More insiders, more insider trading: Evidence from private-equity buyouts ☆

Viral V. Acharya [a,b,c,d],*, Timothy C. Johnson [e]

[a] New York University, Stern School of Business, United States
[b] London Business School, United Kingdom
[c] Coller Institute of Private Equity at London Business School, United Kingdom
[d] Center for Economic Policy Research (CEPR), United Kingdom
[e] University of Illinois at Urbana-Champaign, United States

ARTICLE INFO

ABSTRACT

Prior theoretical research has found that, in the absence of regulation, a greater number of insiders leads to more insider trading. We show that optimal regulation features detection and punishment policies that become stricter as the number of insiders increases, reducing insider trading in equilibrium. We construct measures of the likelihood of insider activity prior to bid announcements of private-equity buyouts during the period 2000–2006 and relate these to the number of financing participants. Suspicious stock and options activity is associated with more equity participants, while suspicious bond and CDS activity is associated with more debt participants — consistent with models of limited competition among insiders but inconsistent with our model of optimal regulation.

© 2010 Elsevier Ltd All rights reserved.

1. Introduction

The unprecedented buyout wave in the first part of this decade was accompanied — if press accounts are to be believed — by an unprecedented degree of insider trading. (Section 3.2 below summarizes some of the informal studies documenting this trend.) Perhaps this is merely a matter of scale: a larger number of deals means more opportunities for insider trades. However, a plausible alternative is that a novel characteristic of the buyout wave — larger financing syndicates — played a role in fostering a greater degree of information exploitation. This possibility has not escaped the attention of industry observers: the use of larger pools of participants on both the debt and equity sides (compared to similar deals in the past) naturally means that there have been more people with advance knowledge of the deals. It almost seems like a truism to observe that having more insiders should lead to more insider trading.

Yet this hypothesis is both untested and, upon reflection, not actually self-evident. Is it really clear that 20 insiders will exploit the same information to a greater extent than would 10? Might this not lead to a greater likelihood of detection and punishment? More generally,

how is insider trading affected by the *joint* effects of competition and regulation?

Absent regulation, if insiders have the same information (e.g., advance knowledge of a takeover bid) and trading is continuous, any number $N > 1$ of informed traders will drive prices immediately to their full-information level. Holden and Subrahmanyam (1992) and Back, Cao, and Willard (2000) show this in the case of homogeneously informed risk-neutral insiders in discrete-time and continuous-time settings, respectively. Holden and Subrahmanyam (1994) and Baruch (2002) further show in discrete-time and continuous-time settings, respectively, that if insiders are risk-averse, then the effect is even stronger since the risk-averse informed trader is concerned about future price risk from uncertain noise trades. The aggressive nature of insider trading induced by multiplicity of insiders is weaker in the model of Foster and Viswanathan (1996), in which traders have heterogeneous information and therefore continue to retain some monopoly power. The result on more insiders leading to more insider trading would also obtain in a one-time exchange if insiders engaged in Bertrand competition. This could occur, for instance, if the informed players were also dealers who compete via price for limited uninformed order flow, in what seems a reasonable depiction of some markets, such as the credit derivatives market.

This intuition, however, cannot be applied to most markets, since insider trading is, in fact, regulated by explicit enforcement regimes. Whether or not a greater number of insiders will lead to more insider trading generally must depend upon (among other things) the nature of the enforcement regime and the penalty functions that insiders face. To take a simple example, suppose regulators investigate a deal if and only if the pre-deal volume of stock trades exceeds a known and fixed threshold, V, and that, conditional on an investigation being initiated, detection and (dire) punishment are certain. The equilibrium outcome is that N informed traders each trade up to V/N shares, so that total trading does not rise with N. In fact, it is not difficult to see that such an enforcement regime might even be *optimal*: commitment to a ceiling on illegal trade creates a negative externality of each insider's trade on other insiders and this makes the ceiling to some extent self-enforcing.

Optimal regulation of insider trading, in the case of a single insider, is studied by DeMarzo, Fishman, and Hagerty (1998). We develop a model that generalizes the setting of that paper to N insiders. In the decision to trade, each insider ignores the cost imposed on other insiders in approaching the enforcement ceiling. This effect induces greater insider trading and lower liquidity (unless checked by regulation). We show formally that, under a fairly general set of conditions, the optimal regulation features an enforcement ceiling that gets *stricter* as the number of insiders increases. We conclude that empirical evidence of insider trading increasing with the number of insiders is unlikely to be consistent with optimal enforcement.

The empirical part of the paper provides an opportunity to investigate these issues in a setting that is attractive along a number of dimensions. We study trading activity in stocks, options, bonds, and CDS markets over the period 2000–2006 for the interval immediately preceding buyout announcements by private-equity acquirors of public firms in the United States.[1] Prior to a bid announcement, there is a well-defined set of players that possess valuable, short-lived, non-public information. The number of informed parties has nothing to do with information production: the quantum of information is the same for all deals. Moreover, reputation considerations are also unlikely to play a large role because information can be exploited anonymously in the stock and options markets. On the other hand, insider trading is definitely illegal and subject to severe penalties for bonds, stocks, and options.

Our primary findings are that insider trading in stock and options markets is more likely if there is a larger size of equity syndicate, whereas insider trading in CDS and bond markets is increasing in the size of debt syndicates. These results imply that insider trading becomes more likely with more insiders in spite of the presence of regulation. (We stress that we do not claim that any of the entities we count is literally guilty of prohibited activity.)

Our preferred explanation concerns the *nature* of the enforcement regime. If each potential insider regards the likelihood of detection (and the probable penalty upon detection) as independent of the number N of insiders, then one would expect a rising number of informed players to result in a rising amount of illegal behavior. Our model in fact implies that such a regime is suboptimal because the harm to market liquidity from allowing more insiders to trade can be efficiently avoided by imposing an enforcement ceiling. We conclude that our evidence is consistent with recent claims in the financial press that the enforcement of insider trading preceding the leveraged buyouts of public firms has not adjusted to the evolving institutional structure of such buyouts, particularly with regard to the increasing number of insiders.

Beyond this theoretical explanation, it also seems clear that allowing the total amount of informed trading to rise with N creates dangerous incentives. To the extent that insiders can *choose* N — e.g., one can always tip off one's friends — there could be a positive net benefit to doing so. If expected individual punishment actually weakens with N, this would create an externality, making it safer for more agents to trade together.[2]

---

[1] Acharya and Johnson (2007) also examine the possibility that insider trading increases with the number of insiders by analyzing the CDS market. However, the exclusive focus on the CDS setting is not the most attractive for issues we examine. For all practical purposes, there is no regulatory effort to curb such activity in the CDS markets, and any self-regulation is achieved only through recommended practices and implicit contracts with counterparties and other insiders, e.g., syndicate banks. Acharya and Johnson (2007) hypothesized that the number of banks with access to private information about a borrower would contribute to the amount of suspicious activity, documenting supportive empirical evidence.

[2] Such a policy would entail a social dimension to insider trading, as hypothesized by Glaeser, Sacerdote, and Scheinkman (1996) and Sah (1991), raising the possibility of "crime wave" equilibria. In a recent paper, Bond and Hagerty (2005) study such possibilities, and show how particular enforcement regimes might promote them.

In highlighting the interaction of enforcement and competition, our work speaks to both the literature on the dynamics of asymmetric information, and the literature on the design and efficiency of regulation. An additional contribution of the paper is to offer a new empirical finding relating syndicate structure to insider activity. Although there is an enormous body of work documenting differences in information asymmetry across firms (and across markets), and although these differences are widely thought to have important consequences for market dynamics, there has been perhaps surprisingly little study of *why* these differences arise. Exploiting the opportunity presented by recent developments in the takeover market, we offer an initial contribution in this direction. Large firms with more bank relationships and with larger potential bidding syndicates are likely to see leakage of non-public information. Furthermore, we find that information leakage is segmented in that the relevant number of insiders for stock-based markets is based on equity participants, whereas that for credit markets is based on debt syndicate members, suggestive of localized information flows within these markets.

The outline of the paper is as follows. The next section presents our theoretical setting and formalizes the main result on optimal enforcement of insider trading with N insiders. Section 3 describes the period we examine and our sample of LBO bids. Section 4 explains our construction of the main dependent and independent variables. Section 5 presents the empirical results and considers alternative interpretations and robustness. The final section summarizes the paper and concludes.

2. Competitive insider trading under regulation

A central question in market microstructure is the behavior of informed traders under conditions of limited liquidity and competition. The question we investigate is how informed trading policies change in the presence of explicit regulatory constraints. In attempting to understand the influence of institutional structure on price determination and trading volume, such regulatory constraints should be expected to have first-order effects: after all, in most of the world's security markets there are outright prohibitions against the exploitation of material non-public information. Moreover, from a regulatory perspective, efficient design of the enforcement regime requires an understanding of the endogenous responses induced in those who possess such information.

The latter problem — efficient policy design — is the focus of this section. We describe a model for analyzing the optimal regulation of insider trading based on the framework of DeMarzo, Fishman, and Hagerty (1998), who analyze the issue when there is a single insider. We wish to understand how the solution to the problem changes when the number of insiders varies. When there are $N > 1$ insiders, we can regard insider trading decisions as being constrained both by explicit (external) regulation and by the implicit (internal) regulation imposed by competition among the insiders. We are interested in the interplay of these two forces. In particular, our goal is to characterize conditions under which optimal external regulation should become *stricter* for larger values of N.

Our model is necessarily stylized, and we highlight several dimensions along which the problem could be generalized. Given our setting, however, we are able to establish that, under relatively mild conditions, a regulator facing three states of the world — with 0, 1, or $N > 1$ informed players — should adopt an enforcement policy under which no insider trading at all is tolerated in the third state. We interpret this result as supporting the assertion that under efficient regulation, one should not expect to observe in equilibrium more insider trading with more insiders. (In a sense, this can be seen as complementing the analogous result when there is no regulation: under perfect competition and risk-neutrality, any number $N > 1$ of insiders will instantly drive prices to their full-information level.) After describing the setting and stating the result formally, we discuss the main intuition behind it.

The model involves a single risky asset (the stock) traded in a single period. There are four sets of players. First, there are uninformed liquidity traders who randomly buy or sell the stock for unmodeled reasons. Second, there are N insiders who are risk neutral, and who know the terminal value of the stock with certainty. Third, there is a risk-neutral market maker who must set bid and ask prices, b and a, subject to a zero-profit condition, and must sell whatever quantity is demanded at a and buy whatever is offered at b. An important simplification is that the market maker cannot condition these prices on the size of the incoming orders.

The fourth actor is the regulator, whose objective is to maximize the welfare of the uninformed traders, which is equivalent to minimizing the bid–ask spread (or maximizing liquidity). We defer for the moment a formal specification of the regulator's problem.

For simplicity, we assume that there are only two possible terminal stock prices, S_T. With probability p_H there is a takeover bid of value $H > \overline{S}$; with probability $1 - p_H$ the asset has a liquidation value of $(\overline{S} - p_H H)/(1 - p_H) < \overline{S}$, so the expected payoff is $\overline{S}$.

If there is no takeover bid, then there are no insiders. If there is a bid, then N takes on some value $N \geq 1$. The regulator and the insiders observe this value. If there is a bid, insiders can decide to (illegally) buy at the ask price. Clearly the bid price will never exceed H, so insiders are never sellers. (They are not paid negative fines for selling.)

We assume that the buying and selling demands of the uninformed traders are independent of one another and of prices. Since there are never any insider sales, the zero-expected-profit condition applied to the market maker's purchases ensures that $b = \overline{S}$. The spread is thus determined by the ask price, a. Denote the uninformed buying demand by Y. If the total informed buying demand is X, then the zero-profit condition applied to the market maker's sales requires $E[(X+Y)(a - S_T)] = 0$, where E denotes the expectation over uninformed buying demand, the number of insiders, and the terminal stock price. This, in turn, implies

$$a = \frac{\overline{S} + H \frac{EX}{EY}}{1 + \frac{EX}{EY}}.$$

Given this price, the nth insider determines his demand, x_n, in order to maximize expected profit net of expected penalties, while taking the policies of the other insiders as given. Since the insiders are identical, we consider only symmetric equilibria in which their demands are identical.

To further analyze the insiders' policy, we need to specify the form of the penalty and detection functions that they face as part of the regulatory environment. This, in turn, requires us to define the regulator's problem. Note that the regulator's objective of minimizing the bid–ask spread is equivalent to minimizing a, which is equivalent to minimizing expected informed trade, because the expression for a given above is increasing in EX.

The regulatory problem is to decide when to undertake costly enforcement actions, subject to a budget constraint. We assume that, following the realization of the stock's value, the regulator can choose to investigate all trading for a cost equal to $c_0 + c_1 N$. If an investigation is undertaken, insider trading is detected and punished with certainty. Note that having costs increase with N gives the regulator an incentive to devote fewer resources to enforcement when there are many insiders.

We assume that the punishment incurred upon detection by any insider who trades is disgorgement of his trading profit plus a fixed penalty, P_0. While endogenizing the penalty is an interesting direction for generalization, it is not unreasonable to view it as determined by statutes and courts, and therefore outside the control of the regulatory agency. The choice of a fixed penalty is important for our results. In practice, insiders subject to prosecution can also face a variable component, which we could accommodate as well.[3] However, it seems realistic to model even modest violations as incurring large fixed costs, if for nothing other than mounting a criminal defense and incurring reputational damage.

In deciding when to undertake an investigation, the regulator is subject to the resource constraint that the policy's expected cost must be no more than some fixed budget K. The background assumption is that the regulator oversees many repetitions of the game, so that the constraint need not hold ex post for every event. A related assumption is that the regulator can and does commit ex ante to its enforcement policy. Undertaking a potentially costly investigation may not be worthwhile ex post. But commitment is necessary for the threat to be credible. Moreover, the commitment itself is valuable. If the government could renege on the enforcement budget in high-N outcomes, this would induce a positive externality of insiders' trades on each other. Bond and Hagerty (2005) analyze how this scenario can generate "crime wave" equilibria.

In choosing its enforcement policy, the regulator can only condition upon the observed variables N and the total volume of buy orders $Z=X+Y$. In a similar setting, DeMarzo, Fishman, and Hagerty (1998) establish that, if

the distribution of Y satisfies the monotone likelihood ratio property (MLRP), then it suffices to restrict attention to non-random policies that specify investigation if and only if Z exceeds some threshold V. (In this context, MLRP is equivalent to log concavity of the density function of Y.) For that reason, we consider such threshold policies. The problem then boils down to the optimal choice of V. In general, this ceiling will depend on the distribution of noise trader demand and (via the budget constraint) the distribution of N, as well as on the outcomes N and Z.

