] The *p*-values of the *J*-statistics from optimal GMM estimates of the models are not reported here, but exhibit a pattern similar to the HJ statistics.

these factors (panel B) for the 1981–2004 period.[30] A few observations can be made at this point. First, the correlations among the SMB, HML, and takeover factors are fairly high. Of particular interest is the positive correlation between HML and takeover (50.54%, see panel A). This may raise two concerns—that any detected importance of the takeover factor might be spuriously due to this correlation, or that a cross-section based on book-to-market will handicap the takeover factor relative to the book-to-market factor. To alleviate such concerns, we will investigate the performance of the takeover factor in the CSRs when the HML factor is excluded. As an additional robustness test, we also form an alternative set of test portfolios based on takeover vulnerabilities. Finally, we note that the cross-sectional correlation of the HML and takeover betas has the opposite sign and equals only −0.63%. The cross-sectional correlation of the UMD (momentum) factor with the takeover beta is rather large, about 73%, again the opposite sign of their time series correlation of about −34%.

We first focus on the 100 portfolios based on decile sorts of book-to-market and size and report the importance of the takeover factor in various specifications.[31,32] The annualized coefficients from the second-stage CSR are presented in Table 8. Panel A uses the data for 1981–2004 from the logit estimation of Table 2 (panel A). Panel B presents results for the takeover factor constructed using 10-year rolling estimation windows for 1991–2004.

Model 1 in panel A of Table 8 is the benchmark four-factor model. As is well known, the Fama-French factors are priced and the model generates an $R^2$ of 14.54%.[33] Model 2 adds the takeover factor. Consistent with our theory, we find that the takeover factor is important in explaining cross-sectional differences in equity returns. The annual risk premium associated with this factor is rather high and equals 8.00% ($t$-stat = 3.05). However, it is useful to note that the average beta on this factor is only 0.05. Thus, the average annualized risk premium associated with this factor is much lower and is equal to 0.4%. It is also striking that the $R^2$ of the regression significantly increases to 34.35%.[34] Finally, the HJ-dist shows that pricing errors decrease substantially after the

[30] Since the cross-sectional betas are from a multivariate regression, these betas incorporate the time-series correlation structure between the factors, and are also specific to the asset pricing model employed. The univariate betas, which we do not use, would have a correlation structure that would be much more similar to the time-series correlation of the factors. The multivariate beta correlation matrix reported here is for the model including all five factors and using the 100 book-to-market and size-sorted portfolios.

[31] We thank Ken French for making the returns on the 100 book-to-market/size portfolios available on his website.

[32] We also use 25 portfolios instead of 100 based on these characteristics. The results are statistically significant in three of the four models. For the 25 book-to-market/size portfolios, with the Fama-French four-factor model, the takeover factor is not significant, perhaps due to lack of variability that is not explained by the HML factor.

[33] The computed $R^2$ are using GLS with a constant. The significance of the takeover factor is robust in models without a constant, which are available on request.

[34] Since the Fama-French model does not accurately price small- and high-growth stocks, we check if the performance of the takeover factor is driven by these extreme portfolios. We remove from the cross-section of 100 portfolios those 5 portfolios that correspond to the smallest size decile and highest growth (below the median book-to-market). Our results, available on request, are robust to this. Our results are also robust to removing all 10 portfolios of the smallest size decile.

**Table 8**
**Cross-sectional regressions using 100 book-to-market/size-sorted portfolios**

| | FF4 | FF4 + takeover | CAPM | CAPM + takeover | FF4 | FF4 + takeover | CAPM | CAPM + takeover |
|---|---|---|---|---|---|---|---|---|
| | Panel A: 1981–2004 | | | | Panel B: 1991–2004, rolling estimation | | | |
| Constant | 0.18 (8.36) | 0.17 (7.49) | 0.20 (9.84) | 0.19 (9.00) | 0.16 (8.29) | 0.15 (7.36) | 0.16 (9.01) | 0.15 (8.28) |
| Market | −0.11 (2.84) | −0.10 (2.53) | −0.12 (3.23) | −0.11 (2.97) | −0.07 (1.70) | −0.06 (1.45) | −0.07 (1.74) | −0.07 (1.56) |
| SMB | 0.02 (0.69) | 0.02 (0.69) | | | 0.04 (1.02) | 0.04 (1.05) | | |
| HML | 0.05 (2.07) | 0.05 (2.06) | | | 0.04 (1.17) | 0.04 (1.16) | | |
| Mom | 0.11 (2.33) | 0.11 (2.22) | | | 0.00 (−0.04) | 0.00 (0.02) | | |
| Takeover | | 0.08 (3.05) | | 0.07 (2.90) | | 0.13 (2.88) | | 0.12 (2.70) |
| $R^2$ | 14.54% | 34.35% | 5.20% | 13.39% | 11.06% | 18.58% | 4.03% | 4.13% |
| H-J statistic | 0.69 | 0.60 | 0.76 | 0.67 | 0.79 | 0.75 | 0.84 | 0.80 |
| | 0.37% | 24.77% | 0.00% | 1.72% | 24.80% | 51.10% | 7.70% | 18.90% |

