Endogenous Limits to Arbitrage and Price Informativeness

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Abstract

Theory suggests that traders are reluctant to trade on negative private information about a corporate decision if their trading can cause a reversal of the decision. We provide evidence for such endogenous limits to arbitrage in the context of mergers. Starting from the observation that an acquirer termination fee increases the cost of canceling a transaction, potentially breaking the feedback loop between prices and decisions, we find the amount of firm-specific information in post-announcement acquirer stock prices to be significantly higher if a termination fee is used than when none is used.

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1 Introduction

The various versions of the market efficiency hypothesis state that security prices reflect available information. This view is consistent with a mechanism where information is collected and processed by market participants and their subsequent trading causes prices to reflect this information. The traditional view is that the process through which prices become informative ends here. A recent literature has investigated the possibility that corporate managers pay attention to stock prices and adjust their corporate decision in accordance with what they learn. When managerial decisions are affected by stock prices, there will be a feedback loop between prices and decisions that may impact the degree of market efficiency. Edmans, Goldstein, and Jiang (2015) argue that feedback between prices and decisions may deter investors from trading on private information. In a merger setting, the intuition for their argument is as follows: If speculators have private negative information about a proposed merger, they will short the buyer's stock. But if the acquirer management pays attention to the market reaction to the proposed merger, they will realize that the deal is bad and cancel the transaction. This leaves the speculators with a losing short position. Thus, the speculators will not short the stock in the first place, the acquirer price won't reflect as much negative information as is available to speculators, and the management will execute a suboptimal decision. Consequently, welfare-reducing endogenous limits to arbitrage have occurred.

This paper's main contribution is to provide empirical evidence on this particular form of limits to arbitrage and its impact on price informativeness. Endogenous limits to arbitrage will be absent if the feedback loop between stock prices and management decisions is broken. In the context of mergers, the feedback loop can be broken when an acquirer has agreed to compensate the seller if the transaction is cancelled by the acquirer. Such compensation is commonly referred to as a termination fee. A large acquirer termination fee will make deal cancellation more costly and, consequently, less likely. This should induce informed traders to trade on their information making prices informative. We use a large sample of merger announcements and compare the information content of post announcement stock prices for acquirers committed to pay a termination fee to the corresponding measure for acquirers without such commitment. We find that the post announcement stock price of acquirers in deals with acquirer termination fees contains considerably more firm specific information than the stock price of acquirers that have not agreed to pay a termination fee. In our sample, our price-informativeness measure more than doubles compared to the sample average when a termination fee is used. This finding is consistent with the existence of endogenous limits to arbitrage and demonstrates their significant impact on price efficiency.

We reinforce this evidence in several ways. First, the feedback from stock prices to management decisions is asymmetric. While a negative stock price response might induce the management to reconsider its decision, a positive response shows that the market agrees with management making deal cancellation less likely. Thus, only after receiving a negative signal, will speculators be concerned about their trading causing deal cancellation. It follows that only in such cases will a termination fee ease these concerns and increase trading and price informativeness. To test this refined implication of endogenous limits to arbitrage, we split the sample into a subsample of mergers likely to come with a negative signal and a subsample where a positive signal was more likely. In support of the theory, we find that the use of an acquirer termination fee increases price informativeness only in the bad-signal subsample, while the estimated effect is insignificant and close to zero in the good-signal subsample.

Second, we address endogeneity concerns arising from the fact that the inclusion of a termination fee is a choice. When negotiating the terms of a merger transaction, the acquirer and the seller negotiate over multiple issues simultaneously. The inclusion of an acquirer termination fee will in some transactions be part of the set of issues to agree upon. Thus, the inclusion of an acquirer termination fee in the merger agreement will be a choice variable related to other observable and unobservable variables. If the choice of including a termination fee is related to unobservable variables that also influence price informativeness, we are facing an endogeneity problem that can potentially bias our estimate of the effect of including a termination fee. A case in point are private benefits the management of the acquirer might derive from running a larger post-merger firm. In that case, the management is more likely to complete the transaction if the stock market reaction is negative, which should increase speculators' willingness to trade on private information and therefore increase price informativeness. Such private benefits are also likely to affect the bargaining process leading to the merger agreement. One can think of different ways private benefits could affect the bargaining outcome, with some economic mechanisms making the inclusion of an acquirer termination fee more likely while others making it less likely. Since the different mechanisms produce different biases in the OLS regression, the comparison between the OLS and the IV regression outcomes allows us to shed some light on the mechanism at work.

To deal with this potential endogeneity problem, we introduce instrumental variables for the usage of acquirer termination fees. In the 1994 case *Paramount v. QVC Network*, the court ruled that the stock lockup option granted by Paramount to a prospective bidder Viacom was invalid, while the $100 million termination fee was upheld. The court viewed the termination fee to be a reasonable amount to cover Viacom's expenses in connection with the merger process. After this court ruling, the number of mergers with acquirer termination fees increased sharply.

We use a dummy variable that equals one for the period 1994–2018 and zero for the period 1986–1993 as an instrumental variable for the usage of acquirer termination fee. We also identify a structural break in the influence of acquirer, target, and deal characteristics on the use of termination fees following the ruling. Consequently, we incorporate interaction terms of the dummy variable with these characteristics as additional instruments. These variables act as instruments for the acquirer termination fee in a standard instrumental variable approach and in an endogenous switching model. The latter is used because the assumptions underlying the instrumental variable approach do not fully agree with the suspected economic mechanism causing the endogeneity problem. The coefficient estimates of these two approaches greatly exceed the OLS-estimates and reveal an economically significant effect of endogenous limits to arbitrage. The sign of the OLS-bias is informative in itself and suggests that the target management recognizes opportunistic merger motives of the acquirer management. In such cases, the target management is less concerned about the transaction failing and instead uses its bargaining power to secure other concessions, rather than a termination fee, in the bargaining leading to the merger agreement.

Finally, we show that our results are not driven by the trading of merger arbitrageurs. Their trading involves buying the target's stock and shorting the stock of the acquirer in such a way that the short and long side perfectly match each other if and when the transaction closes. Merger arbitrageurs aim to pocket the gap typically seen between the target stock price right after the announcement and the offered acquisition price. While they care about the merger success probability, they do not care about the quality of the merger per se. Their strategy always involves shorting the bidder. It is not clear whether their trading makes prices informative in the same way the trading of informed traders does. For that reason, we investigate whether our earlier results hold in a subsample of mergers with non-listed targets. In these cases, merger arbitrageurs cannot acquire a long position in the target, making the merger arbitrage trade impossible. We find that all our findings also hold in this subsample.

This paper adds to the literature on the feedback between the stock market and corporate decisions. Dow and Gorton (1997), Subrahmanyam and Titman (2001), Bond, Goldstein, and Prescott (2010) and Dow, Goldstein, and Guembel (2017) show theoretically how stock prices may affect corporate decisions. Baker, Stein, and Wurgler (2003), Chen, Goldstein, and Jiang (2007) and Ferreira, Ferreira, and Raposo (2011) show that prices can affect corporate investment decisions as well as the corporate governance. Luo (2005) and Kau, Linck, and Rubin (2008) study the feedback effect in takeovers. They find that acquiring firms pay attention to the market reaction to the transaction announcement and document a larger probability of deal cancellation when bidder announcement return is negative. Edmans, Goldstein, and Jiang (2011) show, using an instrument for undervaluation, that prices can trigger an acquisition. While the previously mentioned empirical literature studies the feedback from security prices to corporate decisions, our paper provides evidence on how the feedback between prices and corporate decisions results in endogenous limits to arbitrage and, consequently, influences market efficiency.

A paper closer to ours is Gorton, Huang, and Kang (2017), which study the interaction between information production and CEO replacement. Assuming that informed traders always trade, the authors show that the link between trading and CEO replacement reduces traders incentive to get informed about CEO quality. The authors shed light on this mechanism by studying the empirical relation between CEO turnover and the probability of informed trading (PIN, see Easley, Hvidkjaer, and O'Hara, 2002). Our paper, and the model of Edmans, Goldstein, and Jiang (2015), differ in that we focus on whether feedback between prices and corporate decisions destroys the trading incentive of informed traders. Morover, since it has been shown that PIN struggles to capture price informativeness around corporate events known to generate relevant firm specific information, we follow Roll (1988) and Chen, Goldstein, and Jiang (2007) and use firm specific stock return variation as a measure of price informativeness.[1] This measure has been found to be highly correlated with stock prices' ability to predict firms' future earnings (Chen, Goldstein, and Jiang, 2007) and to be considered by a firm's management for investment decisions (Veldkamp, 2006).

Our paper is also related to the takeover deal protection literature, which has mostly been concerned with target termination fees. There is little work on understanding acquirer termination fees. Exceptions include Bates and Lemmon (2003) and Chen, Mahmudi, Virani, and Zhao (2021). Bates and Lemmon (2003) point out that an acquirer termination fee can be regarded as an insurance for targets. They support this view with evidence showing that acquirer termination provisions are more common in deals with higher cost of deal failure for targets. Chen, Mahmudi, Virani, and Zhao (2021) provide an alternative view. They point out that the acquirer termination fee provides the bidder with an option to cancel the transaction at the cost of paying the termination fee. We find support for the insurance motive. An acquirer termination fee is more likely if the acquirer has difficulties financing the deal or is about the same size as the target. In addition, we find evidence that deals that are more likely to be pursued for the private benefits of bidder management are less likely to include an acquirer termination fee.

The rest of the paper is organized as follows. Section 2 introduces the main empirical hypotheses. Section 3 describes our data. Section 4 outlines the empirical approach and section 5 presents our main empirical findings. Section 6 investigates the robustness of the main findings in a subsample where merger arbitrageurs are absent. Section 7 concludes.

[1] Aktas, de Bodt, Declerck, and Van Oppens (2007) find that PIN is lower before merger and acquisition announcements, which contradicts the evidence on information leakages prior to merger announcements. Collin-Dufresne (2015) finds that PIN is lower on days which Schedule 13D filers trade, and concludes that PIN does not reveal the presence of informed traders.

2 Hypotheses

The development of testable hypotheses and empirical investigation of endogenous limits to arbitrage is inspired by the model of Edmans, Goldstein, and Jiang (2015) (henceforth the EGJ model) and Boleslavsky, Kelly, and Taylor (2017). In the EGJ model, a speculator has private information about the cash flow effects of an investment announced by a firm. The informed speculator would like to trade on and profit from this information while the firm's management would like to learn more about the value of the investment from the stock market reaction. A speculator who has negative private information worries that the investment will be abandoned if his private information is revealed by selling or shorting the stock. The price recovery caused by canceling the investment will result in a loss for the speculator, who adjusts his trading behavior in such a way that trades are less likely to reveal his negative information when combined with the random trades of noise traders. His expected profit from trading therefore depends on the probability of revelation and trading costs, which can be substantial if shorting is involved. If the probability of revelation or the trading costs are too high, the speculator refrains from trading, prices won't be informative, and the management will execute a suboptimal decision. In such a scenario, the feedback loop between trading and corporate decisions endogenously caused limits to arbitrage and a welfare loss.