Return now to the insiders' problem. With enforcement triggered when $Z > V_N$, insiders are detected when there is an unexpectedly large amount of uninformed demand. Specifically, the probability of punishment for each of the insiders is $1 - F(V_N - \sum_{j=1}^{N} x_j)$ where F denotes the cumulative density function of noise trader demand Y. The ith insider's objective (expected payoff) if he trades x_i shares is thus

$$(H-a)x_i F\left(V_N - \sum_{j=1}^{N} x_j\right) - P_0\left[1 - F\left(V_N - \sum_{j=1}^{N} x_j\right)\right]. \tag{1}$$

(If he trades no shares, his payoff is zero.) So, conditional on trade, his first-order condition can be written

$$x_i + \frac{P_0}{(H-a)} = \frac{F'}{F}, \tag{2}$$

where the argument to F and F' is the same as in the previous expression.

We now introduce a convenient parametric assumption. The distribution function of uninformed demand is assumed to belong to the logistic family.

Assumption 1.

$$F(y) = \frac{1}{1 + e^{-(y-m)/s}},$$

where $E(Y) = m$ and s is a scale factor. This distribution satisfies the MLRP. It also has the property that F'/F is invertible:

$$\frac{F'}{F}(y) = w \Rightarrow y = m + s \log\left(\frac{w-s}{s}\right).$$

This handy fact allows us to characterize insiders' optimal demand given the outcome N and the enforcement ceiling. It further yields an important characterization of the participation decision. That is, it yields conditions under which the insider will choose not to trade at all. These results are given in Appendix A. Note that the insiders' problem and the market maker's problem are coupled. The ask price itself depends on the optimal trade demand and hence on the participation decision in each state.

Given the solution to these two problems, the regulator then computes the expected cost of any policy $V(N)$ as

$$E[C(V(N))] = E[(1 - F(V(N) - Nx^*))(c_0 + c_1 N)],$$

where x^* denotes $x^*(V(N))$, the optimal demand of an individual insider, and the expectation is taken over the distribution of the number of insiders as well as over noise trader demand. The regulator then chooses the

---

[3] For example, under U.S. law, the S.E.C. or private plaintiffs can pursue civil penalties up to a maximum that is proportional to the illegal profits. Insiders are also subject to criminal prosecution, under which punishments have a fixed monetary ceiling and can also include jail.

enforcement ceiling that minimizes total insider trade, such that the overall policy satisfies the budget condition in expectation:

$$\min_{V(N)} E[Nx^*|V(N)] \quad \text{such that } E[C(V(N))] \le K.$$

Completely solving the equilibrium analytically is infeasible. However, it is intuitively clear that, with enough resources, the regulator will be able to set the enforcement ceilings low enough to deter all insider activity in all states. Likewise, as the regulator's budget is reduced, it could still turn out to be optimal to loosen enforcement in some states while maintaining zero-trade policies in other states. The heart of our results is that, in fact, it is optimal to maintain a policy that for high N, all insider trading activity is deterred.

To be specific, and to further simplify, assume that the number of insiders only be 1 or N. In that case, we can think of regulatory policies as points in the $V_1 \times V_N$ plane. The optimal responses, a and (x_1, x_N), of market makers and insiders, respectively, can then be seen as partitioning this plane into four regions depending on the insiders' participation/non-participation.

- Zone I is the set of policy points for which $(x_1=0, x_N=0)$.
- Zone II is the set for which $(x_1 = 0, x_N > 0)$.
- Zone III is the set $(x_1 > 0, x_N = 0)$.
- Zone IV is the set $(x_1 > 0, x_N > 0)$.

What we aim to establish is that the optimum usually resides in Zone III. More specifically, it resides on the Zone III side of the border with Zone IV. The meaning of this is that it will be optimal for the regulator to pick the enforcement ceiling V_N to be the highest value that is low enough to force the insiders not to trade when there are $N > 1$ of them.

To continue the simplification, we make an explicit assumption about the distribution of the number of insiders.

Assumption 2. Conditional on a takeover bid, the number of insiders assumes the values 1 or $N > 2$ with equal probability.

Next, define some intermediate quantities: $\pi \equiv P_0/(H-\overline{S})$ and $\theta \equiv \pi p_H/2m$. Note that these are both positive. We require one parametric assumption that we explain below.

Assumption 3. Assume $e^{N-1} > 1 + 1/\theta$.

We now state the main result.

Proposition 1. *Under Assumptions* 1–3, *for any budget amount K such that perfect enforcement is not achievable, the optimal policy lies in Zone* III.

The proof of the proposition is in Appendix A, which also describes a solution algorithm enabling explicit computation of the equilibrium.

In effect, the proposition says that the regulator achieves maximum market liquidity by having no tolerance for "crime sprees." Two effects drive this result. First, and most important, the nature of an enforcement ceiling

on illegal activity is that it creates a limited amount of available profit, which the insiders must split. When that available profit — roughly $(H-\overline{S})V/N$ — is less than the expected penalty, all insiders obey the law, creating a discontinuous jump in the regulator's objective. Second, when enforcement is relaxed away from this benchmark, insiders trade more aggressively when there are more of them. As noted by DeMarzo, Fishman, and Hagerty (1998), when there are multiple insiders, each internalizes less than the full cost (in terms of expected punishments) of the decision to trade more. A consequence of this externality is that raising the enforcement ceiling when N is large results in a greater increase in insider trading than the same increase in the ceiling when N is small. This means that the regulator's objective function is harmed more in the former case, which favors solving the problem by being more lax for outcomes with fewer insiders.

The Assumption 3 imposed in establishing the proposition ensures that these effects are reasonably large. When θ is very small, it becomes harder to show their influence. To see why, write it as

$$\theta = \frac{1}{2} p_H \left(\frac{\pi}{s}\right) \left(\frac{s}{m}\right).$$

This is small when the likelihood of the high-insider state or the probability of a deal are very small (the first two terms). It is also small if punishments are small relative to available profits.[4] A little more subtly, θ is also small when the signal-to-noise ratio in the uninformed trading demand, m/s, is large. When volume is highly predictable (absent a bid), insider trading is riskier. In that case, the externality effect described above is also weaker. Notice that the condition on θ only matters in comparison with the number of insiders in the competitive state.

While the condition is not entirely without loss of generality, it is not stringent in terms of practical numbers. If s/m is the ratio of unexpected to expected volume, a typical daily number might be 0.25. If the regulator ensures that punishment conditional on apprehension is 10 times the average level of insider profit ($\pi/s = 10$), and if the probability of a bid is 1%, then the condition just requires $N > 4$. Moreover, it is important to point out that the condition is far from necessary. It simply provides a succinct way to bound some complex expressions in the proof (Appendix A points out some alternative conditions that might be weaker in some situations). Thus, one could not reverse the assumption and prove the opposite of the proposition, namely, that weaker enforcement with high N is optimal.[5] It is true qualitatively, however, that as the factors in the θ expression shrink or if N is small, the regulatory benefit of the optimal policy becomes harder to discern.

Finally, we note that the assumptions that 1 and N are equally likely outcomes and that $N > 2$ are not essential and are only imposed for simplicity.

---

[4] The Appendix shows that insiders only trade when they will purchase at least s shares. Hence $\pi/s = P_0/(s(H-\overline{S}))$ is a risk-reward ratio to the marginal insider.

[5] We thank the referee for raising this possibility.

To summarize, we have presented a stylized model of the optimal regulation of insider trade when the number of insiders can vary. The model implies that regulation should be more stringent when the number of insiders turns out to be large. The analysis uses some simplifying assumptions, including exogenous penalties, non-discretionary noise trade, and no conditioning of execution prices upon orders. While relaxing these assumptions presents interesting avenues for further research, the case we study still reveals some important economic mechanisms. The main insight is that regulatory efficiency entails tougher enforcement when individuals internalize less than the full marginal punishment cost. This observation frames our empirical work below, which investigates whether observed outcomes are consistent with this notion of regulatory efficiency. In particular, we focus on suspicious trading activity preceding the public announcements of leveraged buyout transactions, which over time have increasingly featured a large number of participating institutions.

3. Background and sample

This section describes the setting for our study. We begin with an overview of some of the industry developments that characterized leveraged buyouts during the sample period 2000–2006.

From 2003 through the second quarter of 2007, there was a dramatic rise in the merger and acquisition (M&A) activity around the world. From a low of $1.2 trillion in 2002, the pace of activity increased to $3.7 trillion by the end of 2006 (according to Thomson Financial Services, as of September 5, 2006). In 2005, there were 200 buyouts in the United States with an aggregate value of $850 billion; the corresponding numbers for Europe were 1300 buyouts and 125 billion euros. Compared to the merger boom of the late 1980s, which was financed primarily by public equity and junk bonds, this merger boom was primarily driven by the availability of syndicated bank debt and the tremendous growth of private-equity funds. Like the buyout wave in the 1980s, deal prices also escalated. See, for example, Acharya, Franks, and Servaes (2007), who report that the growth in LBO transactions during the period 2001–2005 matches the growth from 1985 to the peak of 1988–1989, and that the EBITDA to valuation ratio for targets, relative to the market average, declined to a low of 4% in 2006, as compared to 5% in 1989. With this secular increase in the volume and number of LBO transactions, there have been some important developments in the nature of institutional participation in their equity and debt financing. We review these next.

3.1. Broadening of participation in debt and equity syndicates

The syndicated loan market became a major source of deal financing during the 2001–2006 M&A boom. In 2006, the $233 billion of LBO deal volume in the United States was funded in part by about $125 billion of such loans (Altman, 2007). Broader figures for overall debt issuance show that there was a surge in syndicated debt financing relative to corporate bond issuance. In 2001, both these issuances were around $1.5 trillion, whereas in 2005, syndicated debt financing had grown to be about twice as large — a total of $3.75 trillion relative to corporate bond issuance of $1.75 trillion (according to Merrill Lynch Research based on Dealogic database).

The growth of the syndicated debt market was made possible in part by a deepening of the number of participating lenders beyond the traditional large commercial banks. According to Reuter's Loan Pricing Corporation, institutions other than banks, including hedge funds and other "alternative investment" vehicles, assumed more than 60% of loans issued in 2005. As participation in initial syndicates expanded, so too did the (previously rare) practice of secondary market trading. Important to note for our study are the ramifications of wider participation in initial and secondary markets have for the flow of information. Holders of any stake in a syndicated loan are entitled to all the non-public information gathered by the lead banks in their capacity as monitors of the borrowing firm.

A related and equally important development in the recent buyout wave was the increase in the syndication among private-equity firms on individual deals. Twenty-one "club" deals — involving more than one acquiror — were announced in 2006, valued at $176.5 billion, double the amount in 2005.[6] As deals grew larger, the portfolio diversification motives of private-equity pools essentially ruled out complete equity funding by a single firm. This was especially true in 2005 and 2006 as deals started expanding to relatively large (over $1 billion in value) public firms. Such "clubbing" of deals greatly expanded the universe of people privy to the negotiation process leading up to the launch of a buyout bid.

3.2. An increase in insider trade?

The broadening of participation in debt and equity syndicates, and the increasing role played by hedge funds, is widely believed to have increased the incidence of insider trading in a number of different markets prior to buyout announcements. The reason for this belief seems to be well founded. The list of insiders on deals now includes bidders, investment bankers, lawyers, and lenders, as well as the management of the target company. The TXU Corp. bid of 2007, for example, involved 7 investment banks and 12 law firms. As the pool of people with inside information expands, the potential for inappropriate use of material non-public information clearly increases. Also, when public companies attract interest from would-be acquirors, they often sound out other potential buyers or conduct confidential auctions in search of better prices, further swelling the circle of insiders. It has been well recognized in the media that the increasing size of LBOs and the increase in the number of

---

[6] Cogut, C., Sorkin, D., Sudol, K., "Global Overview", in Simpson Thacher & Bartlett LLP, "Getting the Deal Through", *Private Equity 2007*.

participants are perhaps responsible for the recent surge in insider trading prior to LBO announcements.[7]

Our paper does not directly address the question of whether insider trade actually *did* increase (relative to earlier periods) during 2000–2006. However our investigation is certainly motivated by the perception that it did. This perception is not purely anecdotal. In fact, it finds support in a number of non-academic studies across different markets.[8]

3.3. Changing incentives?

The description above (and our analysis below) suggests that insider trading could have been especially prevalent in the recent buyout wave because of broader participation in takeover activity.

As such, insider activity would be expected to respond to changes in rewards. As noted above, private-equity firms did appear, by some measures, to pay higher and higher prices as the buyout wave progressed. However, from a historical perspective, the rewards for advance knowledge of a bid do not appear to have increased. In our sample of bids (described below), the average six-day return (from $t-5$ to t, where t represents the announcement date) to target stocks was 13.2%. By comparison, Andrade, Mitchell, and Stafford (2001) report average three-day returns of about 16% for all mergers during the 26-year period 1973–1998. (The latter number is a risk-adjusted abnormal return, the former number is a raw return.) That average was virtually constant across decades, and was more than 20% for cash-only deals (i.e., deals comparable to the buyouts we study). Thus, if anything, the rewards to advance trading seem to have been lower during the period 2000–2006.

An alternative hypothesis, motivated by our theoretical analysis, is that the rise in such trading was due to a laxer enforcement climate.

To the extent that insider trading takes place in new over-the-counter derivatives markets such as that for CDS instruments, the assertion is certainly true. As mentioned in the introduction, there are few (if any) laws against such trading in any jurisdiction. And, in the United States, there was not even a clear regulator during our sample period with purview over credit derivatives.

In markets that are explicitly regulated, governments have also faced other new complications. One issue that has made enforcement difficult has been the rise of cross-border trading. New institutions also complicate monitoring. Hedge funds, in particular, are more opaque and less subject to the responsibility of protecting non-public information (via "Chinese walls").

Despite the challenges, and despite the perception among some participants that enforcement has been lenient, we know of no direct evidence that regulators have achieved less success during the period of our sample. Hence, we rely on our model's implications for inference regarding the enforcement climate. In particular, we investigate the link between measures of suspicious trading and the number of insiders for such inference (Proposition 1).

3.4. Sample

We construct a sample of buyouts of public companies from January 1, 2000 to December 31, 2006. We do not extend the sample backwards because, as described above, merger and acquisition activity in the preceding decades was markedly different along a number of dimensions.

Our data come from Thomson Financial (formerly SDC) and consist of bid events. We select bids by private financial buyers of public companies for which the value of the bid exceeds $100 million. We impose a few other selection criteria (described in Appendix B) whose aim is to select private-equity buyouts rather than ordinary acquisitions by (possibly private) operating companies or subsidiaries.

We do not have direct information on the formal structure of a proposed acquisition or its anticipated capital structure. So we cannot necessarily describe all our bids as "LBOs". (There is a field in the database flagging events Thomson determines to be LBOs, but the procedure for the designation is not clear; all bids so described are in our sample, but we also include several bids without this designation.) We also do not require that the bid necessarily be successful or completed. As of November 2007, only 60% of our bids had resulted in a completed deal (with about half of the rest having resulted in acquisition by a different bidder).