We report the results for various cross-sectional GLS regressions of mean excess returns of the 100 book-to-market/size-sorted test portfolios (from the whole CRSP/Compustat universe) regressed on their factor-betas. The multivariate factor-betas are estimated in a time-series regression of each test portfolio on a constant and the particular factor. In panel A, we use the takeover factor estimated from the 1981–2004 logit estimation in panel A of Table 2. In panel B, we use the takeover factor estimated using 10-year rolling logit estimation windows, such that factor returns can be calculated out of sample for 1991–2004. For each model, we report the coefficients in the first row and their *t*-statistics below in parentheses, where standard errors are adjusted for the estimation risk in the betas (see Shanken, 1992) plus the $R^2$ and the Hansen-Jagannathan statistics and their asymptotic *p*-values. The other included factors are the market (VW CRSP index), SMB, HML, and Mom (the Carhart momentum factor).

takeover factor is included. For the four-factor model, the test of zero pricing errors is still roundly rejected ($p$-value $= 0.37\%$), but for the five-factor model, it is not rejected at conventional levels ($p$-value $= 24.77\%$).[35]

To ensure that our results are indeed not driven by the correlations of the takeover factor with the other factors, especially with the book-to-market (HML) factor, we test an additional model. Model 4 considers a two-factor model including only the market portfolio and the takeover factor. As found earlier, the coefficient on the takeover factor is positive and significant, and the associated annual risk premium remains similar (7% with a $t$-stat of 2.90). Notably, the simple two-factor model with the market and the takeover factor still generates an $R^2$ of 13.39%.

Next, in panel B of Table 8 we consider the performance of the takeover factor constructed from 10-year rolling regressions over 1991–2004. For this much shorter time period, none of the four factors in the four-factor model are significant, which indicates that this period may be too short to reliably estimate CSRs. However, in the five-factor model, the takeover factor has a large coefficient of 13% that is clearly significant ($t$-stat $= 2.88$), while the two-factor model of the market portfolio and the takeover factor gives very similar results. Therefore, the pricing ability of the takeover factor is robust to using the 10-year rolling estimation windows and the shorter time period.

Concluding, an economically motivated portfolio constructed to capture differences in firms' exposure to shocks in aggregate fundamentals and discount rates (proxied by the takeover likelihood) is important in explaining the cross-section of equity returns. The increase in $R^2$, relative to existing models that are empirically successful, is remarkably large and shows the importance of accounting for the state variables relating to a time-varying risk premium. These results show that asset pricing models should take into account the difference between variations in price of risk and variations in aggregate fundamentals (e.g., through the use of the takeover factor presented here). Finally, as the increase in the $R^2$ is primarily driven by those portfolios with the larger stocks and higher book-to-market, it seems that expected returns of large and high growth stocks are more affected by these variations.[36]

### 4.3 Takeover vulnerability: risk or characteristics?

The earlier results show that the takeover factor is important in explaining the cross-section of the stock returns even when the cross-section is formed based on book-to-market and the model includes the book-to-market factor. Given that

[35] We also computed the empirical $p$-values assuming normality as in Hodrick and Zhang (2001) using Monte Carlo simulations under each model holding exactly. Ahn and Gadarowski (2004) indicate that the small sample properties of the HJ-dist can be quite far from the asymptotic distribution and depend on the number of assets and the number of time periods. These $p$-values indicate a similar pattern as the asymptotic $p$-values.