The primary goal of this paper is to investigate the existence of such endogenous limits to arbitrage and assess their impact on price informativeness. To empirically investigate this idea, we require a situation where the feedback loop is sometimes switched off. Anything that induces the management to hold on to a planned investment even if the market reacts negatively will lift the limits to arbitrage. Merger decisions provide such a situation. An acquirer termination fee is a device that creates a commitment for the acquirer to go through with an announced merger even if the market reacts negatively to the merger announcement. Since not all mergers involve acquirer termination fees, mergers are a suitable setup to test for the existence and consequences of endogenous limits to arbitrage. Mergers involving acquirer termination fees should represent trading opportunities for informed speculators and prices should reflect their information. Mergers without such termination fees, should be affected by endogenous limits to arbitrage and, as a consequence, prices should be less informative. If endogenous limits to arbitrage exist, then we should observe that:

Hypothesis 1a. Price informativeness is higher for acquirer stocks following the announcement of mergers that include an acquirer termination fee than for acquirers in mergers not including an acquirer termination fee.

That the bidder has agreed to an acquirer termination fee does not mean that it has tied its hands and will go through with the merger at any cost. Bidder management is likely to compare the value implications of the market signal to to the size of the termination fee. The larger the termination fee, the more likely this comparison favors continuing with the merger. Thus, if endogenous limits to arbitrage exists, then

Hypothesis 1b. The price informativeness is increasing in the size of the acquirer termination fee.

Only informed speculators with negative information need to worry about deal withdrawal. Positive private information will be welcomed by the management, making a withdrawal highly unlikely. We, therefore, expect endogenous limits to arbitrage to arise for bad mergers only:

Hypothesis 2. The hypothesized relation between price informativeness and acquirer termination fee is only present when speculators receive bad information about the merger. When speculators receive good information about the merger, price informativeness should be unaffected by the usage of an acquirer termination fee.

3 Data, variable definitions and descriptive statistics

3.1 Data and sample selection

Data on merger transactions is from Refinitiv SDC Platinum (formerly Thomson Reuters SDC Platinum). Table 1 details the sample selection procedure. We start with all mergers by U.S. and publicly listed acquirers announced between 1986 and 2018 where the deal value is available. This results in a sample of 22,254 mergers. Next we require the acquirer to be present in the data from Center for Research in Security Prices (CRSP) and in Compustat. We also require the bidder to be present in CRSP with common shares (share codes 10 or 11) and the announcement date to be within the CRSP time-series of daily returns. This reduces the sample to 17,550 mergers. We only keep the first announcement in a bidding contest, require stock price and number of shares to be available 42 days prior to the announcement (to compute market capitalization), require Compustat total assets to be available for the fiscal year-end prior to the announcement, the deal value divided by bidder market capitalization to be larger than 1%, and dropping utilities. This reduces the sample to 14,545 mergers. In most of our analysis, we include a set of independent variables (covariates) that further reduce the sample size. This gives a sample of 12,519 over the sample period 1986–2018.

Panel A of Figure 1 breaks down the number of mergers by year. The well known surge in merger activity during the 1990s is clearly present in our data. Panel B shows the number of transactions that includes an acquirer terminations fee by year. Prior to 1994 there are few termination agreements. The number jumps in 1994 and stays at a higher level after that. The jump in the usage of acquirer termination fees corresponds to the *Paramount v. QVC Network* ruling where the court found that the $100 million termination fee reasonably reflected Viacom's expenses in connection with the merger. We later exploit this as an instrument for the usage of acquirer termination fee.

3.2 Key variables

This section describes how we measure price informativeness, the existence of an acquirer termination fee agreement, and the size of an acquirer termination fee. We also provide descriptive statistics on these and other variables used in the study. All variable definitions are summarized in Appendix A Table A1.

Price informativeness. To empirically examine our hypotheses 1a, 1b and 2, we measure stock price informativeness as firm specific stock return variation (also referred to as price non-synchronicity). Roll (1988) finds that stock price movements, beyond general market movements, cannot be explained by identifiable news releases. Thus, firm specific price movements are consistent with speculators gathering and trading on their private information. The intuition is as follows: Stock return movements are driven by information from two sources, general public information and firm-specific private information. Public information is reflected in market return and industry return. The part of return that cannot be explained by market return and industry return must be driven by the firm-specific private information. Based on this argument, if the stock prices of a firm contain more firm-specific private information, the return of the stock should be less correlated with the market or industry return, and therefore has greater firm-specific return variation. Prior empirical studies have provided evidence supporting the notion that firm-specific return variation reflects private information and has features important for our study. For example, Durnev, Morck, Yeung, and Zarowin (2003) find that firm-specific return variation indicates that prices are informative since it is associated with stock prices' ability to predict firms' future earnings. Chen, Goldstein, and Jiang (2007) find that the amount of private information in stock prices has a strong effect on the sensitivity of investments to stock prices, suggesting that the management "listens more to the market" when firm-specific return variation is high.[2]

[2] Other papers have used similar proxies and documented effects that are consistent with the view that these proxies capture the amount of private information in prices. E.g., Morck, Yeung, and Yu (2000) find that countries with a well-developed financial system in general have higher firm-specific return variation. Veldkamp (2006) shows that traders rely more on signals common to many firms if firm-specific information is hard to obtain. This leads to a drop in price non-synchronicity. See Chen, Goldstein, and Jiang (2007) for further examples.

We capture firm specific stock return on day t using the following regression model:

$$r_{it} = a_i + b_i r_{mt} + c_{ij} r_{jt} + s_i smb_t + h_i hml_t + \epsilon_{it}, \tag{1}$$

where r_{it} is the daily excess return for firm i on day t, r_{mt} is the market excess return, smb_t is the size factor, hml_t is the book-to-market factor, r_{jt} is the return on industry j in which firm i is a member, and $\{a, b, c, s, h\}$ are parameters to be estimated.[3] The firm specific stock return is measured using the residuals, $\hat{\epsilon}_{it}$, from the above regression. The model is estimated using daily trading dates $t+2$ through $t+22$, where $t=0$ is the merger announcement date. The R-squared from this regression, denoted R_i^2, measures the variation in r_{it} explained by the market, the industry, the size factor and the book-to-market factor. Consequently, $(1-R_i^2)$ captures the variation in r_{it} not explained by the market and the firm's own industry. We use

$$y_i = \ln\left(\frac{1-R_i^2}{R_i^2}\right) \tag{2}$$

as our measure of stock price informativeness. The higher this number the more firm specific information is captured by stock prices.

Termination Fee D. Dummy variable equal to one if the merger agreement includes a bidder termination fee and the Termination Fee Size, as defined below, is larger than its 25%-quantile (of all nonzero Termination Fee Size observations) and zero otherwise. With this cutoff we aim to identify the fee size necessary to induce the bidder management to hold on to a deal. We find that the use of a termination fee leads to a strong and significant reduction of the deal cancellation probability if the fee exceeds its 25%-quantile (see section 3.3 for the details).

Termination Fee Size. Hypothesis 1b captures the idea that bidder managers are less likely to cancel a transaction if they have agreed to a large termination fee. We normalize the termination fee by dividing its dollar value by the market capitalization of the acquirer, measured 42 trading dates prior to the merger announcement. For the regression analysis, we want to include termination

[3]Industries are defined according to the Fama-French 30 industry portfolios.

fees that are sufficiently large to be considered costly to pay by the acquirer. For that reason, the variable Termination Fee Size is computed as the normalized termination fee multiplied by the Termination Fee dummy defined above.

Panel A of Table 2 reports descriptive statistics for the termination fee measures and the dependent variables used in this study. We have 12,519 observations for the "Termination Fee D" variable. An average value of 0.07 implies that 7% of mergers in our sample include an acquirer termination fee with a size exceeding the 25%-quantile. The second row in Panel A shows that, conditional on Termination Fee D being equal to one, the termination fee is on average 3% of acquirer market capitalization. This is a significant fee that will have to be paid to the target if the acquirer decides to cancel the transaction. It exceeds the 2.24% average dividend yield of the S&P500 index recorded over that sample period.

Panel A also shows that mean and median price informativeness are very similar at 0.73 and 0.72 respectively. The mean price informativeness for mergers with an acquirer termination fee is not statistically different from that for mergers without a termination fee. To uncover the effect of a termination fee, we need to partial out the impact of other determinants of price informativeness. Panel A also reports the frequency of withdrawn transactions. On average 9% of announced transactions get canceled. The cancellation frequency in the sample of transactions with an acquirer termination fee and the sample without such a fee are almost the same. However, as we show in the next section, the usage of an acquirer termination fee has a highly significant and sizable impact on the withdrawal probability if other determinants of deal cancellation are properly controlled for.

Panel B of Table 2 presents summary statistics for all variables used as control variables in this paper. Notice that the deal value is significantly larger for deals with Termination Fee D equal to one. In other words, small deals tend not to contain an acquirer termination fee. Similarly, transactions with publicly listed targets are more likely to include an acquirer termination fee. Acquirers with a termination fee also tend to have worse past one-year stock market performance, higher runup and lower announcement abnormal return.

3.3 Acquirer termination fee and likelihood of deal cancellation

The premise of our analysis is that an acquirer termination fee makes it less likely that the acquirer cancels the transaction. Only then is a termination fee a commitment to go through with the merger—potentially inducing informed speculators to trade on their private information. In this section, we show that deals involving an acquirer termination fee are considerably less likely to be canceled.

Define Withdrawn as a dummy variable equal to one if the deal was canceled and zero if the deal was completed. Table 3 documents the results of a probit regression with Withdrawn as the dependent variable. Column (1) of Table 3 reports the results of a probit model which uses four dummy variables based on the four quartiles of normalized termination fee (i.e., the dollar value of the termination fee divided by the market capitalization of the acquirer.) The dummy variables representing the second to the fourth highest quartiles have a significantly negative effect on the withdrawal probability. This result guided our decision to set our termination fee dummy variable equal to one for all deals that have a termination fee exceeding the 25%-quantile. Column (2) of Table 3 shows that the probability of deal cancellation is lower if the transaction includes a large termination fee than when it does not (z-statistic of -3.86). Using a termination fee decreases the probability of deal cancellation by 34% (average marginal effect). This finding is consistent with our premise that a termination fee increases the cost of canceling a merger and thus cancellation becomes less likely. Column (3) shows that the termination fee dummy has an even larger effect on the probability of cancellation when also controlling for termination fee size. The positive marginal effect of termination fee size is consistent with the finding in Column (1) that the effect of a termination fee is not as pronounced if a very large fee is used.

The likelihood of deal cancellation is also increasing in the value of the transaction and is higher if the target is listed. Unsurprisingly, the deal is more likely to be canceled if the bid for the target is unsolicited. The deal is less likely to be canceled if the market capitalization of the bidder is large and if the bidder and the target operate in the same industry. Finally, we find that the announcement effect has no impact on the cancellation probability but the runup prior to the announcement does.