Table 1 presents a summary of the number of bids and their size (10th percentile, median, and 90th percentile) year by year for the bids that meet our initial selection criteria. The most striking feature of the table is the rise in number of transactions in 2006 (81 deals) and the substantial increase in the size of deals after 2003. The median transaction size is around $200–300 million until 2003, but exceeds $600 million thereafter. Note that the 10th percentile sizes also exhibit such a trend but less dramatically, capturing the fact that there has been a surge in very large transactions since 2003. The smallest number of transactions as well as the smallest median

---

[7] "How many people can you have knowing a secret and keep it a secret?" asks John Coffee, a securities law expert at Columbia University in New York. "Under about 10 people. I think Wall Street can keep a secret. But much beyond that, I don't know." ("Insider Trading," Bloomberg Markets, August 2007.) In fact, Keown and Pinkerton (1981), studying the abnormal stock price reactions prior to public M&A transactions, recognized this possibility quite early in the literature: "You start with a handful of people, but when you get close to doing something the circle expands pretty quickly. You have to bring in directors, two or three firms of lawyers, investment bankers, public relations people, and financial printers, and everybody's got a secretary. If the deal is a big one, you might need a syndicate of banks to finance it. Every time you let in another person, the chance of a leak increases geometrically."

[8] See "Boom time for suspicious trades," *Financial Times*, August 6, 2007, for a study documenting that suspicious trading occurred in equity markets ahead of 49% of all North American deals between 2003 and 2007; "Insider Trading Concerns Rise as Stock Options Surge", Bloomberg, May 7, 2007, which showed that pre-bid volume in equity options jumped 221% compared to the average for the prior 50 days for the 17 biggest U.S. takeovers in 2006; and "Moving the Market – Tracking the Numbers," *Wall Street Journal*, December 14, 2006, citing a study by a firm called Credit Derivatives Research, found unusual spikes in CDS fees ahead of news or reported rumors concerning 30 LBOs in 2006.

Table 1
Deal size ($ millions) by bid year.

Year	# of obs.	10th %ile	Median	90th %ile
2000	28	122	257	2175
2001	5	133	256	1995
2002	15	115	314	1464
2003	10	109	244	443
2004	15	401	1208	2931
2005	41	197	605	2344
2006	64	357	1496	17 074

The table presents the year-by-year averages of the number of bids and the 10th percentile, median and 90th percentile of proposed deal value ($mil). The overall sample consists of 178 bids over the period 2000–2006 from Thomson Financial (SDC) database. Sample selection criteria are described in Appendix B.

Table 2
Target characteristics.

	#Obs	10th %ile	Median	90th %ile
Market cap	178	97	446	3321
M/B	170	0.87	1.64	3.39
D/V	175	0.00	0.30	0.61
β	178	0.11	0.78	1.61
σ	178	0.22	0.43	0.79
Stock volume	178	0.02	0.20	1.51
ILLIQ	178	0.0005	0.0077	0.3170
Turnover	178	0.15%	0.59%	1.53%
S&P credit rating	81	BBB-	BB	B

The table presents the 10th percentile, median, and 90th percentile of firm characteristics for the part of our buyout sample that matches with CRSP and Compustat. The measures are based on the six-month (calendar) pre-announcement period. They are calculated per firm per day, and then averaged across the six-month period for each firm. Illiquidity, stock turnover, firm size, market beta, and volatility are based on items reported in CRSP. Market to book, leverage, and S&P credit rating are based on items reported in Compustat. Market cap is equity market value in millions of dollars. M/B is equity value divided by the book value of common equity. D/V is the book value of long-term debt divided by the sum of this value and the market value of equity. σ is the annualized volatility of daily stock returns. β is estimated with respect to the CRSP value-weighted index. Stock volume is in millions of shares per day. ILLIQ is measured using the Amihud (2002) ratio, computed daily and averaged over the entire period, and is in units of 10^{-6}. Turnover is in percent per day. S&P credit rating is averaged for each firm over the six-month period.

size were in 2001, which coincides with a recession in the United States along with historically high default rates. This reflects the somewhat cyclical nature of flows into the buyout industry, a fact that has been reinforced by the virtual drying up of buyouts following the credit market squeeze in the summer of 2007.

Some other aspects of the sample deserve mention. Consistent with our sample being buyouts, the median acquiror bid is for the entire 100% of the target equity. Almost all the acquirors are private-equity houses or their consortia, with a few exceptions reflecting stakes by individuals and venture capital firms. The data span a large number of industries, the decomposition by value being: services (SIC 70–89) 32%; manufacturing (SIC 20–39) 26%; retail trade (SIC 52–59) 14%; finance, insurance, and real estate (SIC 60–67) 13%;

transportations and public utilities (SIC 40–49) 8%; wholesale trade (SIC 50–51) 4%; and mining (SIC 10–14) 1%.

Different subsets of this overall sample are used in different tests below, depending on the additional data needed and the market being investigated.

Out of our initial sample, 212 targets match with Center for Research in Security Prices and Compustat data. Out of these 212 targets, 178 had active syndicated loan data on Loan Pricing Corporation's DealScan database as of the bid date. These targets together had 857 loan facilities with a median maturity of 3.5 years and a median facility size of $135 million. The median number of lead banks and participating banks in these loans is five and six, respectively. The 10th percentile values are 0 and 1, respectively, whereas the 90th percentile values are 22 and 12, respectively. Summary equity characteristics of these 178 firms for the six-month pre-announcement window are provided in Table 2. The median target size is $446 million. Median stock turnover is 0.59% per day. Target firms in general tend to have low leverage (median debt to firm value ratio is 0.30 with 10th percentile leverage being zero) and are from stable sectors (median beta of 0.78). The tests below employ this 178 Thomson-CRSP-Compustat-LPC matched sample.

4. Empirical strategy

This section describes our methodology for testing for a link between the financing structure of a takeover bid and the likelihood of insider trading prior to that bid. Our first step is to construct measures of suspicious pre-bid trade for each event. These then become our dependent variables in the main regressions, which utilize measures of the number of informed insiders as the primary independent variables.

4.1. Measuring insider activity

Insider trades can only be directly measured with detailed transaction data and knowledge of the informed status of all traders. While this information can be obtained by government investigators, even they do not have the resources to gather it systematically for large samples. Instead, typical monitoring relies initially on broader statistics that can be indicative of suspicious activity. We take that approach here, constructing statistics that we postulate to have a monotonic relation with insider activity. To be precise, we postulate that the relation is monotone *given that a bid did occur. There is no assumption that the measures are unconditional predictors of insider trading, or of imminent bids.* Meulbroek (1992) presents direct evidence that pre-bid trading by insiders results in abnormal price and volume changes. While our measures flag unusual trading activity in a number of ways, we have no way of ascertaining the degree of effectiveness of any one of them in truly identifying illegal activity. As noted by Jarrell and Poulsen (1989), unusual activity can precede takeover bids for other reasons, including the building of "toeholds" by the acquiror and speculation based on public information. To the extent

that all of our measures are noisy, the methodology is biased against being able to detect any association between the suspiciousness of trade and our explanatory variables.

It is also worthwhile to point out that our statistics could, in principle, be measuring two distinct things: the *likelihood* of (some) insider activity prior to a particular bid and the *amount* of (all) such activity. Some of our measures might be more sensitive to one than the other. But, as a practical matter, we have very limited ability to distinguish which (if either) we are capturing.

4.1.1. Stock market measures

All of our target companies had publicly traded stock prior to the bid announcement date. For each, we construct measures of unusually heavy trade or unusually large positive price movement in a five-day window immediately preceding the bid. Our methodology consists of two stages. First, we design a regression specification to describe "normal" (or expected) values of each series (volume and returns). We run this regression using daily data for a three-month period preceding the bid. Second, we apply a metric to the regression residuals in the pre-event window to flag the occurrence of suspicious activity on any individual day. There is no single best way to do each step, so we try a few alternatives.

To provide an unconditional variant and a conditional variant, we use two regression specifications, one with just a constant and the other with a constant, lagged volume and returns, day-of-week dummies, and contemporaneous volume and returns for the market index.

Volume and return data come from CRSP. We use the CRSP value-weighted return for the market return, and the S&P500 volume for market volume.

Notice that the last specification includes contemporaneous information. The purpose of these measures is to describe returns and volume given *all* information about the date in question, whether or not the information was known prior to that date. More detailed specifications could include dummies for earnings announcements or other news events. It turns out that our results are largely insensitive to the specific variables chosen.

Given these regressions, we use two functional forms to capture the presence of large positive residuals in the five trading days before the bid: MAX, the maximum of the daily standardized residuals, and SUM, the sum of the positive standardized residuals. The first measure is sensitive to unusually large individual days and the second is sensitive to cumulatively large abnormal trading. The SUM measure is very close to the cumulative abnormal return measure (CAR) employed in much of the event study literature. (Restricting the sum to positive residuals yields a number that is analogous to an *F*-statistic for the test of the one-sided null that none of the residuals is significantly positive.) MAX might miss activity of a strategic insider who acts like a Kyle-type monopolist. SUM (like CAR) might miss intense bursts of activity by competitive insiders.

It is important to clarify that we are interested in the cross-sectional variation in our measures over our sample of events, not the time-series variation for each firm. If a given firm gets a MAX score of 5.0, for example, that certainly seems suspicious (for a standardized residual in a time-series regression). But our methodology does not require us to render a statistical verdict on each deal. Our goal is not to assess whether insider trading took place in any particular instances. Rather it is to analyze the variations in the likelihood of such trading across bids.

Fig. 1 shows the histogram of the MAX measure of abnormal stock returns computed using the simplest regression with just a constant term employed to explain returns. As a benchmark, Fig. 2 shows the histogram of the same measure computed from the five-day window three months before the bid. While the overall frequency distributions look similar, the histogram for the five-day window immediately before the bid shows a significantly fatter right tail. There are nine outcomes greater than five standard deviations from the mean in the first window, whereas there is only one such outcome in the second one. This provides some evidence that the cross-sectional variation in our measures during the five-day window prior to bid announcements appears to be picking up activity related to the bid, and differs markedly from that observed in "normal" periods.

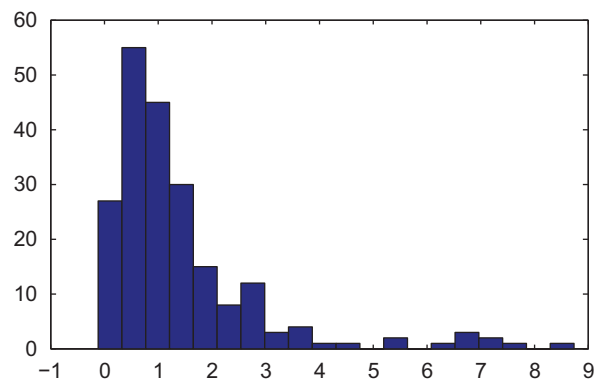**Fig. 1.** The figure shows the histogram of the MAX abnormal stock return measure computed using the pre-bid window from day −5 to −1, where abnormal returns are calculated as residuals in a regression of returns on a constant in the three-month pre-bid window.
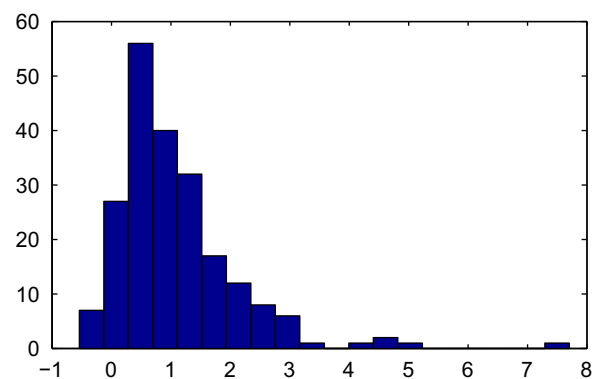

Fig. 2. The figure shows the histogram of the MAX abnormal stock return measure computed over the five-day window that starts three months before the bid.

4.1.2. Options market measures

Of the 212 target firms in our sample, 84 had traded options. Options can offer an insider a cheap way to leverage private information, and options trading has featured prominently in several enforcement cases brought by the SEC. We build measures of unusual options market activity in a similar fashion to our stock market volume measures.

Options present some unique issues with aggregation. A given target company with listed options will typically have dozens of available contracts to trade on any given day (i.e., puts or calls each with several maturities and strike prices). An insider could, in principle, profit from trading in any one of these. We want a single statistic to capture activity across all of them. This adds an additional layer to our procedure.

Our information for this market comes from the OptionMetrics database, and includes daily transactions volume and end-of-day prices for all U.S. listed options. We focus on volume measures and define the following two aggregate statistics: the total number of calls traded and the delta-weighted sum of all traded calls.

We use call volume because it is more efficient to speculate on the upside of a stock with calls than with puts. (We have built similar measures using call and put volume with very similar results; these are omitted for brevity.) The first statistic is self-explanatory. The second statistic weighs calls by delta, $\delta \equiv \partial C/\partial S$. This is a measure of the effective number of shares of stock exposure each commands. It is thus directly comparable to stock volume. Option deltas are computed by OptionMetrics using end-of-day pricing and implied volatilities based on a binomial model that accounts for the American feature of the options. We employ a third statistic that weighs each call by the sensitivity of its returns to the returns of the underlying stock. This number, given by $S\delta/C$, is more sensitive to the options that one might expect speculators to prefer, that is, those with the most "bang for the buck," but is highly sensitive to whether the strike price of the option would make the option in-the-money or out-of-the-money after the bid announcement.

Having computed each of these for all days, we then fit regression specifications to describe the expected value of each. The independent variables in these specifications are similar to those used for the two regressions employed for stocks: the first specification has just a constant and the second specification has a constant, contemporaneous market volume and returns, lagged volume and returns of the underlying stock, and a lagged dependent variable.

An additional complication in these regressions is the presence of a substantial number of zero-volume observations for some firms, that is, days for which no options traded of any strike or expiration. The presence of these days makes the data highly non-normal, and leads to potentially serious misspecification problems with ordinary least squares. To deal with this, we estimate a Heckman (1978) two-stage selection model, which fits the probability of any trade as a function of the regressors and then estimates the volume contingent on trade for the positive trade days. This procedure yields appropriate residuals and residual standard errors for zero and non-zero observations. As with the stock data, we then apply the metrics MAX and SUM to these residuals.