[36] Only considering the 50 portfolios of largest size stocks, the $R^2$ of the four-factor model equals 6.5% and rises to 15.93% when the takeover factor is added. Only considering the 50 portfolios of highest book-to-market stocks, the $R^2$ of the four-factor model equals 12.63% and rises to 25.36% if the takeover factor is added. Results are not reported to save space and are available upon request.

our takeover factor was constructed using several firm-specific characteristics, this section considers the natural question of whether the CSR results are indeed because of covariance (i.e., the takeover factor is priced) or because of correlation with these characteristics (i.e., the characteristics are correlated with average returns).[37]

We investigate the cross-sectional pricing performance of the takeover factor when average portfolio characteristics are added for two sets of test portfolios. The first is the set of 100 book-to-market/size-sorted portfolios constructed by us using the full Compustat sample over 1981–2004 that was used for the logit estimation (i.e., no missing data for any of the independent variables in the logit in Table 2). The second is the set of 100 portfolios based on estimated takeover vulnerabilities from the logit in panel A of Table 2. Since this cross-section is thus not based on book-to-market characteristics, this also addresses concerns that arise from the correlation between the book-to-market and the takeover factors. For each portfolio, we also calculate the time-series average of the takeover likelihood and of each of the independent variables in the logit estimation.

In panel A of Table 9, we report the results for the same four models as before without the average characteristics. The results using the set of 100 takeover-likelihood-sorted portfolios show that the takeover factor is important in explaining cross-sectional differences in stock returns. Moreover, HML (the book-to-market factor) is not significant at all, and the $R^2$ increases substantially if the takeover factor is added.

The results for the four-factor model using 100 book-to-market/size-sorted portfolios are generally similar to the corresponding results in panel A of Table 8. However, the coefficient on the takeover factor in the five-factor model is about double in size and much more significant now that the factor is constructed from the same cross-section as the set of test portfolios. Perhaps even more interestingly, after the takeover factor is added, the HML factor is no longer significant (coefficient of 0.04 with a $t$-stat of 1.80 in the four-factor model but dropping to a coefficient of 0.02 with a $t$-stat of 0.78 in the five-factor model). This suggests that part of the pricing ability of the HML factor may be due to picking up exposure to those state variables that describe time variation in expected returns, which the takeover factor is our proposed proxy for.

In panel B of Table 9, we focus on the four- and five-factor models (with and without the takeover factor), but with average characteristics added to the CSRs. We either add only the average takeover likelihood, or the average of the full set of each of the eight independent variables in the logit.[38]

Using the set of 100 takeover-likelihood-sorted portfolios, the average takeover likelihood is quite significant ($t$-stat = 3.29) when added to the four-factor model and increases the $R^2$ from 16.87% (see panel A) to 25.31%. If the

[37] See, for example, Daniel and Titman (1997).

[38] Multicollinearity prevents adding the takeover likelihood to the full set of eight characteristics.