Finally, using the realizations of the covariates and setting the termination fee dummy to zero, we find from the probit model, that the majority of transactions with a termination fee has a predicted probability of deal cancellation above the mean predicted probability. In other words, termination fees are mostly used when the probability of deal cancellation is high. But when a termination fee is used, it comes with a significantly reduced probability of deal cancellation.

In sum, transactions with an acquirer termination fee have a significantly lower probability of deal cancellation than otherwise similar transactions without a termination fee. Thus, the findings in Table 3 justify using acquirer termination fees to separate feedback equilibria, where endogenous limits to arbitrage might exists, from no feedback equilibria, where such limits don't exist. With that result established, we can move on to test for the existence of limits to arbitrage by analyzing price informativeness.

4 Empirical Strategy

4.1 Ordinary Least Squares estimates

In the previous section, we showed that a deal is considerably more likely to be completed if it includes an acquirer termination fee. Hence, the acquirer termination fee has the potential to shut down the endogenous-limits-to-arbitrage mechanism—if it exists. Thus, we should observe more informative prices after a merger announcement if the merger agreement contains an acquirer termination fee. This is Hypothesis 1a. A simple approach to test this hypothesis is to regress price informativeness of the bidder in merger i, denoted by y_i, on the termination fee dummy t_i and a vector of control variables x_i:

$$y_i = \alpha + x_i\beta + \gamma t_i + e_i \tag{3}$$

The coefficient γ represents the effect of shutting down the endogenous-limits-to-arbitrage mechanism with a termination fee. However, specific assumptions need to hold for γ to capture that effect. A particular concern is that the termination fee choice must not depend on unobservable variables that also affect price informativeness. We discuss why economic theory suggests such unobservable factors are confounding the termination fee choice and how we deal with this challenge in the next sections. To facilitate this discussion and derive estimable equations, we cast the relation between termination fee use and price informativeness in the form of the potential outcome model (4) and (5):[4]

$$y_{0i} = \alpha_0 + x_i\beta_0 + u_{0i} \quad \text{if} \quad t_i = 0 \tag{4}$$

$$y_{1i} = \alpha_1 + x_i\beta_1 + u_{1i} \quad \text{if} \quad t_i = 1 \tag{5}$$

where y_{1i} and y_{0i} are the price informative measures after the announcement of merger i if a termination fee is used ($t_i = 1$) or not used ($t_i = 0$). The vector x_i contains other determinants of price informativeness, whose effects might depend on the presence of a termination fee. The error terms u_{0i} and u_{1i} satisfy $E[u_{gi}|x_i] = 0$ for $g = 0, 1$, are unobservable to us, but potentially known by the parties involved in the transaction and trading, and possibly related to the termination fee choice (see section 4.2).

If the inclusion of a termination fee induces informed traders to participate and make prices more informative, then the difference $y_{1i} - y_{0i}$ should be positive on average, i.e., $E[y_{1i} - y_{0i}] > 0$. $E[y_{1i} - y_{0i}]$ corresponds to the average treatment effect (ATE), where the treatment is the use of an acquirer termination fee. A slightly less demanding but sufficient test for the presence of endogenous limits to arbitrage is to test whether the deals including a termination fee exhibited a higher price informativeness, i.e., $E[y_{1i} - y_{0i} \mid t_i = 1] > 0$. Notice that $E[y_{1i} - y_{0i} \mid t_i = 1]$ corresponds to the treatment effect on the treated (ATT) and only requires the counterfactual y_{0i} for the treated. Thus, Hypothesis 1a can be recast as the ATE or the ATT being positive.

Since we observe only y_{1i} for a merger with a termination fee and only y_{0i} for a merger without a termination fee, the expectation of the respective counterfactual outcome has to be estimated from the data. In the cross section of mergers, we can express the observed price informativeness as: $y_i = y_{0i} + t_i\,(y_{1i} - y_{i0})$. Plugging equations (4) and (5) into this relation allows us to write

[4]The discussion draws on Wooldridge (2010) and Maddala (1983).

observed price informativeness in the form of a switching regression:

$$y_i = \alpha_0 + x_i\beta_0 + (\alpha_1 - \alpha_0)t_i + t_i x_i(\beta_1 - \beta_0) + u_{0i} + t_i(u_{1i} - u_{0i}) \tag{6}$$

If we are able to consistently estimate the parameters of this switching regression from the set of observable variables (y_i, t_i, x_i), then we can calculate the treatment effects and test our hypotheses. To succeed, we need to solve the problem that t_i might be correlated with the error terms u_{0i} and u_{1i}. The same problem plagues the dummy-variable regression (3), where $\beta_1 = \beta_0$ and $\gamma = \alpha_1 - \alpha_0$. We will approach the estimation problem in three different ways: OLS, using instrumental variables, and an endogenous switching regression. The later two approaches were chosen because finance theory and previous empirical results suggest that, in our merger sample, the first approach is vulnerable to endogeneity problems.

Our first approach (OLS) to estimating the treatment effects is based on the assumption that, conditional on a set of covariates x_i, (y_{i0}, y_{i1}) and t_i are mean-independent. It is common to refer to this as the unconfoundedness assumption.[5] The assumption is violated if a variable is omitted that influences both price informativeness and termination fee choice. Given model (6), unconfoundedness implies that t_i and u_{i1}, as well as t_i and u_{i0} are independent conditional on x_i and we can subsume the error terms as in equation (7):

$$y_i = \alpha_0 + x_i\beta_0 + (\alpha_1 - \alpha_0)t_i + t_i x_i(\beta_1 - \beta_0) + \epsilon_i \tag{7}$$

Given the unconfoundedness assumption, we can use the regression results of equation (7) to estimate a conditional $ATE(x_i) = \alpha_1 - \alpha_0 + (\beta_1 - \beta_0)x_i$. Moreover, we have $ATE(x_i) = ATT(x_i)$. The ATE is then derived by averaging $ATE(x_i)$ over all merger observations i and ATT is derived by averaging over the mergers with a termination fee. The two treatment effects are only identified if for every covariate vector x_i there are both mergers with and mergers without termination fee. This is referred to as the "overlap" assumption (See, for example, Imbens and Wooldridge, 2009).

[5] The literature also uses this term for the stricter assumption of conditional independence also known as ignorability of treatment. For our purpose, the weaker mean-independence assumption ($E(y_{gi} \mid x_i, t_i) = E(y_{gi} \mid x_i)$ for $g = 0, 1$) is sufficient.

We attempt to ensure that overlap is fulfilled by removing observations with extreme propensity scores and run regression (7) on such adjusted samples.

4.2 Sources of endogeneity and the bias of OLS estimates

The consistency of the OLS estimates hinges on the unconfoundedness assumption. Existing theories and empirical evidence on managerial merger motives, however, give reason to doubt that this assumption holds. In this section, we will discuss potential economic mechanisms that suggest unobservable variables could affect both the termination fee choice and price informativeness. The discussion attempts to accomplish two goals. First, we identify private benefits of control as a potential source of endogeneity and suggest appropriate econometric techniques that deal with this problem. Second, we show that an estimate of the OLS bias can tell us something about the mechanism through which private benefits of control influence price informativeness and the termination fee choice.

An obvious concern is that the acquirer management pursues an acquisition for its own private benefits. Such private benefits can come in various forms. Underdiversified managers might pursue an acquisition to diversify their personal wealth (Amihud and Lev, 1981). Since managing larger firms tends to bring along increased reputation and higher salary, managers might use the free cash flow to increase the size of the firm (Jensen, 1986). Managers might conduct acquisitions to improve their job security by either increasing the firm's dependence on the current management or by acquiring divisions at which they might be better managers (Shleifer and Vishny, 1989).[6] Empirical evidence consistent with managerial private benefits as acquisition motives are provided in several papers. See, for example, Morck, Shleifer, and Vishny (1990), Lang, Stulz, and Walkling (1991), Berkovitch and Narayanan (1993), Masulis, Wang, and Xie (2007), and Harford, Humphery-Jenner, and Powell (2012).

Whatever the source of private benefits, we expect bidder managers to be more willing to overpay and be more adamant in completing an acquisition if it provides such private benefits. This

[6] Earlier discussions of deviations from shareholder maximization can be found in Baumol (1959) and Williamson (1964), who suggest sales and expense maximization, respectively, as potential managerial objectives. Marris (1964) develops a theory of the firm based on sales growth maximization and Mueller (1969) builds on this theory to develop an explanation for conglomerate mergers.

also implies that they are more likely to ignore a negative stock price reaction to the announcement. If informed traders are aware of the private benefits, they are more likely to trade and price informativeness should increase. In sum, private benefits of control should make prices more informative.

Next we look at how private benefits of control might affect the termination fee choice. There is little theoretical work on why acquirer termination fees are included in merger agreements. Bates and Lemmon (2003) and Chen, Mahmudi, Virani, and Zhao (2021) provide arguments for why a termination fee is optimal for acquirer shareholders in certain circumstances. If the acquirer management pursues a merger for its own private benefits, we see at least three ways in which the termination fee choice could be affected. First, acquirer managers with private benefits of control will prefer a merger that allows their control to be retained. They are less likely to cancel such mergers and, consequently, will perceive the potential cost of agreeing to a termination fee as small. Second, managers might sign an expensive termination fee agreement to have an excuse vis-a-vis the shareholders for holding on to a deal that might later turn out to be bad for the shareholders. In both these cases, private benefits make the use of a termination fee more likely.

If, as above, private benefits are positively correlated with the use of termination fees and positively correlated with price informativeness, the OLS estimate of the effect of a termination fee in equation (7) will be biased upward. This happens because the termination fee also captures the effect of private benefits.

The third possibility produces the opposite bias. If the target management realizes that the acquirer management is unlikely to cancel the transaction because of significant private benefits, target management might decide to use its bargaining power to obtain other concessions than a termination fee. This argument implies that it is less likely to observe a termination fee when the acquirer management has private benefits of control. If the informed traders are aware of the managerial motive for the merger, they will trade on their information even if no termination fee is used. This behavior inflates the average price informativeness of the mergers without such a fee, lowering the OLS coefficient estimate of the termination fee variable.

In light of the discussion above and the existing empirical evidence on private benefits as a merger motive, it is plausible that there are significant correlations between the unobservable determinants of the termination fee choice and the error terms u_{0i} and u_{1i}. In Appendix B, we show how the OLS estimator is quantitatively affected by a non-zero correlation between these error terms. To deal with this endogeneity problem, we will use two different approaches: Instrumental variable (IV) estimation and an endogenous switching regression.