4.1.3. Credit derivatives measures

Of our target firms, 22 had credit default swaps that traded at the time of, and during the three months prior to, the bid. Data for CDS markets is problematic in that no actual transaction records exist. However, several vendors compile indicative end-of-day quotes from market makers, making it possible to compute daily changes in quoted fees. We regress these changes (in logs) on the following two sets of explanatory variables: a constant; and a constant, lagged dependent variable, lagged returns and volume of the firm's stock, day-of-the-week dummies, contemporaneous stock returns and volume, and the contemporaneous change in the BAA-AAA yield spread.

As with the other markets, we then construct the metrics MAX and SUM to yield our measures of unusual activity in the pre-bid window.

4.1.4. Bond market measures

While default swaps might be especially well suited to informed hedging of LBO risk by exposed creditors, the same logic applies to bonds issued by the target firms. We examine unusual activity in primary debt markets as well, which could differ from that in CDS markets for two reasons. First, coverage is different. Not all firms with corporate bonds have active CDS markets. Second, the regulatory regime can differ substantially in that trading in primary debt securities is more clearly subject to U.S. insider trading regulations, and market activity is directly monitored by the SEC.

Beginning in July 2002, corporate-bond trade data are available from the TRACE database. Some volume information is available on TRACE, but it is problematic in that reporting requirements were not uniform over the sample period, trades with non-U.S. entities and in private debt issues (144a) are not covered, and reported transaction sizes are truncated, with different ceilings for differently rated bonds. For the firms in our sample, we have transaction information for at least one debt issue during the three-month pre-bid period for 34 targets. Since some firms have several issues, we form these into portfolios weighted by issue size and study their daily returns. We again employ two different specifications of expected returns for each firm's bond portfolio employing as explanatory variables: the first has just a constant, and the second has a constant, a lagged dependent variable, lagged returns and volume of the firm's stock, day-of-week-dummies, contemporaneous stock returns and volume, and the contemporaneous change in the BAA-AAA yield spread.

Note that, unlike our measures in other markets, here we are interested in suspiciously *negative* residuals (corresponding to increases in yields or CDS fees). To facilitate interpretation of the results, we multiply the bond residuals by minus one, so that the signs of explanatory variables have the same meaning across markets. We then apply the MAX and SUM metrics to the negative residuals.

4.1.5. Related methodologies

Measures of suspicious trading prior to takeover bids and other news events have been reported in a number of

papers. Unusual pre-announcement stock trading activity was first documented by Mandelker (1974) and Keown and Pinkerton (1981). Keown and Pinkerton used daily returns and weekly volume, whereas Mandelker only used monthly returns. Similar results for options appear in Jayaraman, Frye, and Sabherwal (2001) and Arnold, Erwin, Nail, and Bos (2000). Recently, Gao and Oler (2004) and Cao, Chen, and Griffin (2005) have constructed measures of buyer-initiated and seller-initiated volume (in stock and options, respectively) prior to takeover announcements, using the former to identify presumably informed orders. Poteshman (2006) compiles distributional information for several summary statistics of options market activity and uses these to address the unusualness of trading in airline stock options prior to September 11, 2001. Berndt and Ostrovnaya (2007) use changes in option-implied volatility, as well as CDS fees and stock returns, to analyze the pre-event flow of information in a sample that overlaps with our own.

Relative to this literature, our paper does not claim to offer sharper evidence of insider activity and does not address the information flow across different markets. Rather, our focus is on understanding *why* insider trading occurs or when it becomes more likely. We are not aware of other work that examines cross-sectional determinants of informed trade in an event-study context.

4.2. Measuring the number of insiders

The primary tests in the paper are regressions of the measures of unusual pre-bid market activity on the characteristics of the takeover bid. In particular, we want to assess the role of the number of entities involved in financing the bid. To do this, we form separate measures of participation on the debt and equity sides.

For the equity side, our main information comes from the event descriptions provided by Thomson Financial. These descriptions list the major participants in each bid, which we simply count. There is certainly some degree of irregularity in this process as the database does not purport to provide an exhaustive list of participants for each deal. Nor is it even clear that it follows a consistent procedure for deciding which entities to list. Typically, LBOs involve equity stakes being taken by key officers and managers of the target entity, meaning that, technically, a large number of individuals are among the providers of equity finance. The data set appears to list individuals only in rare cases, presumably where they were key instigators or took very large stakes. We have cross-checked the counts we obtained with those obtained from another data provider for a subsample of our events and found good agreement. As mentioned in the Introduction, we do not interpret our counts literally, but only as monotonic (not necessarily linear) transformation of the true number of informed insiders. Of course, the usual argument about noisy data applies here: to the extent that our count is corrupted by random error, it is less likely that our regressions will uncover any relation with suspicious trade.

For debt finance, we follow Acharya and Johnson (2007) in counting the number of participants in syndicated loans

to the target company at the time of the deal. This definition is appropriate when, as in most LBOs, the target company is itself the borrowing entity for the debt used in the deal. It is not appropriate, for example, when the target is merged into an acquiror that itself assumes the additional debt. In our event sample, we were unable to identify instances of this. To be more accurate, we were unable to identify syndicated loans to any of the acquiring entities (e.g., Blackstone or KKR) that could be identified as having been used to finance particular bids. On the other hand, many of our targets did, in fact, take on additional debt following successful bids.

Data on syndicated borrowing come from the Loan Pricing Corporation's DealScan database. It provides lists of all participating entities, and identifies in particular those with lead-bank roles. We count these banks in two ways, reflecting different possibilities for which banks might have been informed prior to a bid.

The most narrow measure is restricted to the set of syndicated loans entered into within the six months after the bid event, and includes only the lead banks for these loans. These lead banks, which are providing takeover finance, would almost certainly have provided the bidders with prior commitments, and hence would have known a great deal about the bids. Not all our deals include records of loans specifically taken out to finance the buyout. (This could be because the borrowing entity had a separate name in Dealscan, or because the financing was not completed as of the time of this writing.) A second measure counts all lead banks in facilities active on or after the date of the bid. This adds to the previous count the target's main banks having ongoing relationships at the time of the bid. Whether or not they ultimately provided deal finance, these banks are likely to have been approached as potential lenders.

A third method of counting includes all bank participants in facilities that were active on or within six months after the date of the bid. This count includes non-lead banks, with which lead banks might have been obligated to share material non-public information in advance of the bid. Non-lead participants can include hedge funds and other investment firms that could have weaker incentives to abide by confidentiality agreements. Ivashina and Sun (2007) report indirect evidence of such leakage in that equity portfolios of institutions that hold stock and loans of the same company significantly outperform comparable investors that do not invest in the loan market. Even the last count of bank relationships is clearly only a lower bound, since it ignores all non-syndicated loans and commitments. In addition, a number of our target firms had no information on bank loan relationships. One could interpret this occurrence as firms having no banking relationships. However, we instead simply exclude these bid events from our tests involving bank counts.

Table 3 shows the distribution of our tabulation of providers of debt and equity finance for the sample bids. (The table uses the second definition of debt participants, that is, lead banks for loans outstanding at the time of the bid or within six months thereafter.) Nearly half of our deals had only a single buyer, whereas among those having syndicated loans, the median number of banks is 5.

Table 3
Target characteristics by number of participants.

	Equity participants				Debt participants		
	1	2	> 2	> 1	< 4	4–7	> 7
Obs	79	64	35	99	62	61	55
Market cap	351	402	1,122	515	230	319	1,605
Premium (pct)	24.2	28.6	21.6	26.0	26.2	27.1	21.6
M/B	1.65	1.63	1.74	1.64	1.49	1.63	1.79
D/V	0.29	0.31	0.24	0.30	0.28	0.27	0.33
β	0.81	0.81	0.67	0.78	0.69	0.78	0.82
σ	0.41	0.49	0.36	0.43	0.47	0.46	0.34
ILLIQ	0.009	0.012	0.002	0.006	0.039	0.014	0.001
Turnover	0.61%	0.59%	0.58%	0.58%	0.48%	0.60%	0.74%

The table shows the median characteristics of target companies selected according to the number of debt and equity participants in the buyout bid. The first four columns sort targets according to equity participants; the last four use the number of lead banks. The premium, reported by SDC, is the bid premium in excess of the stock price one week prior to announcement. Other measures are based the six-month (calendar) pre-announcement period. They are calculated per firm per day, and then averaged across the six-month period for each firm. Illiquidity, stock turnover, firm size, market beta and volatility are based on items reported in CRSP. Market to book value and leverage are based on items reported in Compustat. Market cap is the equity market value in millions of dollars. M/B is equity value divided by the book value of common equity. D/V is the book value of long-term debt divided by the sum of this value and the market value of equity. σ is the annualized volatility of daily stock returns. β is estimated with respect to the CRSP value-weighted index. Stock volume is in millions of shares per day. ILLIQ is measured using the Amihud (2002) ratio, computed daily and averaged over the entire period (in units of 10^{-6}). Turnover is in percent per day.

The table also reports characteristics of the target firms broken down by the number of participants in the bid. This allows us to assess the degree to which the number of participants is exogenous to the takeover process. As expected, the size of the target company is a key determinant of the number of participants. Companies with more than two equity participants are, on average, over twice as large as those with exactly two. Companies with more than seven lead banks are roughly seven times as large as those with fewer than four. Despite this strong size dependence, there is little noteworthy variation in target balance sheet characteristics or in the bid premium. Likewise, there is perhaps surprisingly little variation in the stock market risk measures: stocks with more participants (of either type) have somewhat lower volatility. Stocks with fewer lead banks are more illiquid (using the measure of Amihud, 2002).

While it is clear that deal size largely determines the number of providers of finance, the pairwise correlation between all lead banks and equity participants is only 0.22. Equity participants and deal size have a correlation of 0.36, whereas banks and size have a correlation of 0.51. These correlations confirm that deal size is not the only determinant of syndicate sizes, and that equity and debt syndicate sizes are driven by different considerations.

The strong correlation between the number of bid participants and the target size has significant implications for our empirical work, however. A common assumption in the empirical literature is that asymmetric information problems are worse for smaller firms. If so, then controlling for size should be important in order to

isolate the independent effect of the number of insiders. On the other hand, size can also be viewed as an additional measure of the number of insiders. Larger companies have more officers and directors and more investment bankers and lawyers. We have no other measure of these. Moreover, our measures of the size of the bidding syndicates are fairly crude and could also tend to undercount informed players for larger companies. For example, larger targets might be more likely to have competing bidders, each involving its own banks and advisors. We count only entities participating in a single bid. This line of reasoning suggests that target market capitalization (or deal value) should actually enter positively in our regressions.

5. Results

We now present our primary regression results. The null hypothesis is that (due to efficient enforcement of regulation) the number of participants in a financing syndicate is not related to the degree of suspicious pre-bid trading activity. The alternative of interest is that there is a positive relation with one or more of our measures of syndicate size (including the capitalization of the target firm). We present our basic results for the equity (and equity-linked) and debt (and CDS) markets separately. We then consider some alternative specifications and robustness checks. A final subsection considers interpretation of our findings.

5.1. Equity and options markets

The top panel of Table 4 shows our basic tests using stock market data. The primary finding is that the number of equity participants in a deal is significantly positively associated with the degree of suspicious stock market activity in every regression shown. In contrast, there is little association between number of the target's lead banks (the debt participants measure used in the table) and unusual activity, although the coefficient is always positive. This finding hints at the possibility, to which we return below, that different types of insiders exploit distinct markets. We find a significant positive association between target size and pre-bid volume (although not returns). As mentioned in Section 3, a positive effect of firm size is consistent with size as a proxy for uncounted insiders. There is also no distinct pattern separating the results for the two metrics MAX and SUM. This initial evidence is thus in favor of the view that a larger number of insiders leads to more insider trading.

The bottom panel of Table 4 presents our results for unusual option activity. The table repeats the tests used in the stock market but with slightly different regression specifications, as described in Section 3. In addition, we report results for the two methods of volume aggregation — raw volume and delta-weighted volume.

The overall conclusion from this table is that there is some evidence that the number of equity participants is linked to suspicious activity in options markets. There is no significant effect from the number of debt participants

Table 4
Equity market regressions.

Panel I: Stock market								
	Returns				Volume			
	MAX		SUM		MAX		SUM	
	Uncond	Cond	Uncond	Cond	Uncond	Cond	Uncond	Cond
Equity participants	0.3074	0.3423	0.2809	0.3067	0.4355	0.3980	0.4363	0.4177
	(2.83)	(3.23)	(1.96)	(2.16)	(3.03)	(2.91)	(2.35)	(2.32)
Debt participants	0.0450	0.0373	0.0614	0.0673	0.0089	0.0128	0.0235	0.0359
	(1.48)	(1.26)	(1.53)	(1.69)	(0.22)	(0.34)	(0.45)	(0.71)
Target size	−0.0236	−0.0102	−0.0329	−0.0445	0.2893	0.2758	0.4670	0.4579
	(0.21)	(0.09)	(0.22)	(0.30)	(1.92)	(1.96)	(2.44)	(2.46)

Panel II: Options market								
	Raw volume				Delta weighted			
	MAX		SUM		MAX		SUM	
	Uncond	Cond	Uncond	Cond	Uncond	Cond	Uncond	Cond
Equity participants	0.3323	0.3985	0.5376	0.3378	0.4285	0.4470	0.6402	0.3739
	(1.78)	(2.24)	(1.60)	(1.27)	(2.32)	(2.41)	(1.90)	(1.40)
Debt participants	−0.0137	−0.0049	−0.0135	0.0051	0.0042	−0.0042	−0.0123	−0.0093
	(0.26)	(0.10)	(0.14)	(0.07)	(0.08)	(0.11)	(0.13)	(0.13)
Target size	0.1787	−0.421	0.6731	0.1701	0.0351	−0.0983	0.3714	0.0473
	(0.71)	(0.18)	(1.49)	(0.48)	(0.14)	(0.40)	(0.82)	(0.13)

The table presents regression results for different measures of unusual pre-bid stock and options market activity on bid characteristics. The equity participants variable is the number of distinct bidding entities listed by Thomson Financial. Debt participants is the number of lead banks for syndicated loans to the target at the date of, or within six months after, the bid. Target size is the market value of the target at the bid price. The unusual trading measures are defined by a regression specification of expected returns or volume in three months of daily data prior to the bid; a metric is applied to the standardized residuals from these regressions in a five-day pre-bid window to detect unusually large values. Columns labeled "uncond" use residuals from a regression of returns or volume on a constant. Columns labeled cond use residuals from a regression on a vector of lagged predictors. The residual metrics are labeled MAX and SUM. (See Section 3 for descriptions of each.) The top panel employs stock market data from 178 bid events. The dependent variables in the four left columns are the suspicious activity measures constructed from returns. In the four right columns, the dependent variables are the suspicious activity measures constructed from volume. The bottom panel employs equity-options market data from 83 bid-events. The dependent variables in the four left columns are the suspicious activity measures constructed from raw call volume. In the four right columns, the dependent variables are the suspicious activity measures constructed from delta-weighted call volume. OLS t statistics are shown in parentheses.

or from target size. The effect of the number of equity participants is stronger using the MAX metric than the SUM metric. Recall that the latter was included to account for the possible effect of stealth trading (or order splitting) by insiders, which could result in positive unexplained volume on each of several days but with no one day looking unusual. Our finding here is not that such strategies do not take place but that the number of insiders does not appear to be a better predictor of its occurrence than of unusual single days.