**Table 9**
**Cross-sectional regressions controlling for characteristics**

| | Using 100 logit-sorted portfolios | | | | Using 100 book-to-market/size-sorted portfolios | | | |
|---|---|---|---|---|---|---|---|---|
| | FF4 | FF4 + takeover | CAPM | CAPM + takeover | FF4 | FF4 + takeover | CAPM | CAPM + takeover |
| Panel A: Without characteristics | | | | | | | | |
| Constant | 0.11 (3.47) | 0.10 (2.89) | 0.11 (3.59) | 0.10 (3.05) | 0.22 (7.55) | 0.18 (5.49) | 0.23 (8.19) | 0.17 (5.57) |
| Market | −0.06 (1.31) | −0.04 (0.98) | −0.06 (1.30) | −0.04 (0.97) | −0.16 (3.80) | −0.13 (2.92) | −0.16 (3.87) | −0.11 (2.63) |
| SMB | 0.04 (1.44) | 0.03 (1.30) | | | 0.01 (0.22) | 0.01 (0.55) | | |
| HML | 0.02 (0.74) | 0.005 (0.16) | | | 0.04 (1.80) | 0.02 (0.78) | | |
| Mom | −0.01 (0.11) | 0.02 (0.48) | | | −0.15 (3.05) | −0.15 (3.01) | | |
| Takeover | | 0.12 (3.47) | | 0.11 (3.27) | | 0.17 (7.07) | | 0.15 (6.72) |
| $R^2$ | 16.87% | 34.76% | 2.73% | 18.18% | 10.03% | 27.88% | 12.71% | 28.62% |
| H-J statistic | 0.55 | 0.43 | 0.58 | 0.50 | 0.75 | 0.63 | 0.79 | 0.66 |
| | 68.18% | 99.97% | 48.60% | 96.98% | 0.01% | 8.80% | 0.00% | 2.97% |
| | FF4 | FF4 + takeover | FF4 | FF4 + takeover | FF4 | FF4 + takeover | FF4 | FF4 + takeover |
| Panel B: Including characteristics | | | | | | | | |
| Constant | 0.07 (1.96) | 0.07 (1.87) | 0.08 (1.18) | 0.09 (1.27) | −0.10 (2.32) | −0.09 (2.11) | 0.03 (0.44) | 0.02 (0.29) |
| Market | −0.07 (1.51) | −0.05 (1.18) | −0.06 (1.25) | −0.06 (1.20) | −0.13 (3.14) | −0.12 (2.91) | −0.10 (2.23) | −0.10 (2.12) |
| SMB | 0.00 (0.05) | 0.01 (0.17) | 0.02 (0.75) | 0.04 (1.10) | −0.06 (2.21) | −0.05 (1.90) | 0.00 (0.12) | 0.01 (0.24) |
| HML | 0.00 (0.09) | −0.01 (0.22) | 0.00 (0.14) | −0.01 (0.29) | −0.01 (0.29) | −0.01 (0.45) | 0.04 (1.38) | 0.03 (1.19) |
| Mom | 0.02 (0.37) | 0.04 (0.74) | 0.02 (0.33) | 0.04 (0.68) | −0.07 (1.43) | −0.07 (1.54) | −0.11 (2.10) | −0.11 (2.12) |
| Takeover | | 0.08 (2.36) | | 0.05 (1.94) | | 0.06 (2.09) | | 0.08 (2.88) |
| Logit | 2.64 (3.29) | 2.00 (2.33) | | | 12.04 (10.37) | 11.19 (8.85) | | |
| $Q$ | | | −0.01 (2.05) | −0.01 (1.70) | | | 0.01 (1.52) | 0.01 (1.72) |
| PPE | | | 0.01 (0.20) | 0.02 (0.30) | | | 0.51 (7.32) | 0.50 (7.10) |
| ln(Cash) | | | −0.03 (1.33) | −0.03 (1.16) | | | 0.06 (3.90) | 0.07 (4.00) |
| BLOCK | | | 0.04 (2.05) | 0.03 (1.54) | | | 0.10 (3.17) | 0.09 (2.85) |
| ln(Mktcap) | | | 0.05 (1.71) | 0.04 (1.60) | | | −0.07 (3.96) | −0.07 (4.05) |
| Industry | | | −0.02 (0.64) | −0.04 (1.13) | | | 0.09 (1.49) | 0.10 (1.56) |
| Leverage | | | 0.21 (1.41) | 0.17 (1.16) | | | 0.19 (1.75) | 0.20 (1.85) |
| ROA | | | −0.18 (0.95) | −0.25 (1.26) | | | 0.47 (7.14) | 0.46 (6.88) |
| $R^2$ | 25.31% | 36.11% | 29.89% | 39.20% | 45.50% | 47.60% | 55.86% | 56.37% |
| H-J statistic | 0.55 | 0.43 | 0.55 | 0.49 | 0.75 | 0.63 | 0.75 | 0.63 |
| | 68.05% | 99.96% | 67.82% | 98.13% | 0.00% | 9.52% | 0.00% | 9.03% |

We report the results for various cross-sectional GLS regressions of mean excess returns of two sets of test portfolios regressed on their factor-betas without average characteristics in panel A, and with average characteristics in panel B. The time period is 1981–2004. The multivariate factor-betas are estimated in a time-series regression of each test portfolio on a constant and the particular factors. The first set of test portfolios is the set of 100 logit-sorted, value-weighted portfolios, sorted according to the takeover likelihood from the 1981–2004 logit estimation in panel A of Table 2. The second set of test portfolios is the set of 100 book-to-market/size-sorted, value-weighted portfolios, from independent decile sorts on market capitalization and book-to-market, using the set of firms with complete information for the logit model in panel A of Table 1. The value-weighted average characteristics of each of the variables in the logit model for those firms in the portfolio are added as additional controls in panel B (see Table 1 for a description). Logit is the average takeover likelihood from the fitted logit estimation. For each model, we report the coefficients in the first row and their $t$-statistics below in parentheses, where standard errors for the beta coefficients are adjusted for the estimation risk in the betas (see Shanken, 1992) plus the $R^2$ and the Hansen-Jagannathan statistics and their asymptotic $p$-values. See Table 6 for a description of all the factors, and Table 1 for a description of the characteristics.