4.3 IV and Endogenous Switching Regressions

Our first solution to the endogeneity problem is to use an instrument that is correlated with termination fee use but not with the error terms u_{0i} and u_{1i} in equation (6). Bates and Lemmon (2003) and Officer (2003) document that acquirer termination fees became much more common after the *Paramount v. QVC Network* Delaware court ruling in February 1994. The court ruled that the stock lockup option granted by Paramount to a prospective acquirer Viacom was invalid, while the $100 million termination fee was ruled to be a reasonable amount to cover Viacom's expenses in connection with the transaction. Although this concerned a target termination fee, it has also eliminated the legal uncertainty surrounding the use of an acquirer termination fee. This can be seen in Panel B of Figure 1, where we see a sharp increase in the use acquirer termination fees starting in 1994.

We rely on the *Paramount v. QVC Network* court ruling as an exogenous shock to the usage of acquirer termination fees and define Paramount as an indicator variable that is equal to zero for the eight first years of the sample period, 1986-1993, and equal to one for the remaining years. The idea behind the instrument is that the court ruling acts as an exogenous shock that pushed some acquirers that otherwise would not have used a termination fee into using one—possibly because the uncertainty about the legality of a termination fee was resolved. We also interact the Paramount dummy with other variables to account for the possibility that the court ruling affected how acquirer, target, and deal characteristics influence the use of a termination fee.

To implement the instrumental variable approach, we need to make some simplifying assumptions to equation (6). Following Wooldridge (2010, section 21.4.1), we assume that there is no different unobservable effect of the termination fee choice on price informativeness ($u_{0i} = u_{1i} = u_i$)

and no differential impact of observable variables on price informativeness ($\beta_0 = \beta_1$). This brings us back to equation (3). The endogeneity concern is that private benefits contained in u_i are related to the termination fee dummy t_i. Instead of a linear probability model, we employ the procedure suggested in Windmeijer and Santos Silva (1997) and Wooldridge (2010, section 21.4). We start with a probit model for the termination fee choice that contains a vector of instruments, z_i, and the exogenous regressors of the price informative regression, x_i (see equation 8 below). The actual first stage regression (9) uses the fitted probabilities $\Phi(\hat{\theta}_0 + x_i\hat{\theta} + z_i\hat{\theta}_z)$ from the probit model (8) and the regressors from the price informativeness equation, x_i. The second-stage regression estimates the effect of using a termination fee, $(\alpha_1 - \alpha_0)$, where $\hat{t}_i$ are the predicted values from (9).

$$\text{stage 0:} \quad t_i = 1[\theta_0 + x_i\theta + z_i\theta_z + \epsilon_i \geq 0] \quad \text{(probit)} \tag{8}$$

$$\text{stage 1:} \quad t_i = \gamma_0 + x_i\gamma + \gamma_z\Phi(\hat{\theta}_0 + x_i\hat{\theta} + z_i\hat{\theta}_z) + v_i \tag{9}$$

$$\text{stage 2:} \quad y_i = \beta_0 + x_i\beta + (\alpha_1 - \alpha_0)\hat{t}_i + u_i \tag{10}$$

Note that the probit model for the termination fee use does not need to be correctly specified, as it would have to be for consistency if the predicted probabilities from the probit model were directly used in the second stage (see Angrist and Krueger, 2001).[7]

A major concern with the IV approach is that it consistently estimates the effect of the termination fee use on price informativeness only if $u_{0i} = u_{1i}$, allowing us to neglect the term $t_i(u_{1i} - u_{0i})$.[8] However, our discussion in Section 4.2 raises doubts about whether the nature of the impact of private benefits on price informativeness is unaffected by the presence of a termination fee. Private benefits are less likely to serve as a strong indicator of commitment if a merger already includes a sizable termination fee, implying $u_{0i} > u_{1i}$, all else equal. Neglecting this, as the IV approach does, would assign too much impact to private benefits and likely underestimate the effect of a termination fee. The ultimate bias depends also on the covariation with the other control variables, but

[7] What is needed is that the linear projection of t_i on x_i and $\Phi(\theta_0^* + x_i\theta^* + z_i\theta_z^*)$ actually depends on $\Phi(\theta_0^* + x_i\theta^* + z_i\theta_z^*)$, where θ_0^*, θ^*, and θ_z^* represent the plim of the ML estimator of the misspecified model (Wooldridge, 2010, p 940).

[8] See, e.g., Vella and Verbeek (1999) and Heckman (2010) for discussions of which restrictive assumptions one places implicitly on the economic mechanism that has generated the data if one uses the IV approach.

IV will not identify the treatment effect of a termination fee, $\alpha_1 - \alpha_0$, as we show in Appendix B.

To overcome the limitation of the IV approach, we employ an endogenous switching model where the termination fee choice is modeled explicitly.[9] It comes at the cost of having to impose a trivariate normality assumption for the error terms. The original model in equations (4) and (5) is augmented by:

$$t_i = 1[\theta_0 + x_i\theta + z_i\theta_z + \epsilon_i \geq 0] \tag{11}$$

$$(u_0, u_1, \epsilon)|x, z \sim \text{Multivariate Normal} \quad \text{with}$$

$$cov(u_0, u_1, \epsilon) = \begin{bmatrix} \sigma_0^2 & \sigma_{01} & \sigma_{0\epsilon} \\ \sigma_{01} & \sigma_1^2 & \sigma_{1\epsilon} \\ \sigma_{0\epsilon} & \sigma_{1\epsilon} & 1 \end{bmatrix}$$

With these assumptions, an explicit expression for $E[u_{0i} + t_i(u_{1i} - u_{0i})|x_i, t_i]$ exists. The resulting terms can be added to equation (6) and the parameters of interest can then be consistently estimated with OLS or maximum likelihood as the conditional expectation of y_i is linear in the parameters:

$$\begin{aligned} E[y_i|t_i, x_i, z_i] = \alpha_0 + x_i\beta_0 + (\alpha_1 - \alpha_0)t_i + (x_i - \mu_x)(\beta_1 - \beta_0) \\ -(1 - t_i)\sigma_{0\epsilon}\frac{\phi(q_i\theta)}{1 - \Phi(q_i\theta)} + t_i\sigma_{1\epsilon}\frac{\phi(q_i\theta)}{\Phi(q_i\theta)} \end{aligned} \tag{12}$$

where $q = (x, z)$ is the vector of outcome equation covariates and instruments, and $\phi(\cdot)$ and $\Phi(\cdot)$ are the normal density and cumulative distribution function. An additional advantage of the endogenous switching model is that we can test the different mechanisms presented in 4.2 because we obtain estimates of $\sigma_{0\epsilon}$ and $\sigma_{1\epsilon}$. A natural test of endogeneity of the termination fee choice is to test whether $\sigma_{0\epsilon}$ and $\sigma_{1\epsilon}$ are jointly zero.

[9] See section 21.4.2 in Wooldridge (2010). Applications of this method to finance can be found in Campa and Kedia (2002) and Fang (2005).

5 Results

5.1 Evidence from OLS on endogenous limits to arbitrage

We present the basic OLS regression of equation (7) as a benchmark and reference point for the later approaches. We start with testing Hypothesis 1a and 1b and regress price informativeness, defined in equation (2), on the Termination Fee Dummy or Termination Fee Size and the set of independent variables from Panel B of Table 2. In addition, we control for a trend in the average annual price informativeness across firms in the cross-section. The time-series pattern of this annual average resembles a quadratic trend until 2011. We deal with this by either including a quadratic trend plus year dummies for the years from 2012 to 2018 or by including year fixed effects. For all specifications we include industry fixed effects. Table 4 documents our findings.

The coefficient of the termination fee dummy is positive and statistically significant in both specifications (columns (1) and (3)). The point estimates of 0.113 (quadratic trend, specification (1)) and 0.123 (year fixed effects, specification (3)) represent the average increase in price informativeness if an acquirer termination fee is used. Given the specification in equation (7), the coefficient of the termination fee dummy can be interpreted as the average treatment effect (ATE) of using a termination fee on price informativeness.[10] The treatment effects are equivalent to 15% and 17% of the sample average price informativeness of 0.73 and around 10% of its standard deviation as documented in Table 2.

Our estimates are based on the assumption that for each merger with a termination fee, there exists a comparable counterfactual observation among the mergers without a termination fee. Lack of available counterfactuals, referred to as the overlap problem in section 4.1, are more likely when the probability of using a termination fee is close to zero or close to one. We therefore address this potential issue by trimming observations with a probability of using a termination in the tails of the distribution. When trimming observations below p and above $1-p$ for $p \in \{0.001, 0.01, 0.05, 0.1\}$ we estimate treatment effects between 0.101 and 0.124 with t-values between 2.48 and 2.93.[11] Since

[10] In this specification, the ATE is identical to the ATT. Panel A of Table A5 in Appendix A reports treatment effects and t-values.

[11] After trimming the sample, the sample size varies between 2,500 and 10,000 observations. These findings are not tabulated in a Table.

point estimates and statistical significance is close to what we report in columns (1) and (3) of Table 4, lack of overlap does not seem to cause any issues when estimating the effect of termination fees on price informativeness in our data.

In column (2) and (4), Termination Fee Size is used as the independent variable of interest. The coefficient is again positive and highly statistically significant. Thus, price informativeness is increasing in the size of the acquirer termination fee. These two sets of findings support our Hypotheses 1a and 1b.

The above initial strong support for hypotheses 1a and 1b is indicative of endogenous limits to arbitrage being present when no termination fee is used. However, the estimated coefficients are obtained from lumping together good and bad acquisitions and are akin to an average of the effects in those two samples. If endogenous limits to arbitrage are indeed the mechanism behind the effect of a termination fee on price informativeness, then we should observe a significant effect only among bad mergers (Hypothesis 2).

Ideally, we would test this hypothesis in a subsample of mergers where speculators have received a negative private signal about the merger quality. However, the nature of a private signal is that it is unobservable for all but the person receiving the signal. If there exist good proxies for signals about merger quality and the information is available for the acquirer when making the merger decision, the signal cannot be considered private.

Instead we focus on identifying mergers where the likelihood of a bad signal is higher than the likelihood of a good signal. To this end we rely on the acquirer abnormal announcement return. If the market reaction to the merger is negative on the announcement day, it is more likely that the merger is bad and speculators have more likely received a negative private signal about merger quality.

The variable Announcement AR for acquirer i on day t is defined as the difference between realized return and normal return on the day after the announcement.[12] We expect signals to

[12] Announcement returns are typically estimated using a two- or three-day window around the announcement. There are several reasons to follow this approach. One is that it captures the announcement returns even for deals that are announced after market close on the announcement date. Another is to capture the effect of initial trades on subsequent trading. Since we are interested in the initial signal received by the market on the announcement day and later measure the effect of subsequent trading, we use a one-day abnormal announcement return as a proxy for deal quality. We use the day after the announcement to capture the signal also for deals that are announced after

informed traders about the deal quality to have been weaker if the announcement return turned out to be close to zero and stronger if it turned out to be further away from zero. We therefore define mergers likely to have generated a bad private signal as deals with an abnormal announcement return below -0.02. Deals likely to have generated a good private signal are deals with an abnormal announcement return above 0.02.[13] We test Hypothesis 2 by running the previous regressions of price informativeness on the termination fee and control variables for each of the two subsamples separately.