5.2. CDS and bond markets

Our primary results for credit derivatives are shown in the top panel of Table 5. The table presents the basic regressions of unusual (log) changes in CDS fees, and uses the first two of the three different methods of counting banks explained in Section 4.2.[9]

---
[9] Results for the third measure (all syndicate banks, pre-LBO financing as well as LBO financing) are weaker and available upon request. Intuitively, we expect lead banks to be more important for insider trading tests.

The first method, which counts all lead banks in all active syndicated loans to the targets, shows no significant association with unusual pre-bid CDS activity using the MAX metric or the SUM metric. However, when we count only lead banks for facilities activated *after* the bid (hoping to measure the banks that participated in the LBO financing), the association is positive and statistically quite significant.

Our sample size is quite small, and we cannot be confident that *only* lead banks financing LBOs are associated with information leakage. But such could be the case if, in general, other lead banks of the target are not approached by the bidder as potential sources of funds and do not learn of the impending bid via information-sharing within existing syndicates.

The table also provides the complementary result to our earlier finding that equity participants alone matter for stock and options market trading. Here, debt participants in the LBO financing alone seem to matter for credit market trading. This could be evidence of a preferred habitat — or comparative advantage — for traders of different types. Banks might be better positioned to move quickly in credit derivatives. The finding also implies that cross-market arbitrage is less than perfect, as discussed later.

Table 5
Credit market regressions.

Panel I: CDS market								
	All lead banks				LBO lead banks			
	MAX		SUM		MAX		SUM	
	Uncond	Cond	Uncond	Cond	Uncond	Cond	Uncond	Cond
Equity	0.3457	0.3153	0.3907	0.3392	−0.0260	−0.0191	−0.1714	−0.1311
participants	(0.77)	(0.79)	(0.65)	(0.68)	(0.08)	(0.06)	(0.42)	(0.39)
Debt	0.0826	0.0708	0.1556	0.1210	0.5251	0.4712	0.8035	0.6691
participants	(0.77)	(0.74)	(1.09)	(1.01)	(4.31)	(4.33)	(5.57)	(5.56)
Target	0.6271	0.5046	0.8789	0.6720	0.5824	0.4583	0.8646	0.6442
size	(1.12)	(1.00)	(1.17)	(1.07)	(1.49)	(1.32)	(1.87)	(1.67)

Panel II: Bond market								
	All lead banks				LBO lead banks			
	MAX		SUM		MAX		SUM	
	Uncond	Cond	Uncond	Cond	Uncond	Cond	Uncond	Cond
Equity	0.3027	0.2419	0.1932	0.1459	0.1838	0.1544	−0.0471	−0.0415
participants	(2.04)	(1.65)	(0.90)	(0.68)	(1.20)	(0.98)	(0.22)	(0.18)
Debt	0.0744	0.0800	0.1444	0.1468	0.1445	0.1114	0.2907	0.2336
participants	(1.78)	(1.94)	(2.39)	(2.44)	(2.40)	(1.79)	(3.48)	(2.64)
Target	0.1251	−0.1184	0.3787	−0.1193	0.2871	0.0674	0.6907	0.2142
size	(0.63)	(0.61)	(1.33)	(0.42)	(1.82)	(0.42)	(3.16)	(0.93)

The table presents regression results for different measures of unusual pre-bid CDS and corporate bond market activity on bid characteristics. The equity participants variable is the number of distinct bidding entities listed by Thomson Financial. Debt participants is defined as follows. In the four left columns, it is the number of all lead banks taking part in syndicated loans to the target at the date of, or within six months after, the bid. In the four right columns, it is the number of lead banks for syndicated loans originated after the bid. Target size is the market value of the target at the bid price. The unusual trading measures are defined by a regression specification of expected returns or volume in three months of daily data prior to the bid; a metric is applied to the standardized residuals from these regressions in a five-day pre-bid window to detect unusually large values. Columns labeled "uncond" use residuals from a regression of returns or volume on a constant. Columns labeled "cond" use residuals from a regression on a vector of lagged predictors. The residual metrics are labeled MAX and SUM. (See Section 3 for descriptions of each.) The top panel employs CDS log changes from 22 bid-events. The bottom panel employs the returns on an issue-size weighted portfolio of corporate bonds (multiplied by minus one) from 34 bid events. OLS *t*-statistics are shown in parentheses.

We have a somewhat larger sample of targets with bond market information. The results using their measures of unusual activity are shown in the lower panel of Table 5. Based on the SUM metric, there is a convincing positive association with the number of debt participants in nearly every specification. The evidence for such an association is somewhat weaker using the MAX metric, which could be indicative of a slower reaction to informed trading in the bond market. In contrast with the CDS sample, the positive association between suspicious activity and the number of banks is pronounced for all lead banks as well as only the post-bid lead banks. This is consistent with evidence in Acharya and Johnson (2007) that having a large number of participant banks increases information leakage. This channel is widely believed to be due to increased hedge fund participation in (and purchase of) syndicated debt precisely in order to acquire non-public information.

The results leave open some questions with regard to the different types of debt syndicate measures and their effect on the likelihood of insider trading. For CDS markets, we found that insider trading is more likely with more lead syndicate members for LBO financing, but not with more lead syndicate members in the target's financing inclusive of pre-LBO financing. Although possibly just due to a very small sample size, this feature of the

data fits our understanding of deal timing based upon conversations with bankers involved in LBO deals. This communication suggests the following typical timeline of events. At date-zero, firm A contemplates making a bid for firm T. They approach potential lenders, usually including the target's existing banks, and get a commitment letter from one or more that will become lead banks if a deal ensues. At date-one, firm A actually makes a tender offer for firm T. The commitment letter becomes public (filed with the SEC). LPC news will likely have a story saying who the lead banks are. At date two, the deal is successful and firm A will certainly need the money. The lead banks then solicit other "senior manager" and "bookrunner" banks. Then the bigger group solicits general participants and institutions. At date three, the syndicate is finalized and the loan becomes "active" as per the LPC data. Thus, according to this timeline, and assuming no leakage of information to prospective participant banks, the relevant insiders are captured by the LBO syndicate leads.

5.3. Alternative tests

We perform several tests to assess the robustness of our results. The first set of tests allows for additional controls in our benchmark estimations.

We consider the following set of additional control variables: the bid premium (measured by the raw return of the stock price over the week prior to the announcement), the book-to-market ratio, leverage, the volatility of stock returns, the stock beta, a measure of illiquidity (the ILLIQ ratio of Amihud, 2002) and a measure of activity (stock turnover). We expect at least some of these to be related to the intensity of insider trading. The bid premium captures the gains to be made from trading on insider information and should increase the economic motives for insiders to exploit their private information. We do not have data on the anticipated leverage of the buyout target after being acquired. But the change in leverage would be one determinant of credit spread widening, and would thus be correlated with incentives to exploit private information in the CDS markets. High volatility of stock returns could deter insiders from trading aggressively since, if they have limited capital, they might be unable to diversify away risk related to their position. The systematic component of risk, measured by beta, can also capture this. On the other hand, in debt markets, where the incentive behind insider trading could be to *reduce* firm-specific risk, these measures might work in the opposite way. Finally, insiders can trade larger quantities for the same amount of private information if the underlying markets are more liquid and have greater depth

(smaller price impacts). Thus, we might expect insider trading volume to be higher for liquid markets.

Tables 6 and 7 show the results with additional controls for stock volume, option volume CDS changes, and bond returns, respectively. (The results for stock returns parallel those for stock volume and are omitted for brevity.) The first observation is that, by and large, the relation between suspicious trading and syndicate sizes is the same as in our benchmark results in terms of signs, magnitudes, and statistical significance for both the MAX and SUM measures: stock return and volume and options volume measures are related to equity syndicate size, and CDS and bond return measures are related to number of LBO lead banks. The second observation is that some of the hypotheses proposed above on the coefficients of additional control variables indeed find support. For example, MAX and SUM for stock returns are significantly positively related to the bid premium. Stock volatility leads to significantly lower suspicious trading for the stock market measures but significantly *higher* unusual returns for corporate bonds. Stock liquidity is associated with greater MAX and SUM in stock volume (especially if firm size is considered a proxy for liquidity and also to some extent with turnover as the proxy, although the sign on ILLIQ is the opposite of the predicted one).

Table 6
Equity market regressions with further controls.

	Stock volume				Options volume			
	MAX		SUM		MAX		SUM	
	Uncond	Cond	Uncond	Cond	Uncond	Cond	Uncond	Cond
Equity	0.4477	0.4271	0.6254	0.4347	0.4090	0.4281	0.6798	0.3351
participants	(3.08)	(3.07)	(2.65)	(2.39)	(2.11)	(2.18)	(1.94)	(1.20)
Debt	0.0005	0.0023	0.0303	0.0257	−0.0210	−0.0222	−0.0441	−0.0374
participants	(0.01)	(0.06)	(0.44)	(0.49)	(0.33)	(0.35)	(0.39)	(0.41)
Target	0.6290	0.4926	1.5042	0.8841	0.3958	0.2818	1.7109	0.7160
size	(2.14)	(1.76)	(3.15)	(2.40)	(0.65)	(0.45)	(1.55)	(0.81)
Bid	0.0105	0.0102	0.0172	0.0149	−0.0115	−0.0068	−0.0091	−0.0141
premium	(1.27)	(1.30)	(1.28)	(1.44)	(0.78)	(0.46)	(0.34)	(0.66)
Book/	0.0329	−0.1587	0.4546	−0.1191	0.9513	0.6910	1.7316	1.5295
market	(0.07)	(0.34)	(0.57)	(0.19)	(1.05)	(0.75)	(1.06)	(1.17)
Leverage	0.2033	0.2253	0.3756	0.2524	0.5973	0.4143	1.0429	0.6747
	(1.14)	(1.33)	(1.32)	(1.13)	(0.73)	(0.50)	(0.70)	(0.57)
σ	−2.1030	−2.0202	−3.3534	−2.7070	−1.0117	−0.3032	−5.5573	−2.2833
	(2.15)	(2.17)	(2.11)	(2.21)	(0.50)	(0.15)	(1.53)	(0.79)
β	−0.0346	−0.0037	−0.0693	−0.0216	−0.1599	−0.1865	−0.0075	−0.2784
	(0.12)	(0.01)	(0.14)	(0.06)	(0.38)	(0.43)	(0.01)	(0.45)
ILLIQ	0.0119	0.0089	0.0264	0.0149	0.0195	0.0191	0.0888	0.0398
	(1.57)	(1.23)	(2.14)	(1.57)	(0.55)	(0.53)	(1.38)	(0.77)
Turnover	0.2118	0.1895	0.4046	0.2453	0.1219	0.1094	0.7355	0.2446
	(1.38)	(1.29)	(1.62)	(1.27)	(0.37)	(0.32)	(1.22)	(0.51)
$F_{1,2}$	[0.005]	[0.006]	[0.016]	[0.040]	[0.147]	[0.130]	[0.195]	[0.516]
$F_{1,2,3}$	[0.000]	[0.001]	[0.000]	[0.000]	[0.201]	[0.203]	[0.106]	[0.542]

The table presents regression results for different measures of unusual pre-bid volume in equity and equity options markets on bid characteristics. The dependent variables are as described in the caption to Table 4. There are 177 bid events in the equity sample and 83 in the options sample. The options results use delta-weighted call volume. The independent variables are defined as follows. Bid premium is the bid premium in percent over the stock price one week prior to announcement. Book/market is the book value of target equity divided by bid value. Leverage is the target enterprise value divided by the bid value of equity. σ is the annualized volatility of the target in the three months prior to the bid. β is its beta with respect to the CRSP value-weighted index in the same period. ILLIQ is the cross-sectional rank of the illiquidity measure of Amihud (2002). Turnover is the annualized volume in shares divided by shares outstanding. OLS t statistics are shown in parentheses. The table also presents results (p values) for the F test that the two participant coefficients are jointly zero ($F_{1,2}$) and for the test that all three coefficients are jointly zero ($F_{1,2,3}$).

Table 7
Credit market regressions with further controls.

	CDS changes				Bond returns			
	MAX		SUM		MAX		SUM	
	Uncond	Cond	Uncond	Cond	Uncond	Cond	Uncond	Cond
Equity	−0.0314	−0.0507	−0.1986	−0.1399	0.0023	−0.0092	−0.3505	−0.2317
participants	(0.09)	(0.18)	(0.48)	(0.42)	(0.02)	(0.07)	(1.98)	(1.22)
Debt	0.5356	0.4865	0.8010	0.6692	0.1996	0.1623	0.3771	0.3001
participants	(4.82)	(5.28)	(5.92)	(6.18)	(4.15)	(3.37)	(5.48)	(4.08)
Target	2.0147	2.0975	1.3132	1.4790	0.8388	0.2448	1.1086	−0.1460
size	(1.88)	(2.35)	(1.01)	(1.42)	(1.67)	(0.49)	(1.54)	(0.19)
Bid	−0.0099	−0.0034	0.0080	0.0161	0.0144	0.0366	0.0087	0.0397
premium	(0.25)	(0.10)	(0.16)	(0.41)	(1.07)	(2.72)	(0.46)	(1.93)
Book/	1.1872	1.4507	2.2029	1.8767	−1.8469	−0.9955	−2.7245	−0.9401
market	(0.71)	(1.04)	(1.08)	(1.15)	(2.56)	(1.38)	(2.64)	(0.85)
Leverage	−0.3720	−0.2823	−0.5478	−0.5622	−0.0554	0.3590	0.0081	0.9214
	(0.61)	(0.55)	(0.73)	(0.94)	(0.22)	(1.41)	(0.02)	(2.37)
σ	−7.6141	−7.9478	−7.9919	−7.2540	5.6964	4.5615	10.3336	6.0964
	(1.37)	(1.72)	(1.18)	(1.34)	(2.93)	(2.34)	(3.72)	(2.05)
β	1.1477	1.2714	2.0486	1.7537	−0.5689	−0.2498	−1.3981	−0.4198
	(0.95)	(1.26)	(1.39)	(1.48)	(1.42)	(0.62)	(2.45)	(0.69)
ILLIQ	0.3199	0.3503	0.1349	0.2090	0.0659	0.0051	0.0366	−0.0827
	(1.67)	(2.20)	(0.57)	(1.12)	(0.92)	(0.07)	(0.36)	(0.76)
Turnover	0.5724	0.6758	0.0522	0.1963	−0.2577	−0.3682	−0.4624	−0.5319
	(1.13)	(1.61)	(0.09)	(0.40)	(1.11)	(1.59)	(1.40)	(1.50)
$F_{1,2}$	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.002]	[0.000]	[0.000]
$F_{1,2,3}$	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]

The table presents regression results for different measures of unusual pre-bid CDS changes and negative bond returns on bid characteristics. There are 22 bid events in the CDS sample and 34 in the bond sample. The dependent variables are as described in the caption to Table 5. The debt participants variable is the number of lead banks for syndicated loans originated after the bid. The control variables are as described in the caption to Table 6 OLS *t* statistics are shown in parentheses. The table also presents results (*p* values) for the *F* test that the two participant coefficients are jointly zero ($F_{1,2}$) and for the test that all three coefficients are jointly zero ($F_{1,2,3}$).