takeover factor is added as well, the factor is significant *albeit* with a smaller coefficient of 0.08 ($t$-stat $= 2.36$, smaller compared to panel A), but the average takeover likelihood remains significant as well. Therefore, both covariance and characteristics seem important, though they are difficult to separate. For example, the correlation of the takeover betas and the average portfolio takeover likelihood for this set of test portfolios equals 73%. Next, we use the full set of eight average characteristics, of which $Q$ and blockholdings are most significant, and find that the takeover factor remains significant (bit less so, coefficient of 0.05 with a $t$-stat of 1.94).[39] Also, the HML factor remains insignificant.

The results using 100 book-to-market/size portfolios are even more interesting. When the average takeover likelihood of each of the portfolios is added to the CSR of the four-factor model, the HML factor becomes insignificant (its coefficient of 0.04 with a $t$-stat of 1.80 from panel A reduces to a coefficient of $-0.01$ with a $t$-stat of 0.29). Therefore, this suggests that the pricing ability of the HML factor may be due to characteristics related to takeover exposure.

However, if the takeover factor and the average takeover likelihood are added to the four-factor model, both are clearly significant (the factor has an annualized coefficient of 0.06 with a $t$-stat of 2.09). Moreover, it is only the addition of the takeover factor that substantially reduces pricing errors as measured by the HJ-dist ("zero pricing error" has a $p$-value of 0% without the takeover factor but including the average takeover likelihood, and a $p$-value of 9.52% with both takeover factor and average takeover likelihood). Next, the takeover factor remains significant (annualized coefficient of 0.08 with a $t$-stat of 2.88) even if all eight characteristics averages are included.

### 4.4 Aggregate fundamentals versus discount rates

The evidence presented in this paper supports the view that firms exposed to takeovers have a higher rate of return. Our interpretation of this evidence, viewed through the theoretical framework presented, would be that takeover targets are more sensitive to aggregate fundamental shocks than to discount-rate shocks. In this section, we shed direct light on this interpretation.

To separate the sensitivity to aggregate fundamental shocks from the sensitivity to discount-rate shocks, we use the two-beta framework proposed by CV. They propose a two-beta model that captures a stock's risk by the loadings on the cash-flow beta and the discount-rate beta. They split the return on the market portfolio into two components: one component reflecting news about the market's future cash flows and the other reflecting news about the market's discount rates. A stock's cash-flow beta measures the stock's return covariance with the former component and its discount-rate beta its return covariance with the latter component.

[39] We again find that there is a severe multicollinearity. For example, the correlations of the takeover beta with the average characteristics are generally high (e.g., 71% with $Q$, 63% with blockholdings, 57% with the dummy of a takeover in that industry in the previous year, and 58% with leverage).

**Table 10**
**Cash-flow betas and takeover vulnerability**

| DR beta | CF beta | Takeover likelihood |
|---|---|---|
| 1.35 | −0.013 | 1.00 |
| 1.33 | 0.064 | 2.00 |
| 1.20 | 0.067 | 3.00 |
| 1.15 | 0.060 | 4.00 |
| 1.18 | 0.082 | 5.00 |
| −0.16 | 0.094 | 5–1 |
| −0.21 | 0.135 | 10–1 |

The table shows the estimated discount-rate (DR) and cash-flow (CF) betas for the takeover-likelihood-sorted portfolios (see the text for a description of the betas, or see CV for details). The time series used is 1981:1–2001:12. All estimated betas are significant at the 1% level and all differences are significant at the 5% level.

We investigate whether firms with higher takeover exposure exhibit a pattern of higher cash-flow betas. As before, we sort firms into portfolios based on their takeover vulnerability using the coefficients estimated in the logit regression. We form five portfolios with an equal number of firms in each portfolio and estimate each portfolio's cash-flow and discount-rate betas. As seen in Table 10, the cash-flow betas exhibit the expected trend: higher takeover vulnerability is associated with higher cash-flow betas. On the other hand, discount-rate betas exhibit a decreasing trend with greater takeover exposure. This evidence thus supports the view that takeover activity is high when aggregate cash flows are high. In fact, this view appears to shed light on the trend in discount-rate betas, as well if takeovers decrease the horizon of the equity holding (Lettau and Wachter, 2005). In any case, there is a little evidence for the view that discount-rate fluctuations, in isolation, motivate acquisition activity.[40]