Table 5 provides the results for the year fixed effects specification.[14] When splitting the sample using the abnormal announcement return as described above, the "Negative signal" subsample consist of 2,854 mergers and the "Positive signal" subsample has 2,919 mergers. We drop 6,746 observations with abnormal announcement returns between -0.02 and 0.02.

Under Hypothesis 2, we would expect that the use of a termination fee has a positive significant effect on price informativeness in the negative-signal subsample, while there should be no effect in the positive-signal subsample. This is exactly what we find. For the negative-signal subsample, the coefficient estimate for the termination fee dummy is 0.15 with a t-statistic of 2.33. The estimated coefficient for the positive-signal subsample is much smaller and not statistically different from zero.

We repeat the analysis with the termination fee size replacing the termination fee dummy and report results in panel B of Table 5. The point estimate for the subsample with a negative signal is positive and statistically significant at conventional levels. The point estimate in the subsample with a positive signal is less than one fifth of the estimate in the negative-signal subsample and statistically indistinguishable from zero.[15]

Taken together, these results provide strong support for the presence of endogenous limits to arbitrage and that a termination fee lifts these limits to arbitrage and produces more informative

market close. The model for computing normal returns is estimated over the period -294 through -43 relative to then announcement date, with a minimum of 126 observations and using the Fama-French three-factor model. The abnormal return is then trimmed at the 1st and 99th percentile.

[13] Results are qualitatively similar with cutoffs ranging from -0.025 through -0.005 and from 0.005 through 0.025. We report the results for the -0.01 and 0.01 cutoff combination in Table A2 in Appendix A.

[14] The results are very similar when we control for Trend and Trend squared.

[15] In Table A2 in Appendix A we show that results are qualitatively the same if we use a cutoff of -0.01 for the negative signal and 0.01 for the positive signal. This Table also shows that results do not change if we control for Trend and Trend squared.

prices. However, there is the risk that the results are affected positively or negatively by an omitted variable like managerial private benefits.

5.2 Evidence from IV on endogenous limits to arbitrage

In this section, we take into account that unobservable variables such as private benefits could influence both the choice of using an acquirer termination fee and price informativeness. We use the adjusted 2SLS procedure with the Paramount dummy and interaction terms of this dummy with deal, bidder, and target characteristics as instrumental variables (see section 4.3).

Table 6 presents the results of four different specifications for the initial probit model for the termination fee choice. The first two columns show the results of specifications without time-related variables. Column (3) shows the results when a quadratic trend is used and column (4) the results for a specification with year dummies. The termination fee use does not exhibit any particular time-series pattern (except for the sharp increase after the Paramount ruling which we exploit as IV). But since price informativeness does exhibit time-series patterns, we must also include these time-related variables in the probit model.

The four probit models show that the Paramount ruling had a strong and significant impact on the use of acquirer termination fees. After the Paramount ruling, a termination fee is considerably more likely. For the termination fee users, the Paramount ruling made fee usage seven times more likely. Columns (2) to (4) reveal that the ruling also changed the way acquirer, target, and deal characteristics influence the decision to use a termination fee. The effect of deal characteristics has increased considerably in magnitude. The impact of deal size and acquirer size has risen by 60 to 70%. We also find that more highly levered bidders are more likely to include a termination fee after the ruling. A potential explanation is that the target is concerned about financing problems leading to a deal breakdown. Paying in cash is associated with a lower propensity to use a termination fee. This finding could be related to both the financing situation of the acquirer and to the availability of free cash flow that the management might use for its own benefit, e.g., for empire building. We will explore this agency issue in more detail in the next section. The size and significance of the interaction terms justify using both the Paramount dummy and interaction terms between this dummy and deal characteristics as instruments.

Table 7 presents the results of the first stage regression (9). The main finding is that the predicted probability from the probit model ($\Phi(\hat{\theta}_0 + x_i\hat{\theta} + z_i\hat{\theta}_z)$) is a highly significant instrument for the termination fee dummy. For all four specifications, comparing the numbers in the row labeled "F-stat weak instr" with the numbers in the row labeled "5% critical value" at the bottom of the Table, it is clear that the F-statistic is much larger than the critical value for a weak instrument—rejecting the null hypothesis of a weak instrument.[16] This is not very surprising since the use of acquirer termination fees has become much more common after the Paramount ruling. When the predicted probability is added to the first stage regression, all other estimated coefficients are close to zero and statistically insignificant. This suggests that the probit model in Table 6 captures the main determinants of the termination fee choice.

Finally, the second-stage regression results using the Paramount instrument are reported in Table 8. Column (1) and column (2) represent second-stage results for models without time-trend variables. Column (3) reports the results for the quadratic trend model and column (4) the results for the model with year fixed effects. In all specifications, the termination fee dummy instrumented with the Paramount dummy and interaction terms is positively related to price informativeness, with point estimates between 0.33 and 0.45. These estimates are almost four times as large as the OLS estimates from Table 4 and amount to around 60% of the sample average and 40% of the standard deviation of our price informativeness measure. The finding that the IV estimate exceeds the OLS estimate is consistent with the economic mechanism from section 4.2 where target managers and informed traders consider certain managerial merger motives as a sign of commitment by the bidder. This induces the informed traders to trade on their information and the target managers to go after other provisions in the merger agreement than an acquirer termination fee.

Next we test the prediction that the inclusion of a termination fee should only impact price informativeness when speculators receive a bad signal about the merger (Hypothesis 2). We follow the same procedure as in Section 5.1 and define mergers likely to have generated a bad private signal as deals with an abnormal announcement return below -0.02. Deals likely to have generated

[16] The F-statistic is from the Montiel-Pflueger robust weak instrument test (Montiel Olea and Pflueger, 2013). The critical values are based on the asymptotic 2SLS-estimator bias at the 5% threshold.

a good private signal are deals with an abnormal announcement return above 0.02.

In Table 9, we report the results for the 2SLS specification with year fixed effects.[17] The first three columns under the headline "Negative signal" show the results of the augmented 2SLS procedure for bad mergers and the remaining three columns the results for good mergers. Again, the instruments have a strong effect on the use of a termination fee, with the Montiel-Pflueger F-statistics comfortably exceeding the 5%-critical value for a 2SLS estimator bias of at least 5%. The important finding is that the termination fee dummy has a highly significant effect on price informativeness in the bad merger sample while it is insignificant in the good merger sample. For the bad merger sample, the coefficient has further increased to 0.508.[18] This coefficient estimate represents both the ATE and the ATT on price informativeness of using a termination fee.[19] This is in line with the endogenous-limits to arbitrage theory, which states that informed traders would only be reluctant to trade on negative information.

Overall, the IV estimation results provide clear support for the presence of endogenous limits to arbitrage. However, we continue our analysis because our earlier discussion on how private benefits might impact price informativeness indicates that an assumption underlying the IV approach may not hold in our sample.

5.3 Evidence from the endogenous switching model

Our concern with the IV approach is that the impact of unobservable variables on price informativeness is assumed to be the same irrespective of whether a termination fee was used or not ($u_{1i} = u_{0i}$). This entails that the relation between the unobservable variables affecting the termination fee choice (ϵ_i) and those affecting price informativeness is assumed to be the same for both the termination fee sample and the no termination fee sample. In terms of covariances, this translates to $\sigma_{0\epsilon} = \sigma_{1\epsilon}$.[20] Our discussion of potential sources of endogeneity suggests that private benefits

[17]The results for the quadratic-trend specification can be found in Table A3 in Appendix A.

[18]If we use an announcement return of below -1% to create the negative-AR sample, then we find a coefficient of 0.507 for this subsample. The positive-AR sample, based on announcement returns above 1% yields an insignificant coefficient of 0.005.

[19]Panel B of Table A5 in Appendix A reports treatment effects and t-values.

[20]Vella and Verbeek (1999) discuss the differences in assumptions between the IV approach and the control function approach (endogenous switching regression in our setting) and the importance of inspecting the economic mechanism

likely have a stronger, more negative effect on price informativeness if no termination fee is used, i.e., $\sigma_{0\epsilon} < \sigma_{1\epsilon} \leq 0$. The endogenous switching model (4), (5), and (11) permits us to overcome this shortcoming of the IV approach because it allows for distinct covariances.

Table 10 contains the results for the endogenous switching model applied to the entire sample. We report the results for the specification with a quadratic trend and the specification with year fixed-effects. The column "Selection" contains the estimated coefficients for the probit model in equation (11) and the column "Outcome" the results of the price informativeness equation. The results for the two specifications of the probit model are similar to the results for the probit model of the IV approach reported in Table 6. The estimated effect of the termination fee on price informativeness is larger than the coefficient obtained via IV. For the quadratic time trend specification (year fixed effects specification), we obtain a coefficient of 0.531 (0.557), which compares to an IV-estimate of 0.426 (0.448). These estimates correspond to more than 70% of the average of price informativeness and to around 50% of its standard deviation.

In panel B of Table 10, we report the estimates of the covariances between the error terms in the termination fee choice equation and the error terms in the price informativeness equation, ($\sigma_{0\epsilon}$ and $\sigma_{1\epsilon}$). They provide three interesting insights. First, both $\sigma_{0\epsilon}$ and $\sigma_{1\epsilon}$ are negative. These results hint at the same economic mechanism behind the termination fee choice as the bias of the OLS coefficient: Acquisitions motivated by private benefits of the bidder management are less likely to involve a termination fee because private benefits and termination fees are substitutes in their role as commitment signals. This commitment is also noticed by informed traders, who are willing to trade on their information. Hence, unobservable private benefits create a negative correlation between the error terms, which produces downward biased OLS estimates (see equation (13) in Appendix B). Second, $\sigma_{0\epsilon}$ is almost twice as large in absolute terms as $\sigma_{1\epsilon}$ indicating that private benefits are not as important as a sign of commitment if a termination fee is already used. Nevertheless, private benefits seem to provide a signal of bidder commitment above and beyond the use of a termination fee, otherwise $\sigma_{1\epsilon}$ would be 0. This difference explains why IV estimates, which are based on the assumption of equal covariances, differ from the estimates of the endogenous switching model (see

giving rise to endogeneity in the first place.

equation (14) in Appendix B). Finally, we compute correlations, $\rho_i = \sigma_{i\epsilon}/\sigma_i$ for $i = 0, 1$, and test whether these are jointly zero. The Wald test of $\rho_{0\epsilon} = \rho_{1\epsilon} = 0$ strongly rejects the null hypothesis of zero correlations and, therefore, the absence of an endogeneity problem ($\chi^2_2 = 8.0$ with a p-value of 0.018 for the quadratic trend specification).

As in the two previous sections, we put the endogenous limits to arbitrage model to a tougher test and run the endogenous switching model on the high and low announcement return subsamples separately. We report the year-fixed effects specification in Table 11.[21] The results are strikingly clear. For acquisitions that experienced a negative announcement return (negative signal), the inclusion of a termination fee leads to an increase in price informativeness of 0.832, which amounts to 75% of its standard deviation and is larger than its average. Within the positive announcement return sample, the inclusion of a termination fee does not have a significant effect on price informativeness.