On the one hand, these results illustrate that the relationship between suspicious trading activity and syndicate sizes is robust to additional controls. On the other hand, the fact that some of the additional controls (bid premium, volatility, and liquidity, in particular) are themselves related to trading activity in a manner that is consistent with informed speculation or hedging lends additional confidence that our measures of suspicious trading are indeed representative of trading by informed agents.

The second test is to address the concern that our bid announcement dates gathered from Thomson Financial are not in fact perfectly accurate. If the dates are late by a few days, then we might observe large trading returns and volumes prior to the recorded announcement date. By going through public records of announcements, we have verified for many instances that this is not the case. We have also checked that in over 70% of our deals, the maximum standardized residual in stock returns occurs on day 0 (around 50%) or day −1 (an additional 20%). To account for the possibility that perhaps the bid information reaches markets a day before the actual announcement date in our data, we examine the window preceding day −1 to calculate the MAX and SUM measures. Simultaneously, we consider a ten-day window before day −1 to account for the fact that in at least some (alleged) insider trading cases discussed in media, the abnormal trading patterns were claimed to have been detected as early as two weeks prior to the bid announcement.

The results from employing this alternate pre-bid window are contained in Table 8. Overall, the link between suspicious trading and syndicate sizes is robust to this change, although the significance is weakened for the bond returns and for the SUM metric in stock return, stock volume, and option volume measures.

A final issue to consider is whether our cross-sectional result that suspicious trading indicators are linked to syndicate sizes is in fact a time-series result. We know from trends in the LBO markets that syndicate sizes have grown secularly over time. It is also plausible that the intensity of analyst following and arbitrage activities, precisely aimed at identifying LBO targets, has increased and perhaps even got better over time. If this were true, then the increase in suspicious trading we identified would simply reflect the greater information acquisition prior to bid announcements. To address this hypothesis, we re-ran the estimations of Table 4 (stock returns and volume regressions) with year dummies. While we do not report the entire estimations, several points are noteworthy. First, the coefficients in the main regressions are unaffected, in size and significance. Second, the year intercepts are neither significant nor monotonically increasing. For instance, the largest year intercept for stock returns is 2001 and for stock volume 2002. Hence, it does not seem that our results are driven by an increase

Table 8
Alternate pre-bid window.

| | Stock volume | | | | Options volume | | | |
| | MAX | | SUM | | MAX | | SUM | |
	Uncond	Cond	Uncond	Cond	Uncond	Cond	Uncond	Cond
Equity	0.3545	0.3319	0.2638	0.2263	0.5044	0.4847	0.5977	0.4179
participants	(2.40)	(2.40)	(0.90)	(0.97)	(2.96)	(2.92)	(1.88)	(1.58)
Debt	−0.0479	−0.0561	−0.0337	−0.0271	−0.0372	−0.0350	−0.1063	−0.0771
participants	(1.16)	(1.45)	(0.41)	(0.41)	(0.76)	(0.73)	(1.16)	(1.01)
Target	0.3183	0.3334	0.8033	0.6092	0.0779	0.0522	0.3693	−0.0004
size	(2.09)	(2.34)	(2.65)	(2.53)	(0.33)	(0.23)	(0.84)	(0.00)

| | CDS changes | | | | Bond returns | | | |
| | MAX | | SUM | | MAX | | SUM | |
	Uncond	Cond	Uncond	Cond	Uncond	Cond	Uncond	Cond
Equity	0.3793	0.2746	0.6916	0.6727	0.1831	0.1622	0.0770	0.1194
participants	(1.50)	(1.24)	(1.72)	(1.98)	(1.11)	(1.08)	(0.32)	(0.50)
Debt	0.3019	0.2694	0.5690	0.4780	0.0700	0.0620	0.1740	0.1280
participants	(3.35)	(3.41)	(3.96)	(3.95)	(1.08)	(1.05)	(1.82)	(1.35)
Target	0.2379	0.2041	0.2104	0.0299	0.2685	0.0351	0.4815	−0.0125
size	(0.82)	(0.80)	(0.46)	(0.08)	(1.59)	(0.23)	(1.93)	(0.05)

The table presents regression results corresponding to Tables 4 and 5 but with the MAX and SUM measures computed over the 10-day pre-bid window from date −11 through date −2. Note that results are reported only for delta-weighted call option volume. For CDS and bond returns, the debt participants variable is measured by the number of LBO lead banks. OLS *t* statistics are shown in parentheses. The number of observations for stock, options, CDS, and bonds is 178 , 83, 22, and 34, respectively.

over time in the extent of information generated about the likelihood of LBO deals. It is difficult to verify this for options and credit markets since we do not have observations in all years. In particular, virtually all the events in the CDS subsample are in 2006.

5.4. Market segmentation or cross-market transmission?

It is important to ascertain that our results for different markets — stock, options, bonds and CDS — are not manifestations of just the underlying result that there is evidence consistent with insider trading in *some* market and that cross-market information flows give rise to related trading activity in other markets.

Table 9 quantifies the extent to which our suspicious activity measures are picking up distinct signals across different markets. For each pair of market indicators, the correlation between suspicious trading measures is computed across targets for which we have information in the two markets. So, for example, the upper rightmost entry indicates that the cross-bid correlation between suspicious stock market returns (using the SUM metric) and suspicious bond market returns (using the MAX metric) is 0.5469.

The table tells us that there is significant correlation across all pairs of markets in unusual activity. This is not surprising. Arbitrageurs ensure that information about a firm revealed in one market is rapidly transmitted across all similarly exposed instruments. When we seek to *explain* the likelihood of suspicious activity, this transmission also implies that anything that helps predict it in one

market will, mechanically, have some predictive power in other markets. This raises the possibility that, for example, true insider activity happens only in the credit derivatives markets and is driven by the number of banks with exposure to the target's debt. Finding a role for the number of creditors in insider trading in another market could then be spurious.

Table 10 addresses the question of whether cross-market transmission of information can account for *all* of the significant role played by number of financing relationships. The table repeats the basic regressions of Tables 4 and 5 using subsamples of bid events for which transmissions between some pairs of markets cannot possibly be a factor. For example, the second and fifth columns repeat the regressions using the stock market indicators but *only* using targets for which there were *no traded options* at the time of the bid. Notice that for this subsample there is no longer a statistically significant positive role for the number of equity participants. While this is partly due to a smaller sample (comparing the fourth and fifth columns, the magnitude of the coefficient is almost the same), it also is consistent with the view that our stock market evidence is simply echoing a significant role for equity participants that exists in the options markets.

The conclusion of the table is quite different, however, when we confine the sample to targets for which there was not a CDS market at the time of the bid. The third and sixth columns show that the role for equity participants is still significant and in some cases stronger than in the full sample. Thus, we can rule out the hypothesis that our equity market results are driven by cross-market

Table 9
Correlations of unusual activity across markets.

		Stock return		Stock volume		Options volume		CDS return		Bond return	
		SUM	MAX	SUM	MAX	SUM	MAX	SUM	MAX	SUM	MAX
Stock return	SUM	–	0.9286 (211)	0.5540 (211)	0.5752 (211)	0.5887 (82)	0.6181 (82)	0.6511 (23)	0.6865 (23)	0.6114 (39)	0.5469 (39)
	MAX		–	0.6074 (211)	0.6689 (211)	0.5113 (82)	0.6127 (82)	0.6440 (23)	0.7103 (23)	0.5735 (39)	0.5743 (39)
Stock volume	SUM			–	0.9420 (211)	0.6187 (82)	0.6014 (82)	0.7138 (23)	0.7226 (23)	0.5320 (39)	0.4180 (39)
	MAX				–	0.5752 (82)	0.6317 (82)	0.7186 (23)	0.7775 (23)	0.5434 (39)	0.4751 (39)
Options volume	SUM					–	0.9302 (82)	0.7041 (19)	0.5988 (19)	0.5493 (26)	0.3900 (26)
	MAX						–	0.7178 (19)	0.6633 (19)	0.5843 (26)	0.4935 (26)
CDS return	SUM							–	0.9482 (23)	0.5124 (19)	0.3594 (19)
	MAX								–	0.3884 (19)	0.2832 (19)
Bond return	SUM									–	0.8915 (39)
	MAX										–

The table shows the sample correlation between our measures of unusual activity in different markets constructed using our conditional specifications (A2,B2,C2,D2,E2) during the five-day pre-bid windows. Options volume uses delta-weighted call volume with Heckman's adjustment. Bond returns are for portfolios. The number of bids events in each pair-wise comparison is shown in parentheses.

Table 10
The effect of insiders in subsamples.

	Dependent: unusual stock returns			Panel I: MAX metric Dependent: unusual stock volume			Dependent: unusual options volume	
	All deals	Excl. targets w/ options	Excl. targets w/ CDS	All deals	Excl. targets w/ options	Excl. targets w/ CDS	All deals	Excl. targets w/ CDS
Equity participants	0.3423 (3.23)	0.2282 (1.46)	0.3760 (3.40)	0.3980 (2.91)	0.3795 (1.66)	0.3710 (2.59)	0.4470 (2.41)	0.5391 (2.67)
Debt participants	0.0373 (1.26)	0.0067 (0.16)	0.0038 (0.11)	0.0128 (0.34)	−0.0722 (1.12)	−0.0234 (0.55)	−0.0042 (0.08)	0.0248 (0.39)
Target size	−0.0102 (0.09)	−0.0794 (0.49)	−0.1039 (0.85)	0.2758 (1.96)	0.6209 (2.58)	0.1992 (1.26)	−0.0983 (0.40)	−0.4949 (1.50)
N obs.	178	97	156	178	97	156	83	63

	Dependent: unusual stock returns			Panel I: SUM metric Dependent: unusual stock volume			Dependent: unusual options volume	
	All deals	Excl. targets w/ options	Excl. targets w/ CDS	All deals	Excl. targets w/ options	Excl. targets w/ CDS	All deals	Excl. targets w/ CDS
Equity participants	0.3067 (2.16)	0.2393 (1.05)	0.3689 (2.40)	0.4177 (2.32)	0.3811 (1.27)	0.3807 (1.96)	0.3739 (1.40)	0.5275 (1.87)
Debt participants	0.0673 (1.69)	0.0039 (0.06)	−0.0002 (0.00)	0.0359 (0.71)	−0.0801 (0.94)	−0.0032 (0.06)	−0.0093 (0.13)	0.0156 (0.17)
Target size	−0.0445 (0.30)	−0.0559 (0.23)	−0.1203 (0.71)	0.4579 (2.46)	0.8203 (2.58)	0.28242 (1.32)	−0.0473 (0.13)	−0.4112 (0.89)
N obs.	178	97	156	178	97	156	83	63

The table performs the regressions in Table 4 on subsamples of bid events for which the target did *not* have securities traded in one or more of the other markets. The table employs measures of unusual activity constructed using our conditional specifications (A2,B2,C2,D2,E2) during a five-day pre-bid windows.

transmission of information from insiders in the CDS market. The final column shows that the same conclusion holds for our options market finding when we exclude targets with credit derivatives.

This finding also supports our view that the relation we document for the equity and equity derivatives markets is distinct from the relation documented by Acharya and Johnson (2007) in the CDS markets. While we are unable to perform the converse tests for credit markets (we have no targets for which there is CDS or bond information but not stock information, and only three for which there are CDS instruments but not

options), we suspect that they would confirm this view because there is no explanatory power in equity markets for the number of debt participants. It is thus unlikely that debt participants' trading in the stock market is leaking to the bond or CDS market and inducing the significant role we find there.

Our bond market result contrasts, however, with that of the CDS markets and reveals that it is not just LBO syndicate size that makes insider trading more likely but also the overall debt syndicate size inclusive of pre-LBO leads. This is more in line with the result of Acharya and Johnson (2007) that information release in CDS markets prior to credit deterioration events increases with the total number of relationship banks of the underlying entity. The bond market result thus suggests that information indeed leaks from LBO syndicate leads to existing bankers of the target and/or to potential participants who will be arranged in case the deal goes through.

5.5. Interpretation of results

Our primary finding from the above results is that a larger number of insiders indeed makes suspicious trading more likely. This is contrary to the predictions in standard informed trading models wherein trading becomes more revealing with more competitors. The standard models would predict that the effect of two informed traders is much the same as that of many. Hence, one explanation for our primary finding is that competition might be limited by the wealth constraints of insiders and/or exposure to timing or market risks — realistic features that are generally not considered in the microstructure literature. However, this hypothesis also raises the possibility that it is not the *extent* of insider activity that has risen in the period under study but its detectability. Insiders might have tended to trade more aggressively, making it easier for the market to deduce their presence and resulting in quicker price reactions. Without detailed trade and trading counterparty information, it is difficult to rule out this alternative.

The second explanation for our finding has to do with the nature of the enforcement regime. If each potential insider regards the likelihood of detection (and the probable penalty upon detection) as independent of the number N of insiders, then one would expect a rising number of informed players to result in a rising amount of illegal behavior. We argued in our theoretical model that such an enforcement regime might be suboptimal because the harm to market liquidity from allowing more insiders to trade can be efficiently avoided by imposing an enforcement ceiling of abnormal trading activity above which detection and penalties are imposed. This ceiling should be optimally declining in N to enable insiders to internalize the cost they impose on other insiders with an additional unit of trade.

Hence, we conclude that a finding that more insiders leading to more insider trading is likely to be consistent with the existence of frictions faced by insiders due to wealth or risk considerations. The relation between

insiders and insider trading, however, is less likely to be consistent with the enforcement regime having optimally responded to the changing institutional structure of LBOs, particularly the larger syndicate sizes.

Beyond our model, however, it also seems clear that allowing the total amount of informed trading to rise with N creates dangerous incentives. To the extent that insiders can *choose N* — e.g., one can always tip off one's friends — there could be a positive net benefit to doing so. If expected individual punishment actually weakens with N, there arises an externality that makes it safer for more agents to trade together. Such a policy would entail a social dimension to insider trading under which "crime wave" equilibria become possible (see, e.g., Sah, 1991; Glaeser, Sacerdote, and Scheinkman, 1996; Bond and Hagerty, 2005).