It is natural to ask what fraction of the observed abnormal returns to the takeover-spread portfolio can be explained by these changes in betas. The difference between the cash-flow betas of firms exposed to takeovers and firms protected from takeovers equals 0.094 (significant at the 5% level). Similarly, the difference between the discount-rate betas of firms exposed to takeovers and of firms protected from takeovers is −0.16 (again, significant at the 5% level). Using the annualized risk premium estimates provided by CV, this would imply an expected return difference of approximately 6.13% (8.8% using decile sorts). While providing support to the view presented in this paper, such a model thus does not completely explain the abnormal returns documented in this paper either. There may be additional factors missing from the simple two-beta model. Further investigation is left for future work.

## 4.5 Out-of-sample robustness check: 1951–1979

In a final out-of-sample robustness check, we use the logit coefficients estimated over 1980–2004 but apply these to the universe of all Compustat firms over the

[40] If discount-rate shocks and cash-flow shocks are negatively, but not perfectly, correlated, it is important to consider the sensitivity of takeovers to each shock in isolation.

**Table 11**
**Abnormal returns and cross-sectional regressions, 1951–1979**

| Takeover likelihood | Mean | Alpha | $t$-stat |
|---|---|---|---|
| Panel A: Abnormal returns related to takeover likelihood, 1951–1979 | | | |
| Using announced and completed takeovers | | | |
| 1 | 7.35% | −0.60% | 0.84 |
| 2 | 9.50% | 0.54% | 0.82 |
| 3 | 12.05% | 1.89% | 2.55 |
| 4 | 13.32% | 2.76% | 3.66 |
| 5 | 17.74% | 6.28% | 5.57 |
| 5–1 | 10.39% | 6.88% | 5.14 |
| 10–1 | 11.74% | 7.37% | 4.34 |

| | 100 BM-size-sorted portfolios | | | | 100 takeover-likelihood-sorted portfolios | | | |
|---|---|---|---|---|---|---|---|---|
| | FF4 | FF4 + takeover | CAPM | CAPM + takeover | FF4 | FF4 + takeover | CAPM | CAPM + takeover |
| Panel B: Cross-sectional regressions, 1951–1979 | | | | | | | | |
| Constant | 0.03 (0.89) | 0.04 (1.00) | 0.05 (1.18) | 0.09 (2.16) | −0.05 (1.46) | −0.04 (1.05) | −0.04 (1.28) | −0.03 (0.75) |
| Market | 0.04 (0.78) | 0.03 (0.59) | 0.03 (0.62) | −0.02 (0.39) | 0.13 (3.11) | 0.13 (2.84) | 0.12 (3.00) | 0.11 (2.57) |
| SMB | 0.02 (1.00) | 0.02 (1.16) | | | 0.01 (0.37) | −0.01 (0.40) | | |
| HML | 0.04 (2.67) | 0.04 (2.77) | | | 0.03 (1.53) | 0.02 (1.16) | | |
| Mom | 9.73 (2.72) | 9.77 (2.72) | | | 3.05 (1.04) | 2.03 (0.67) | | |
| Takeover | | 0.02 (0.74) | | 0.05 (2.21) | | 0.07 (2.71) | | 0.06 (2.46) |
| $R^2$ | 27.89% | 28.26% | 1.27% | 14.20% | 17.84% | 30.50% | 10.38% | 30.77% |
| H-J statistic | 0.49 | 0.49 | 0.59 | 0.57 | 0.55 | 0.51 | 0.61 | 0.58 |
| | 75.82% | 76.46% | 4.93% | 11.47% | 20.81% | 63.05% | 1.70% | 8.87% |

Panel A reports the annualized mean, the annualized abnormal return (alpha), and the corresponding $t$-statistic of five equal-weighted portfolios that are sorted according to their takeover vulnerabilities using logit coefficients estimating the entire Compustat sample for the years 1981–2004, but applied to the entire Compustat sample for 1951–1979. The logit model used is similar to the model in panel A of Table 2, but excluding *BLOCK* and industry. We also report the annualized mean and alpha and the corresponding $t$-statistic of an equally weighted portfolio that buys firms in the highest takeover likelihood category and shorts firms in the lowest category based on quintile (5–1) and decile (10–1) sorts. The alphas are relative to the four-factor Fama-French (1992) and Carhart (1997) models. Panel B reports the corresponding CSRs analogous to Tables 6 and 7.