As is the case for both the OLS Termination Fee dummy coefficient from Section 5.1 and the corresponding IV coefficients from Section 5.2, the above Termination Fee coefficient of 0.832 represents the ATE on price informativeness of using a termination fee in the bad signal sub-sample. However, our endogenous switching regression specification implies that the Average Treatment effect on the Treated (ATT) is different than the ATE. The ATT in the full sample is 0.872 (z-value of 3.04). In the bad acquisition sample the ATT is 1.099 (z-value of 3.83) and in the good acquisition sample the ATT is -0.236 (z-value of -1.00). The ATT is larger than the ATE because the average treatment effect is larger for those firms that use a termination fee.[22]

Finally, we analyze the correlation structure of the error terms in the good and the bad acquisition subsamples. For the bad acquisition subsample, the findings for the covariances between the error terms in the termination fee choice equation and the error terms in the price informativeness equation are similar to the overall sample. However, in the positive abnormal return subsample, the correlation between the error terms is insignificant. Together with the insignificant termination fee dummy for this subsample, this strongly suggests that positively informed traders don't need

[21]The quadratic trend specification is found in Table A4 of Appendix A.
[22]Panel C of Table A5 in Appendix A reports treatment effects and z-values.

a commitment signal to be willing to trade on their information.[23] This is what we expect. The endogenous limits to arbitrage mechanism should only operate in the bad merger sample.

The results of the endogenous switching model strongly support the endogenous limits to arbitrage theory. The treatment effect of using a termination fee on price informativeness is statistically and economically significant. This is particularly true for the sub-sample of mergers where informed speculators are more likely to have received a signal indicating a low-quality merger. For mergers in this sub-sample, we also find support for our endogeneity concern that is related to the presence of private benefits of the bidder management. In the good acquisition sample, neither the termination fee nor the potential source of endogeneity matter for price informativeness. This findings are in line with the asymmetric nature of the endogenous limits to arbitrage mechanism.

6 Whose Trading Makes Prices Informative?

It is well known that merger arbitrageurs are trading in the acquirer around the merger announcement. Merger arbitrage involves buying the target's stock and shorting the stock of the acquirer in such a way that the short and long side perfectly match each other if and when the transaction closes. Merger arbitrageurs profit from a gap typically observed between the target stock price right after the announcement and the offered acquisition price. While they care about the merger success probability, they do not care about the quality of the merger per se. Regardless of the merger's quality, their strategy always involves shorting the bidder. Thus, their trading may or may not make prices informative in the same way the trading of informed traders does. For that reason, we check whether our earlier results hold in a subsample of mergers where merger arbitrageurs are absent. This will be the case if the target is not publicly listed.

Table 12 documents the relation between termination fee and price informativeness for the subsample of mergers with non-listed targets and for the subsample with listed targets. For the latter subsample, we also require that the payment was not made exclusively in cash. This ensures that any merger arbitrage trade, if it was made, always involved the bidder stock.

[23] This findings does not allow us to conclude that the termination fee choice is unaffected by the private benefits of the management in that subsample. It implies that such unobservables are not correlated with unobservables in the price informativeness equation and do not cause an endogeneity problem.

We present the estimates for the coefficient of the termination fee dummy using our three different estimation techniques: OLS, the instrumental variable approach, and the endogenous switching approach. For the non-listed sample, the OLS estimates are similar in magnitude to the original estimates but only marginally significant. The IV estimates and the estimates of the endogenous switching model for this sample are highly significant and larger than the original estimates in Tables 8 and 10. This suggests that the endogeneity bias caused by the bidder management pursuing private benefits is more severe in the subsample of non-listed targets. The estimates for the non-listed sample tend to be larger than the estimates for the listed sample, indicating that the termination fee has a larger impact on price informativeness if merger arbitrageurs are absent.

Overall, the evidence in Table 12 is qualitatively similar to the evidence documented in Section 5. Thus, the support we find in favor of our hypotheses is not driven by the trading of merger arbitrageurs.

7 Conclusion

The theory of endogenous limits to arbitrage, as specified in Edmans, Goldstein, and Jiang (2015), suggests that the possibility of feedback from speculator trading, via information content of prices and order flow, to managerial actions causes speculators to be more reluctant to trade on their information. This leads to endogenously inefficient markets.

In this paper, we investigate whether such endogenous limits are at work in financial markets and measure their impact on price informativeness. Our empirical setting is trading and information flow around merger announcements. Endogenous limits to arbitrage can arise in this context if informed speculators are reluctant to trade on private negative information relevant to the merger, because they fear that the acquirer will cancel the transaction upon learning about merger quality through observation of prices and order flow. Our empirical investigation relies on the idea that prices should be more informative if the feedback loop is broken because it is too costly to cancel the transaction.

An acquirer termination fee, whereby the bidder has agreed to pay a fee to the target if the deal is canceled, can be such a cost that might shut down the endogenous limits to arbitrage mechanism. Consequently, if endogenous limits to arbitrage exist, bidder stock prices should be more informative when a termination fee is used.

As a preliminary step, we show that mergers with an acquirer termination fee are canceled considerably less often than those without such a fee. We then estimate the effect of the use of a termination fee on price informativeness in three different ways: OLS, an instrumental variable approach, and using an endogenous switching model. The results of all three methods support the presence of endogenous limits to arbitrage. Price informativeness improves significantly if an acquirer termination fee is used. More importantly, we find that the asymmetry inherent in the endogenous limits to arbitrage mechanism holds in the data. Only traders who have received a negative signal about the merger quality should be concerned about revealing their private information and potentially causing the transaction to be cancelled. Using positive and negative announcement day returns as a proxy for the likelihood of private positive or negative signals, we find a significant impact of termination fee use on price informativeness only in the subsample of negative signals.

Our results suggest that it is important to control for endogeneity and omitted variables. The inclusion of an acquirer termination fee in the merger agreement is a choice determined in the negotiations between the acquirer and the target, and merger motives of the bidder management are likely to play a crucial role. A large literature identifies managerial private benefits and empire building as merger motives. In such cases, the bidder management is more likely to complete the transaction even if the stock market reaction is negative, which should increase the willingness to trade on private information and thus enhance price informativeness. Such private benefits are also likely to affect the bargaining process leading to the merger agreement. Ex ante, we can think of different ways private benefits could affect the bargaining outcome, with some economic mechanisms making the inclusion of an acquirer termination fee more likely while others making it less likely. Since different mechanisms produce different biases in the OLS regression, the comparison between the OLS and the IV results allows us to shed some light on the mechanism at work.

In the landmark case *Paramount v. QVC Network*, the court ruled in favor of including a termination fee in merger agreements. This decision led to a strong and lasting increase in the usage of acquirer termination fees and altered how bidder, target, and deal characteristics affect the use of these fees. We employ a dummy variable equal to one after this ruling, along with interaction terms of this dummy with other characteristics as powerful and plausibly exogenous instrumental variables in a standard IV approach and an endogenous switching model. We consider the latter method better suited to the economic mechanisms suspected of causing the endogeneity problem. The resulting coefficient estimates are much larger than the OLS estimates. The increase suggests that the target management understands the opportunistic merger motives of the bidder management, is less concerned with the transaction failing, and uses its bargaining power to secure other concessions than a termination fee in the bargaining leading to the merger agreement.

Finally, we investigate whether the difference in price informativeness is due to the trading behavior of merger arbitrageurs, who are concerned with the likelihood of deal completion rather than the quality of the merger itself. We find that the effect is present and even more pronounced in the sample of mergers with non-listed targets, where the merger arbitrage trade is not feasible.

Overall, our results suggest a sizable impact of endogenous limits to arbitrage. The increase in price informativeness brought about by using a termination fee amounts to more than average price informativeness in the data and 74% of its standard deviation. In the sample of non-listed firms it even amounts to an increase that corresponds to 175% of its mean and 116% of its standard deviation.

The implication of our findings is an inevitable welfare loss: If the acquirer in a bad merger has not tied its hands through a termination fee, endogenous limits to arbitrage make prices less informative and an inefficient merger is completed. If the acquirer has tied its hands, prices are informative but because of the commitment an inefficient merger is completed.

References

Aktas, Nihat, Eric de Bodt, Fany Declerck, and Hervè Van Oppens, 2007, The pin anomaly around M&A announcements, *Journal of Financial Markets* 10, 169–191.

Amihud, Yakov, and Baruch Lev, 1981, Risk reduction as a managerial motive for conglomerate mergers, *The Bell Journal of Economics* 12, 605–617.

Angrist, Joshua D., and Alan B. Krueger, 2001, Instrumental variables and the search for identification: From supply and demand to natural experiments, *The Journal of Economic Perspectives* 15, 69–85.

Baker, Malcolm, Jeremy C. Stein, and Jeffrey Wurgler, 2003, When does the market matter? stock prices and the investment of equity-dependent firms, *The Quarterly Journal of Economics* 118, 969–1005.

Bates, Thomas W., and Michael L. Lemmon, 2003, Breaking up is hard to do? an analysis of termination fee provisions and merger outcomes, *Journal of Financial Economics* 69, 469–504.

Baumol, W.J., 1959, *Business behavior, value and growth* (Macmillan).

Berkovitch, Elazar, and M. P. Narayanan, 1993, Motives for takeovers: An empirical investigation, *The Journal of Financial and Quantitative Analysis* 28, 347–362.

Boleslavsky, Raphael, David Kelly, and Curtis R. Taylor, 2017, Selloffs, bailouts, and feedback: Can asset markets inform policy?, *Journal of Economic Theory* 169, 294–343.

Bond, Philip, Itay Goldstein, and Edward Simpson Prescott, 2010, Market-based corrective actions, *Review of Financial Studies* 23, 781–820.

Campa, Jose Manuel, and Simi Kedia, 2002, Explaining the diversification discount, *The Journal of Finance* 57, 1731–1762.

Chen, Qi, Itay Goldstein, and Wei Jiang, 2007, Price informativeness and investment sensitivity to stock price, *Review of Financial Studies* 20, 619–650.

Chen, Zhiyao, Hamed Mahmudi, Aazam Virani, and Xiaofei Zhao, 2021, Why are bidder termination provisions included in takeovers?, *Journal of Financial and Quantitative Analysis* forthcoming.

Collin-Dufresne, Pierreand Fos, Vyacheslav, 2015, Do prices reveal the presence of informed trading?, *The Journal of Finance* 70, 1555–1582.

Dow, James, Itay Goldstein, and Alexander Guembel, 2017, Incentives for information production in markets where prices affect real investment, *Journal of the European Economic Association* 15, 877–909.

Dow, James, and Gary Gorton, 1997, Stock market efficiency and economic efficiency: Is there a connection?, *Journal of Finance* 52, 1087–1129.