Also beyond the scope of the model, there could be implicit penalties due to concerns for probity and honesty.[10] Such implicit penalties are likely to imply a constant likelihood of an informed party becoming an insider, or perhaps even a rising likelihood if there is some "sharing of the blame" as numbers increase. Optimal enforcement of insider trading might thus have to be stricter as insiders increase in number, either because implicit penalties do not cause insiders to internalize the negative externality of their trades on others or because they in fact exacerbate this externality.

An additional noteworthy result is that the effect of syndicate sizes seems to be segmented across equity and credit markets. Our evidence shows that the equity syndicate size makes insider trading more likely in equity and options markets, whereas the debt syndicate size matters for CDS and bond markets. This segmentation potentially tells us something about the nature of information flow from syndicate members to other players in financial markets. For instance, media discussions surrounding insider trading around LBOs seem to suggest that information sometimes leaks from hedge funds, which are often debt syndicate members, to their prime brokers, which are banks and perhaps more naturally inclined to trade in bonds and CDS to hedge their counterparty exposure to the hedge funds. Similarly, a lead arranger of the debt financing is likely to communicate with a bond underwriter to assess whether a certain amount of debt can be placed in the market for the LBO at a desired spread. Finally, equity syndicate members might consult equity analysts to get an idea of the overall sector prospects or an assessment of valuation of the target.

6. Conclusion

In this paper, we derive a theoretical result that the optimal regulation of insider trading should get stricter as the number of insiders increases. We use a cross-section of buyout bids during the period 2000–2006 to examine this hypothesis. We establish a positive link between the variation across events in suspicious pre-bid trading and the variation in the size of the financing syndicate for the

---

[10] We are grateful to an anonymous referee for pointing this out.

bid, and hence to the likely number of agents who would have had advance knowledge of it. We suggest that either imperfect competition among insiders or inefficient enforcement could lead to the findings documented here. A natural objective for future research is to attempt to distinguish between these factors. Characterizing this distinction is important both for regulatory objectives and for the general goal of understanding the dynamics of information asymmetry.

On a practical note, our work illuminates the interaction of several diverse trends in the recent evolution of the capital markets. The rise of private equity and the broadening of participation in syndicated lending during the period 2000–2007 were not isolated institutional developments. Rather, they had important implications beyond corporate finance that may have affected the dynamics of securities markets. We document the effects on returns and volume in several asset markets. A further important goal is to understand the implications of these findings for the dynamics of market liquidity.

Appendix A. Model solution and proof of Proposition 1

We are interested in solving the joint problem of the insiders, the market maker, and the regulators as described in the text. Recall the setting. At time zero the market maker must set the ask price a without knowing whether there will be zero insiders (no takeover), one insider, or N insiders. Regardless of the outcome, the market maker must sell whatever quantity is demanded at a. Here we denote the ex ante uncertain number of insiders by M. We are taking the probabilities of the outcomes 1 and N to be equal (conditional on a deal). The probability that there is a deal (a bid at price H) is denoted p_H.

A procedure for computing everybody's optimal policies is as follows.

1. Consider every point in the $V_1 \times V_N$ plane. At (V_1, V_N), solve the insider/market-maker problem for (x_1, x_N) and a.
   - This step will partition the plane into four regions (or zones) depending on the insiders' participation/non-participation.
   - Let Zone I be the set of points for which $(x_1=0, x_N=0)$. Let Zone II be the set $(x_1 = 0, x_N > 0)$. Let Zone III be the set $(x_1 > 0, x_N = 0)$. Let Zone IV be the set $(x_1 > 0, x_N > 0)$.
2. Given the responses x, compute the regulator's expected cost C at each point.
3. For a given budget K, find the locus of points in the $V_1 \times V_N$ plane having $C=K$.
4. Choose the minimum a value along this budget curve in each zone. The global optimum is the point achieving the minimum over these four.

The first step above is complicated by the fact that the stock's ask price a will not drop out of the insider's optimization. The ask depends on both trade quantities that the insiders adopt, x_1 and x_N. This means that

determining these two policies (and a) is a coupled problem. Also, one must check the participation constraint — that the insiders' objective is positive or else he does not trade — within this system, since the ask price will reflect that choice too.

Computation of the regulator's expected cost is straightforward. However, the isoquants (budget curves) cannot be characterized explicitly, and hence neither can the value of the objective along any such curve. The optimization thus requires a numerical grid search with an implicitly defined constraint. In fact, each zone must be searched separately since the objective (and the budget curves) are discontinuous across zones.

What we now aim to establish is that the optimum resides in Zone III under some reasonable conditions. This is Proposition 1. More specifically, the argument will show that the solution resides on the Zone III side of the border with Zone IV. The meaning of this is that it will be optimal for the regulator to pick V_N to be the highest value that is low enough to force the insiders not to trade when the outcome has N insiders.

The following is a breakdown of the steps in the proof.

- Fix a level of the regulator's objective or, equivalently, a level a^* of the ask price.
- Find an expression for the cost C^{III} of a policy in Zone III achieving a^*.
- Find an expression for the minimal cost C^{II} of a policy in Zone II achieving a^*. Show that it is greater than C^{III}.
- Using monotonicity properties of the cost and ask functions within Zones II and III, conclude that for any fixed level of cost, C^*, there is a policy in Zone III having cost C^* that dominates the lowest ask price achievable in Zone II at cost C^*.
- Find an expression for the minimal cost C^{IV} of a policy in Zone IV achieving a^*. Show that it is greater than C^{III}.
- Using monotonicity properties of the cost and ask functions within Zones IV and III, conclude that for any fixed level of cost, C^*, there is a policy in Zone III having cost C^* that dominates the lowest ask price achievable in Zone IV at cost C^*.

The argument in effect solves the dual problem of minimizing the cost for a given level of the objective. This locates the solution to the primal problem because of the monotonicity properties (to be explained below).

The key equation in the proof is a way of writing the first-order condition (FOC) for the insiders, which relies on the assumption that the distribution of noise trader demand is of the logistic form. Recall that the insiders determine the quantity traded, x, taking as given the number of insiders ($M=1$ or N), the ask price a, the enforcement ceilings, (V_1, V_N), and the policies of the other insiders (if any). Let F denote the CDF of noise trader demand. The ith insider's objective (expected payoff) if he trades x_i shares is

$$(H-a)x_i F\left(V_M - \sum_{j=1}^{M} x_j\right) - P_0 \left[1 - F\left(V_M - \sum_{j=1}^{M} x_j\right)\right].$$

(If he trades no shares, his payoff is zero.) So, conditional on participation, the FOC can be written

$$x_i + \frac{P_0}{(H-a)} = \frac{F'}{F}.$$

Applying the invertibility property of F'/F (cf. Assumption 1 in the text), the FOC says

$$se^{(V_M - \sum_{j=1}^{M} x_j - m)/s} = x_i + \frac{P_0}{(H-a)} - s. \tag{3}$$

If the outcome is $M=1$,

$$se^{(V_1 - x_1 - m)/s} = x_1 + \frac{P_0}{(H-a)} - s. \tag{4}$$

If the outcome is $M=N$,

$$se^{(V_N - Nx_N - m)/s} = x_N + \frac{P_0}{(H-a)} - s \tag{5}$$

since all insiders choose the same policy in equilibrium.

These equations provide necessary conditions for the optimal insider responses given a. It is easy to check that they have unique solutions and that these solutions are indeed maxima (i.e., the second-order condition holds). What is not clear is whether the solutions satisfy the participation constraint, that is, whether they imply a positive objective function. It turns out that this is very easy to check, thanks to the following lemma.

Lemma A.1. *The FOC yields a positive objective function if and only if its solution x^* exceeds s.*

This is easy to prove by just plugging the FOC back into the expression for F and rearranging. The result completely characterizes the borders of the four zones.

As noted above, it is still not easy to actually find x^*. Given V_1 and V_N, one has to solve for x_1, x_N, and a simultaneously, where the system is closed by the equation

$$a = \frac{\bar{S} + H\frac{EX}{EY}}{1 + \frac{EX}{EY}} = \frac{\bar{S} + \frac{Hp_H}{2m}(x_1 + Nx_N)}{1 + \frac{p_H}{2m}(x_1 + Nx_N)}.$$

The quantity we care about, because it appears in Eqs. (3)–(5), is a little simpler:

$$\frac{P_0}{(H-a)} = \frac{P_0}{(H-\bar{S})}\left[1 + \frac{p_H}{2m}(x_1 + Nx_N)\right] \equiv \pi + \theta(x_1 + Nx_N). \tag{6}$$

This yields a nonlinear set of equations for the policies, x. However the FOCs only pertain if the solution has a positive objective. Using the lemma, this is equivalent to replacing the (stylized) equation $x = FOC^{-1}(a)$ by $x = FOC^{-1}(a)1_{\{FOC^{-1}(a) \geq s\}}$.

Now, proceeding with the steps in the proof, we want to compute the regulator's cost for a policy on the border of Zones III and IV having a specified objective which we will denote a^-. This is equivalent to specifying a target amount of total expected insider trading $(p_H/2)(x_1 + Nx_N)$. We will call λ the target for $x_1 + Nx_N$. In Zone III, $x_N = 0$. Specifying a^- or λ is thus the same as specifying the target amount of trading x_1. Call this value x_1^-.

Knowing the desired x_1 and knowing $x_N = 0$ tells us right away the corresponding ceiling V_1 via the FOC for x_1,

which we can write

$$V_1 = s \log\left(\frac{P_0}{(H-a^-)} + x_1^- - s\right) + x_1^- + m - s \log s.$$

The cost of the policy also depends on V_N. We will choose V_N to be the highest possible value that keeps us in Zone III. This choice means that V_N will be epsilon less than the value inside Zone IV having solution $x_N = s$. That gives us a FOC for V_N which says

$$V_N = s \log\left(\frac{P_0}{(H-a^+)} + s - s\right) + Ns + m - s \log s.$$

Unfortunately, this equation depends on the ask price, which is no longer a^- since we are now at a point having $x_N > 0$. To find a^+ we observe that the amount of trading x_1^+ must satisfy

$$V_1 = s \log\left(\frac{P_0}{(H-a^+)} + x_1^+ - s\right) + x_1^+ + m - s \log s.$$

Subtracting the two equations for V_1 and using

$$\frac{P_0}{(H-a^-)} = \pi + \theta(x_1^- + 0) \quad \text{and} \quad \frac{P_0}{(H-a^+)} = \pi + \theta(x_1^+ + Ns)$$

yields an expression for x_1^+ in terms of x_1^-. We will give this expression below. For now, assume that we have solved it and have identified $\Delta^{III} \equiv (x_1^+ - x_1^-)/s$. We then have enough information to explicitly compute the cost of our policy.

The expected cost of any enforcement policy is

$$\frac{1}{2}(c_0 + c_1)(1 - F(V_1 - x_1)) + \frac{1}{2}(c_0 + c_1 N)(1 - F(V_N - Nx_N)).$$

(We now drop the factor of 1/2, which plays no role.) Using our specific functional form for F and the equations for the V just derived, we can evaluate the first and second terms, which we will call C_1^{III} and C_N^{III}. Omitting the algebra, we find

$$C_1^{III} = \frac{c_0 + c_1}{1 + \frac{1}{s}[(\pi - s) + (1+\theta)x_1^-]}, \tag{7}$$

$$C_N^{III} = \frac{c_0 + c_1 N}{1 + \frac{e^N}{s}[(\pi + Ns\theta) + \theta(x_1^- + s\Delta^{III})]}. \tag{8}$$

Given these costs, the next step is to find the lowest achievable cost of any policy in Zone II that also achieves the objective a^-. To locate that policy, we need two auxiliary results.

First, in this zone, level curves of the ask function are lines of fixed V_N. This follows from writing the FOC for x_N as

$$V_N = s \log\left(\frac{P_0}{(H-a)} + x_N - s\right) + Nx_N + m - s \log s$$

and using $P_0/(H-a) = \pi + \theta Nx_N$. This shows that all outcomes with the same x_N (and thus the same a) have the same V.

Second, costs in this zone are strictly decreasing in V_1 for any fixed V_N. This follows from inspection of the cost function, using the fact that V_1 has no effect on a (since all V_1 in this region enforce $x_1 = 0$).

Together these two facts imply that along any level curve of fixed a, the minimum cost occurs at the boundary, i.e., the highest V_1 point that is low enough to force $x_1 = 0$. So it suffices to locate the lowest-cost point on the boundary achieving a^-. We will then compute its cost and compare it to C^{III}.

Given an objective of a^- or, equivalently, given that Nx_N^- equals λ, the FOC

$$V_N = s \log\left(\frac{P_0}{(H-a^-)} + x_N^- - s\right) + Nx_N^- + m - s \log s$$

pins down V_N. Likewise, the requirement that we are on the boundary means that if we raise V_1 by epsilon we will get $x_1 = s$. Hence,

$$V_N = s \log\left(\frac{P_0}{(H-a^+)} + s - s\right) + s + m - s \log s.$$

As before, this involves a new value a^+ that holds across the boundary. To find it, we observe that inside the boundary we also have

$$V_1 = s \log\left(\frac{P_0}{(H-a^+)} + x_N^+ - s\right) + Nx_N^+ + m - s \log s.$$

Putting the two equations for V_N together and using $P_0/(H-a^+) = \pi + \theta(s + Nx_N^+)$ produces an expression for x_N^+ in terms of x_N^-. As above, we will not solve it for now, but instead use it to (implicitly) define $\Delta^{II} \equiv N(x_N^+ - x_N^-)/s$.

When we now plug our values for V_1 and V_N into the distribution function F to compute the expected cost of the policy, we find the terms

$$C_1^{II} = \frac{c_0 + c_1}{1 + \frac{e}{s}\left[(\pi + s\theta) + \theta N\left(x_N^- + \frac{1}{N}s\Delta^{II}\right)\right]}, \tag{9}$$

$$C_N^{II} = \frac{c_0 + c_1 N}{1 + \frac{1}{s}\left[(\pi - s) + \left(\frac{1}{N} + \theta\right)Nx_N^-\right]}. \tag{10}$$

We can now compare terms and find conditions such that C^{II} exceeds C^{III} term-by-term. The expressions are to be compared using the same target objective, so the value Nx_N^- in the C^{II} terms must equal x_1^- in the C^{III} terms. Their common value is λ. Note that term-by-term dominance is stronger than necessary.

Comparing (7) and (10) immediately shows that the latter has both a bigger numerator and a smaller denominator, so it is clearly more expensive. Comparing (8) and (9) is a little harder.