earlier time period of 1951–1979.[41] As we do not have acquisition data available for this earlier period, the required assumption is that the characteristics of takeover targets did not significantly change over the full 1950–2004 period. However, our logit specification for the 1980–2004 sample that we are using for this case leaves out the blockholding variable and the dummy indicating whether there was a takeover in the firm's industry the previous year, as these variables are not available over the earlier time period, but otherwise is identical to the specification of Table 2. The (unreported) logit results (using all announced and completed takeovers, though results are robust to using completed takeovers only) are similar to those reported in Table 2.

Next, we sort the universe of all Compustat firms (with no missing information on any of the logit variables and with stock return data on CRSP) into quintile and decile portfolios based on their takeover likelihood according to the fitted logit coefficients. Table 11 (panel A) presents the mean returns and alphas of the resulting equally weighted portfolios. The takeover-likelihood-spread portfolio buying firms in the highest quintile and selling firms in the lowest quintile of the takeover likelihood has an alpha of 6.88% ($t$-stat = 5.14) over 1951–1979 versus 7.37% ($t$-stat = 4.34) using decile sorts.

Finally, panel B of Table 11 shows the results for CSRs using the takeover factor (i.e., the spread portfolio based on quintile sorts on the takeover likelihood with equal weighting) for 1951–1979. For three out of the four models considered, the takeover factor seems to be priced.[42]

## 5. Conclusion

This paper considers the impact of the takeover likelihood on firm valuation. While takeovers provide profitable exit opportunities for the target shareholders, takeover activity is affected by equity market conditions.

Using a theoretical framework where the price of risk varies over time and is not perfectly related to changes in aggregate fundamentals, we show that takeover exposure is associated with expected returns. We consider two alternative motivations for acquisition activity. The first motivation for acquisitions is driven through agency problems, which are exacerbated during times of positive cash-flow shocks (the "agency" view). This causes firms exposed to takeovers to become more sensitive to shocks in aggregate fundamentals. The second motivation for acquisitions is the valuation of potential synergies (the "synergy" view). When the price of risk is low, the value of these synergies

[41] Another out-of-sample test would be to consider another country with an active takeover market, such as the UK. While that falls outside the scope of this paper, we found two papers that document in independent work that takeover-likelihood-sorted portfolios may generate abnormal returns in other countries as well (see Powell (2004); and Brar, Giamouridis, and Liodakis (2006)).

[42] When using the 100 book-to-market/size sorted portfolios and the takeover factor is added to the four-factor model, its coefficient is positive but not significant, but it is when added to the CAPM. If we use the 100 takeover-likelihood-sorted portfolios, the takeover factor is significant even when added to the four-factor model, and the HML factor is not significant.

is high and firms tend to acquire, thereby increasing the sensitivity of potential targets to the changes in the price of risk. We show that firms exposed to takeovers could have a higher or lower rate of return, depending on the relative importance of two acquisition motives. While the agency view would unambiguously suggest that firms exposed to takeovers should have a higher rate of return, the implications from the synergy view depend on the importance of the investor's intertemporal hedging demands. If such demands are important, then the synergy view would suggest that firms exposed to takeovers should have a lower rate of return.

We document several supporting results. First, we show that a portfolio that buys firms with high takeover vulnerability and sells firms with low takeover vulnerability is associated with annualized abnormal returns of 11.77% relative to the four-factor Fama-French (1992) model augmented with the momentum factor (Carhart, 1997) model between 1981 and 2004. Second, we use the returns to the takeover-spread portfolio to propose a "takeover" factor, which is related to real takeover activity and a firm's exposure to takeovers, and can largely explain the abnormal returns associated with governance-spread portfolios (GIM and CN). Further, the takeover factor explains differences in cross-sectional equity returns, and improves substantially on the four-factor model.

This paper contributes to two different areas of research. First, it contributes to the development of an empirical asset pricing model that captures state variable(s) related to a time-varying risk premium and aggregate discount-rate and cash-flow shocks. The second contribution deals with the importance of corporate governance. Many advocates of governance have cited the positive abnormal returns associated with a better governance to promote governance reform. While the conclusion that governance is associated with a better firm performance might still be correct, the paper warns against the use of these abnormal returns as supporting evidence.

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