Durnev, Artyom, Randall Morck, Bernard Yeung, and Paul Zarowin, 2003, Does greater firm-specific return variation mean more or less informed stock pricing?, *Journal of Accounting Research* 41, 797–836.

Easley, David, Soeren Hvidkjaer, and Maureen O'Hara, 2002, Is information risk a determinant of asset returns?, *Journal of Finance* 57, 2185–2221.

Edmans, Alex, Itay Goldstein, and Wei Jiang, 2011, The real effects of financial markets: The impact of prices on takeovers, *Journal of Finance* 67, 933–971.

———, 2015, Feedback effects, asymmetric trading, and the limits to arbitrage, *American Economic Review* 105, 3766–3797.

Fang, Lily Hua, 2005, Investment bank reputation and the price and quality of underwriting services, *The Journal of Finance* 60, pp. 2729–2761.

Ferreira, Daniel, Miguel A. Ferreira, and Clara C. Raposo, 2011, Board structure and price informativeness, *Journal of Financial Economics* 99, 523–545.

Gorton, Gary B., Lixin Huang, and Qiang Kang, 2017, The limitations of stock market efficiency: Price informativeness and ceo turnover, *Review of Finance* 21, 153–200.

Hansen, B., 2022, *Econometrics* (Princeton University Press).

Harford, Jarrad, Mark Humphery-Jenner, and Ronan Powell, 2012, The sources of value destruction in acquisitions by entrenched managers, *Journal of Financial Economics* 106, 247–261.

Heckman, James, and Salvador Navarro-Lozano, 2004, Using matching, instrumental variables, and control functions to estimate economic choice models, *The Review of Economics and Statistics* 86, 30–57.

Heckman, James J., 2010, Building bridges between structural and program evaluation approaches to evaluating policy, *Journal of Economic Literature* 48, 356–398.

Imbens, Guido W., and Jeffrey M. Wooldridge, 2009, Recent developments in the econometrics of program evaluation, *Journal of Economic Literature* 47, 5–86.

Jensen, Michael C., 1986, Agency costs of free cash flow, corporate finance, and takeovers, *The American Economic Review* 76, 323–329.

Kau, James B., James S. Linck, and Paul H. Rubin, 2008, Do managers listen to the market?, *Journal of Corporate Finance* 14, 347–362.

Lang, Larry H.P., RenéM. Stulz, and Ralph A. Walkling, 1991, A test of the free cash flow hypothesis: The case of bidder returns, *Journal of Financial Economics* 29, 315–335.

Luo, Yuanzhi, 2005, Do insiders learn from outsiders? evidence from mergers and acquisitions, *Journal of Finance* 60, 1951–1982.

Maddala, GS, 1983, *Limited-dependent and qualitative variables in econometrics* . , vol. 149 (Cambridge University Press).

Marris, R., 1964, *The Economic Theory of Managerial Capitalism* (Free Press of Glencoe).

Masulis, RONALD W., CONG Wang, and FEI Xie, 2007, Corporate governance and acquirer returns, *The Journal of Finance* 62, 1851–1889.

Montiel Olea, J. L., and C. E. Pflueger, 2013, A robust test for weak instruments, *Journal of Business and Economic Statistics* 31, 358–369.

Morck, Randall, Andrei Shleifer, and Robert W. Vishny, 1990, Do managerial objectives drive bad acquisitions?, *The Journal of Finance* 45, 31–48.

Morck, Randall, Bernard Yeung, and Wayne Yu, 2000, The information content of stock markets: why do emerging markets have synchronous stock price movements?, *Journal of Financial Economics* 58, 215–260.

Mueller, Dennis C., 1969, A Theory of Conglomerate Mergers, *The Quarterly Journal of Economics* 83, 643–659.

Officer, Micah S., 2003, Termination fees in mergers and acquisitions, *Journal of Financial Economics* 69, 431–467.

Roll, Richard, 1988, R^2, *Journal of Finance* 43, 541–566.

Shleifer, Andrei, and Robert W. Vishny, 1989, Management entrenchment: The case of manager-specific investments, *Journal of Financial Economics* 25, 123–139.

Subrahmanyam, Avanidhar, and Sheridan Titman, 2001, Feedback from stock prices to cash flows, *Journal of Finance* 56, 2389–2413.

Veldkamp, Laura L., 2006, Information Markets and the Comovement of Asset Prices, *The Review of Economic Studies* 73, 823–845.

Vella, Francis, and Marno Verbeek, 1999, Estimating and interpreting models with endogenous treatment effects, *Journal of Business & Economic Statistics* 17, 473–478.

Williamson, Oliver Eaton, 1964, The economics of discretionary behavior: Managerial objectives in a theory of the firm, *(No Title)*.

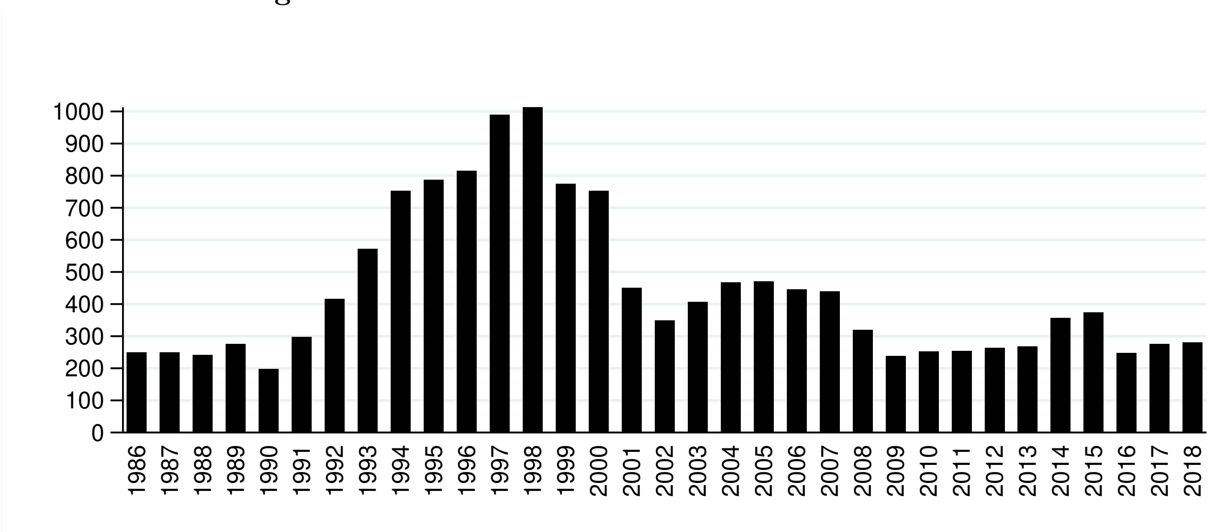
Windmeijer, F. A. G., and J. M. C. Santos Silva, 1997, Endogeneity in count data models: An application to demand for health care, *Journal of Applied Econometrics* 12, 281–294.

Wooldridge, Jeffrey M., 2010, *Econometric Analysis of Cross Section and Panel Data* . chap. 21 (The MIT Press).

Figure 1
Number of mergers and number of mergers where the acquirer has signed a termination agreement, 1986–2018

The sample consists of 14,545 mergers announced between 1986 and 2018. There are 1,327 mergers with an acquirer termination agreement. The merger data is from Refinitiv SDC Platinum (formerly Thomson Reuters SDC Platinum). The figure reflects the sample selection described in Section 3.1 and documented in Table 1.

A. Number of mergers



B. Number of mergers with acquirer termination agreement

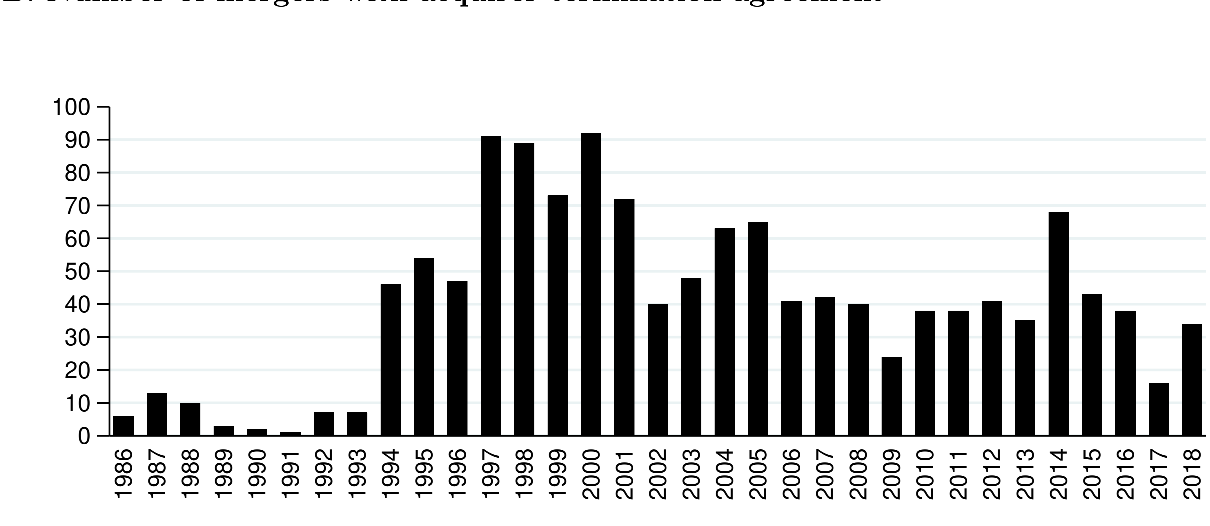


Table 1
Sample selection

The merger data is from Refinitiv SDC Platinum (formerly Thomson Reuters SDC Platinum). PERMNO is the permanent security identifier for security information provided by the Center for Research in Security Prices (CRSP). GVKEY is a permanent company identifier for data provided by Compustat. Relative Deal Size is transaction value divided by bidder market capitalization 42 days prior to the deal announcement date.

Sample selection filters	Num. obs.
Databases: US and Non-US deals + Date announced: 1.1.1986–31.12.2018 + Acquirer Nation: US + Acquirer Public Status: P + Deal value: Where information is available + Type of Deal: Merger	22,254
Match Bidder to PERMNO in CRSP and GVKEY in Compustat	18,362
CRSP Bidder share codes 10 or 11 + Deal ann. date within time-series of CRSP daily returns	17,550
Keeping only the first announcement in a bidding contest	17,454
Non-missing bidder market capitalization and total assets prior to announcement date	16,258
Deal Value divided by bidder Market Capitalization at $t = -42 > 1\%$	14,817
Dropping utilities	14,545
Exogenous variables (covariates) available	12,159

Table 2
Descriptive statistics

The sample consists of 14,545 mergers announced between 1986 and 2018. When the sample is required to have non-missing covariates, the sample size is 12,519. The merger data is from Refinitiv SDC Platinum (formerly Thomson Reuters SDC Platinum). The column named "Diff." reports the differences in means between samples with Termination Fee D equal to zero and Termination Fee D equal to one. The t-value in the last column is the difference in means divided by the standard error of the difference. All variables are defined in Appendix Table A1.