We proceed as follows. First, observe that the numerator in Eq. (8) is $c_0 + c_1 N$. Dividing the numerator and denominator by N makes this numerator strictly smaller than the numerator for (9) which is just $c_0 + c_1$. So it suffices to show that the denominator of (8) scaled by $1/N$ is bigger than that of (9). That is, we need

$$\frac{1}{N} + \frac{e^N}{Ns}[(\pi + Ns\theta) + \theta(\lambda + s\Delta^{III})] > 1 + \frac{e}{s}[(\pi + s\theta) + \theta(\lambda + s\Delta^{II})].$$

We split this in two pieces to deal with the Δ terms separately; and

$$\frac{1}{N} + \frac{e^N}{Ns}[\pi + Ns\theta] > 1 + \frac{e}{s}[\pi + s\theta]$$

as long as *either* $e^{N-1} > 1 + 1/\theta$ or $e^{N-1}/N > 1 + s/\pi$. (Note that $e^N/N > e$ for integers greater than one.) We impose the former condition (Assumption 3 in the text) since it also comes up below.

Turning to the Δ terms, recall that the Δ's are each the solution to a pair of FOCs. We now give them specifically. Subtracting the two equations

$$V_1 = s \log(\pi + \theta(x_1^-) + x_1^- - s) + x_1^- + m - s \log s,$$

$$V_1 = s \log(\pi + \theta(x_1^+ + Ns) + x_1^+ - s) + x_1^+ + m - s \log s$$

and exponentiating gives

$$e^{-\Delta^{III}} = \frac{(\pi - s + Ns\theta) + (1+\theta)(x_1^- + s\Delta^{III})}{(\pi - s) + (1+\theta)x_1^-}.$$

If we approximate the left-hand side by $1 - \Delta^{III}$, we have a linear equation for Δ^{III}. Note that the exact solution will be negative (as will the approximation), because Δ^{III} is the change in insider trading under the $M=1$ outcome when the enforcement policy has had V_N relaxed such that x_N increases from 0 to s. That change raises the market maker's ask, which lowers x_1. The negative sign and the concavity of the left side guarantees that the approximate solution will be more negative. That is, the approximation error is itself negative and is bounded by the approximation itself. Solving the linearized equation, the approximation is

$$\hat{\Delta}^{III} = \frac{-Ns\theta}{(\pi + s\theta) + (1+\theta)x_1^-}. \tag{11}$$

Proceeding likewise for Δ^{II} yields the approximate solution

$$\hat{\Delta}^{II} = \frac{-Ns\theta}{s + N(\pi - s + s\theta) + (1+N\theta)Nx_N^-}. \tag{12}$$

Now the terms we are comparing in (8) and (9) are $(e^N/N)(\lambda + s\Delta^{III})$ and $(e/1)(\lambda + s\Delta^{II})$, and we want the former to be larger, or

$$\frac{e^{N-1}}{N} > \frac{1 + \frac{s\Delta^{II}}{\lambda}}{1 + \frac{s\Delta^{III}}{\lambda}}.$$

Noting that Δ^{II} is negative and that our approximation to Δ^{III} is too negative, it suffices if

$$\frac{e^{N-1}}{N} > \frac{1}{1 + \frac{s\hat{\Delta}^{III}}{\lambda}} = \frac{1}{1 - \frac{Ns^2\theta/\lambda}{(\pi + s\theta) + (1+\theta)\lambda}}$$

or $$\frac{e^{N-1}}{N}\left(1 - \frac{Ns^2\theta/\lambda}{(\pi + s\theta) + (1+\theta)\lambda}\right) > 1.$$

The left side of the last expression exceeds $(e^{N-1}/N)(1 - Ns^2/(s\lambda + \lambda^2))$. And, since $\lambda = Nx_N^- > Ns$, this exceeds $(e^{N-1}/N)(1 - 1/(N+1))$. So we have the desired inequality as long as $e^{N-1}/(N+1) > 1$, which is true for integers greater than two, which is Assumption 2.

We have now shown that, given a fixed level of the ask price, a policy achieving that level can be found in Zone III with a lower expected cost than the lowest-expected-cost policy in Zone II that achieves the same target.

To return from the dual problem to the primal, given any fixed level, K, of expected cost, suppose one locates the lowest-ask policy, $(V_1^{K,II}, V_N^{K,II})$, having that cost in Zone II. Denote that ask a_K. Consider the isoquant of the ask function upon which that policy lies. We have seen that along any level curve of fixed a, the minimum cost occurs at the boundary. Moving along that level curve thus takes us to a point on the boundary with a cost, $\hat{K}$, not greater than K and the same ask, a_K, as the lowest ask achievable with cost equal to K in this zone. We have further located a policy in Zone III with cost $\hat{\hat{K}}$ less than $\hat{K}$ having an ask equal to a_K. We can therefore move along the ask level curve, which are lines of constant V_1, until the cost rises to the original level K. The two isoquants defined by cost = K and ask = a_K thus intersect in the interior of Zone III. Any point along the cost isoquant in the direction of lower V_1 will thus have a strictly better objective than a_k. Hence, the optimal policy for a regulator with budget constraint K cannot be a point in Zone II.

Completing our proof now requires likewise ruling out optimal policies in Zone IV where both $x_1 > 0$ and $x_N > 0$. The steps will be essentially the same. We start by locating a lowest-cost policy in this region for a given target level a or equivalently having $x_1 + Nx_N$ equal to a target λ. The relation between the two is (cf. Eq. (6))

$$\frac{P_0}{(H-a)} = \pi + \theta\lambda.$$

In Zone IV, the first-order conditions for both x_1 and x_N hold since neither is constrained. As we did before, we use the FOCs to substitute in to the expression for $(1-F)$ – the probability of being caught – which appears in the cost function. This yields the two cost terms:

$$C_1^{IV} = \frac{c_0 + c_1}{1 + \frac{1}{s}[(\pi - s)\theta + \theta\lambda + x_1]}, \tag{13}$$

$$C_N^{IV} = \frac{c_0 + c_1 N}{1 + \frac{1}{s}\left[(\pi - s) + \left(\frac{1}{N} + \theta\right)\lambda - \frac{x_1}{N}\right]}. \tag{14}$$

Comparing terms (13) and (7), the only difference is that the latter has a λ in the denominator where the former has x_1. But $\lambda > x_1$. So the Zone III term is always smaller. Comparing the denominators of (14) and (8), we want to show

$$e^N[(\pi + Ns\theta) + \theta(\lambda + s\Delta^{III})] > (\pi - s) + \left(\frac{1}{N} + \theta\right)\lambda - \frac{x_1}{N}.$$

We need this to hold at all points on the curve of fixed λ. So the most stringent requirement is that it must hold when x_1 is set to its minimum achievable value in Zone IV, which is s. Recall also that

$$0 > \Delta^{III} > \hat{\Delta}^{III}$$

So it suffices to establish the inequality with our linearized approximation $\hat{\Delta}^{III}$ on the left side. The desired inequality can then be expressed as

$$-e^N s\theta\hat{\Delta}^{III} < e^N (Ns\theta) + e^N\pi + e^N\theta\lambda - (\pi - s) - \theta\lambda - \frac{\lambda}{N} + \frac{s}{N}. \tag{15}$$

The left side is

$$e^N s\theta\frac{Ns\theta}{(\pi + s\theta) + (1 + \theta)x_1^-} < e^N(Ns\theta).$$

So the inequality in (15) holds as long as the sum of the last six terms on the right is positive. For this, it suffices if the terms involving λ are positive, i.e., if

$$\lambda\left(e^N\theta - \theta - \frac{1}{N}\right) > 0 \quad \text{or} \quad e^N > 1 + \frac{1}{\theta N}.$$

But $e^N > e^{N-1} > 1 + 1/\theta > 1 + 1/\theta N$, where the middle inequality is imposed by Assumption 3.

With this, we have shown that, given a fixed level of the ask price, a policy achieving that level can be found in Zone III with a lower expected cost than the lowest expected cost policy in Zone IV that achieves the same target.

Now, the final step. Given any fixed level, K, of expected cost, suppose one locates the lowest-ask policy, $(V_1^{K,IV}, V_N^{K,IV})$, having that cost in Zone IV. Denote that ask a_K. We have located a policy in Zone III with cost $\hat{K}$ less than K having an ask equal to a_K. We can therefore move along the ask-isoquant in Zone III, which are lines of constant V_1, until the cost rises to the original level K. The cost=K and ask=a_K isoquants thus intersect in the interior of Zone III. Any point along the cost isoquant in the direction of lower V_1 will thus have a strictly better objective than a_k. Hence, the optimal policy for a regulator with budget constraint K cannot be a point in Zone IV.

This completes the proof.

Appendix B. Data sources and details

This appendix elaborates on some aspects of our use of the primary databases employed in the empirical work.

Event sample:

In selecting events from the Thomson Financial database, we specify acquisition bids for public U.S. companies and impose the following criteria:

- Value of the bid must exceed 100 million dollars.
- Acquiror must be seeking a controlling stake.
- Bid must be an offer for publicly held securities (not a private holding).
- Bidder must be a private entity or group (perhaps including individuals).
- Bidder's "type" must be given as "Financial" or its business description must make clear that it is an investment vehicle.
- Buyer cannot be a bank, insurance company, or real estate investment trust.

In instances where a target is subject to more than one bid in our sample, we select only one event.[11] For multiple-bid targets, we follow these rules:

- Restrict attention to initial bids by each entity, i.e., not sweetened or subsequent bids.

---

[11] There is one exception: Petco Animal Supplies Inc. was actually taken private twice in our sample, having been re-floated in between.

- If there is more than one bid by private acquirors, take the successful one unless it cannot be determined which bid was successful, or another private bid occurred less than 14 days before the successful one. In these cases, take the earliest bid.
- If a non-private bid occurred less than 14 days before a private one, discard this target.

Our rule favors successful bids only because these deals are more likely to enable us to identify providers of debt finance (as described in Section 3.1).

Counting bank relationships:

We count bank relationships by selecting active tranches of syndicated loans from the LPC DealScan database. Loan participants are then assigned a unique code in accordance with the rules below.

- All subsidiaries/branches/operations of one company are grouped together, e.g., Bank of Nova Scotia, Scotia Capital, BNS International, and Scotiabank Inc. are treated as a single entity.
- Any joint entity involving two or more firms (which themselves appear separately) are treated as the subsidiary of the firm listed first: KZH-Cypress, KZH-Soleil, etc., are all treated as subsidiaries of KZH.
- Firms before and after a merger or acquisition are treated as different companies; e.g., Ag First merged with Farm Credit in 1992 to form AgFirst Farm Credit Bank. The three of them are considered separate lenders.

Corporate bond returns: We identify all corporate bonds of a firm matched by ticker symbol to the TRACE database. To obtain the firm-level bond return, individual bond returns are weighted each day in proportion to their respective issued par value (obtained from the Mergent fixed-income database). Individual bond returns are computed every day as the percentage change from the most recent available closing price for the bond to the given day's closing (that is, last trade) price. Note that this could result in the use of stale prices for some issues that do not trade at (or near) the close of each day. However, if a bond has no trade on a given date, then its return is taken to be zero.

References

Acharya, V.V., Franks, J., Servaes, H., 2007. Private equity: boom and bust? Journal of Applied Corporate Finance 19, 44–53.

Acharya, V.V., Johnson, T.C., 2007. Insider trading in credit derivatives. Journal of Financial Economics 84, 110–141.

Altman, E., 2007. Global debt markets in 2007: new paradigm or the great credit bubble? Journal of Applied Corporate Finance 19, 17–31.

Amihud, Y., 2002. Illiquidity and stock returns: cross-section and time-series effects. Journal of Financial Markets 5, 31–56.

Andrade, G., Mitchell, M., Stafford, E., 2001. New evidence and perspectives on mergers. Journal of Economic Perspectives 15, 103–120.

Arnold, T.M., Erwin, G.R., Nail, L.A., Bos, T., 2000. Speculation or insider trading: informed trading in options markets preceding tender offer announcements. Working paper. University of Alabama, Birmingham.

Back, K., Cao, H.H., Willard, G.A., 2000. Imperfect competition among informed traders. Journal of Finance 55, 2117–2155.

Baruch, S., 2002. Insider trading and risk aversion. Journal of Financial Markets 5, 451–464.

Berndt, A., Ostrovnaya, A., 2007. Information flow between credit default swap, option and equity markets. Working paper, Carnegie Mellon, Pittsburgh, PA.

Bond, P., Hagerty, K., 2005. Preventing crime waves. Working paper, Northwestern University, Evanston, IL.

Cao, C., Chen, Z., Griffin, J.M., 2005. Informational content of option volume prior to takeovers. Journal of Business 78, 1073–1109.

DeMarzo, P.M., Fishman, M.J., Hagerty, K., 1998. The optimal enforcement of insider trading regulations. Journal of Political Economy 106, 602–632.

Foster, F.D., Viswanathan, S., 1996. Strategic trading when agents forecast the forecasts of others. Journal of Finance 51, 1437–1478.

Gao, Y., Oler, D., 2004. Trading in target stocks before takeover announcements. Working paper, Cornell University, Ithaca, NY.

Glaeser, E., Sacerdote, B., Scheinkman, J., 1996. Crime and social interactions. Quarterly Journal of Economics 111, 507–548.

Heckman, J.J., 1978. Sample selection bias as a specification error. Econometrica 47, 153–161.

Holden, C.W., Subrahmanyam, A., 1992. Long-lived private information and imperfect competition. Journal of Finance 47, 247–270.

Holden, C.W., Subrahmanyam, A., 1994. Risk aversion, imperfect competition and long-lived information. Economics Letters 44, 181–190.

Ivashina, V., Sun, Z., 2007. Institutional stock trading on loan market information. Working paper, Harvard Business School, Cambridge, MA.

Jarrell, G.A., Poulsen, A.B., 1989. Stock trading before the announcement of tender offers: insider trading or market anticipation? Journal of Law, Economics, & Organization 5, 225–248.

Jayaraman, N., Frye, M.B., Sabherwal, S., 2001. Informed trading around merger announcements: an empirical test using transaction volume and open interest in options markets. Financial Review 36, 4574.

Keown, A.J., Pinkerton, J.M., 1981. Merger announcements and insider trading activity: an empirical investigation. Journal of Finance 22, 855–869.

Mandelker, G., 1974. Risk and return: the case of merging firms. Journal of Financial Economics 1, 303–336.

Meulbroek, L.K., 1992. An empirical analysis of illegal insider trading. Journal of Finance 47, 1661–1699.

Poteshman, A.M., 2006. Unusual options market activity and the terrorist attacks of September 11, 2001. Journal of Business 79, 1703–1726.

Sah, R.K., 1991. Social osmosis and patterns of crime. Journal of Political Economy 99, 1272–1295.