	Full sample				Termination Fee D = 0		Termination Fee D = 1			
Variable	N	Mean	Median	Std.	N	Mean	N	Mean	Diff.	t-val.
A. Termination fee and dependent variables										
Term fee D	12519	0.07	0.00	0.25						
Term Fee Size	12519	0.00	0.00	0.01	11691	0.00	828	0.03	-0.03	-113.2
Price Informativeness	12519	0.73	0.72	1.12	11691	0.74	828	0.69	0.04	1.04
Withdraw	12159	0.09	0.00	0.29	11336	0.09	823	0.09	0.00	0.31
B. Covariates										
Inst. Ownership	12519	0.46	0.45	0.30	11691	0.45	828	0.52	-0.07	-6.66
Breadth of Ownership	12519	0.06	0.03	0.08	11691	0.06	828	0.06	0.00	-1.11
Ln Deal Value	12519	4.39	4.23	2.00	11691	4.27	828	6.16	-1.89	-27.09
Cashonly	12519	0.29	0.00	0.46	11691	0.30	828	0.15	0.15	9.43
Unsolicited	12519	0.04	0.00	0.18	11691	0.04	828	0.03	0.01	1.04
Target is Public	12519	0.46	0.00	0.50	11691	0.44	828	0.80	-0.36	-20.61
Ln Market Cap	12519	6.36	6.36	2.03	11691	6.34	828	6.66	-0.32	-4.37
Market to book	12519	3.40	2.22	3.71	11691	3.41	828	3.24	0.17	1.29
Leverage	12519	0.22	0.16	0.22	11691	0.22	828	0.22	0.00	0.21
Past one-year CAR	12519	-0.01	-0.01	0.65	11691	-0.01	828	-0.06	0.06	2.37
Volatility	12519	0.48	0.39	0.29	11691	0.48	828	0.48	0.00	-0.16
Lagged price inf.	12519	1.21	1.19	1.22	11691	1.22	828	1.00	0.23	5.22
Normal Volume	12519	11.67	11.72	1.99	11691	11.63	828	12.25	-0.62	-8.73
Runup	12519	0.00	0.00	0.19	11691	0.00	828	0.02	-0.02	-3.11
Announcement AR	12519	0.00	0.00	0.05	11691	0.00	828	-0.01	0.01	5.96
Same Industry	12519	0.77	1.00	0.42	11691	0.77	828	0.81	-0.04	-2.94

Table 3
The probability of canceling a transaction

The sample consists of 14,545 mergers announced between 1986 and 2018. When the sample is required to have non-missing covariates, the sample size is 12,519. The merger data is from Refinitiv SDC Platinum (formerly Thomson Reuters SDC Platinum). The dependent variable is Withdrawn: A dummy variable equal to one if SDC status is "Withdrawn" and equal to zero if SDC status is "Completed." The regressors "1st quartile" to "4th quartile" are dummy variables equal to one if the termination fee size lies in the respective quartiles. The coefficients are estimated using Probit and standard errors are robust to heteroskedasticity. Parentheses contain z-values. All variables are defined in Appendix Table A1.

	Dependent variable: Withdrawn		
Variable	(1)	(2)	(3)
1st quartile	0.048 (0.38)		
2nd quartile	−0.404 (−2.81)		
3rd quartile	−0.349 (−2.77)		
4th quartile	−0.165 (−1.46)		
Term fee D		−0.296 (−3.86)	−0.418 (−4.05)
Term fee size			3.952 (1.88)
Inst. Ownership	0.106 (1.18)	0.108 (1.20)	0.106 (1.18)
Breadth of Ownership	0.640 (1.68)	0.640 (1.68)	0.631 (1.66)
Ln Deal Value	0.170 (10.03)	0.172 (10.19)	0.170 (10.07)
Cashonly	−0.163 (−3.45)	−0.164 (−3.47)	−0.167 (−3.54)
Unsolicited	1.685 (22.76)	1.681 (22.72)	1.683 (22.73)
Target is Public	0.592 (12.96)	0.592 (13.01)	0.593 (13.03)
Ln Firm Value	−0.274 (−9.85)	−0.275 (−9.92)	−0.273 (−9.82)
Market to book	0.004 (0.68)	0.004 (0.68)	0.004 (0.70)
Leverage	0.491 (4.54)	0.497 (4.61)	0.486 (4.49)
Past one-year CAR	−0.030 (−1.08)	−0.030 (−1.10)	−0.029 (−1.05)
Volatility	0.128 (1.33)	0.129 (1.34)	0.124 (1.29)
Lagged price inf.	0.071 (3.82)	0.071 (3.82)	0.071 (3.85)
Normal volume	0.065 (3.32)	0.065 (3.33)	0.065 (3.33)
Runup	−0.295 (−3.02)	−0.294 (−3.02)	−0.296 (−3.04)
Announcement AR	−0.268 (−0.71)	−0.264 (−0.70)	−0.251 (−0.66)
Same Industry	−0.106 (−2.39)	−0.105 (−2.38)	−0.106 (−2.41)
Constant	−1.284 (−5.81)	−1.287 (−5.83)	−1.290 (−5.85)
Year fixed effects	Yes	Yes	Yes
Industry fixed effects	Yes	Yes	Yes
Pseudo R^2	0.235	0.235	0.236
Number of observations	12,159	12,159	12,159

Table 4
Price informativeness and acquirer termination fees

The sample consists of 14,545 mergers announced between 1986 and 2018. When the sample is required to have non-missing covariates, the sample size is 12,519. The merger data is from Refinitiv SDC Platinum (formerly Thomson Reuters SDC Platinum). Coefficients are estimated using OLS and standard errors are robust to heteroskedasticity. T-statistics in parentheses. All variables are defined in Appendix Table A1.

	Dependent variable: Price informativeness			
Variable	(1)	(2)	(3)	(4)
Termination Fee D	0.113 (3.02)		0.123 (3.30)	
Termination Fee Size		4.173 (4.74)		4.306 (4.85)
Inst. Ownership	−0.214 (−5.18)	−0.214 (−5.18)	−0.230 (−5.55)	−0.230 (−5.55)
Breadth of Ownership	0.493 (2.76)	0.494 (2.76)	0.423 (2.39)	0.423 (2.39)
Ln Deal Value	0.034 (4.32)	0.032 (4.15)	0.036 (4.61)	0.035 (4.47)
Cashonly	0.029 (1.37)	0.028 (1.33)	0.038 (1.81)	0.036 (1.75)
Unsolicited	0.066 (1.43)	0.068 (1.47)	0.068 (1.48)	0.069 (1.51)
Target is Public	−0.045 (−2.25)	−0.044 (−2.19)	−0.043 (−2.14)	−0.041 (−2.06)
Ln Market Cap	−0.132 (−10.2)	−0.129 (−10.1)	−0.130 (−9.84)	−0.128 (−9.72)
Market to Book	−0.004 (−1.59)	−0.004 (−1.57)	−0.004 (−1.64)	−0.004 (−1.61)
Leverage	0.217 (4.18)	0.207 (3.99)	0.223 (4.29)	0.213 (4.09)
Past one-year CAR	−0.048 (−3.53)	−0.047 (−3.48)	−0.053 (−3.90)	−0.052 (−3.84)
Volatility	−0.126 (−2.63)	−0.129 (−2.69)	−0.123 (−2.34)	−0.126 (−2.40)
Lagged price inf.	0.195 (21.4)	0.195 (21.5)	0.190 (20.5)	0.190 (20.5)
Normal Volume	−0.043 (−4.47)	−0.044 (−4.52)	−0.043 (−4.42)	−0.044 (−4.47)
Runup	−0.199 (−4.09)	−0.200 (−4.13)	−0.209 (−4.32)	−0.211 (−4.36)
Announcement AR	0.203 (1.06)	0.201 (1.06)	0.136 (0.71)	0.133 (0.70)
Same Industry	−0.053 (−2.39)	−0.053 (−2.39)	−0.054 (−2.45)	−0.054 (−2.44)
Constant	1.936 (20.3)	1.939 (20.4)	1.884 (17.1)	1.886 (17.1)
Trend	0.047 (8.18)	0.047 (8.19)		
Trend squared	−0.003 (−12.9)	−0.003 (−12.9)		
Year fixed effects	2012-2018	2012-2018	Yes	Yes
Industry fixed effects	Yes	Yes	Yes	Yes
R^2	0.279	0.280	0.289	0.289
Number of observations	12,519	12,519	12,519	12,519

Table 5
Price informativeness and acquirer termination fees using subsamples selected on acquirer announcement day abnormal returns below −0.02 and above 0.02

The sample consists of 14,545 mergers announced between 1986 and 2018. When the sample is required to have non-missing covariates, the sample size is 12,519. The merger data is from Refinitiv SDC Platinum (formerly Thomson Reuters SDC Platinum). The column named "Negative signal" contains all observations with acquirer Announcement AR below −0.02. The column named "Positive signal" contains all observations with acquirer Announcement AR above 0.02. Coefficients are estimated using OLS and standard errors are robust to heteroskedasticity. T-statistics in parentheses. All variables are defined in Appendix Table A1.

	Dependent variable: Price informativeness	
	Negative signal	Positive signal
A. Termination Fee Dummy		
Term fee D	0.151 (2.33)	0.067 (0.91)
Inst. Ownership	−0.271 (−3.14)	−0.185 (−2.16)
Breadth of Ownership	0.425 (1.26)	0.620 (1.39)
Ln Deal Value	0.019 (1.09)	−0.003 (−0.18)
Cashonly	0.036 (0.77)	0.102 (2.22)
Unsolicited	0.095 (1.11)	0.060 (0.51)
Target is Public	−0.092 (−2.03)	0.025 (0.57)
Ln Market Cap	−0.123 (−4.69)	−0.100 (−3.88)
Market to Book	−0.004 (−0.79)	−0.009 (−1.72)
Leverage	0.216 (2.02)	0.140 (1.21)
Past one-year CAR	−0.050 (−2.04)	−0.052 (−2.11)
Volatility	−0.182 (−1.96)	−0.049 (−0.48)
Lagged price inf.	0.149 (7.67)	0.129 (6.53)
Normal Volume	−0.041 (−2.14)	−0.067 (−3.59)
Runup	−0.091 (−1.01)	−0.296 (−3.38)
Announcement AR	−0.580 (−1.01)	0.648 (1.39)
Same Industry	−0.001 (−0.02)	−0.094 (−2.18)
Constant	2.020 (7.69)	2.451 (8.89)
Year fixed effects	Yes	Yes
Industry fixed effects	Yes	Yes
R^2	0.269	0.257
Number of obs.	2,854	2,919
B. Termination Fee Size		
Termination Fee Size	5.442 (3.50)	0.958 (0.62)
Other Control var.	Yes	Yes
Year fixed effects	Yes	Yes
Industry fixed effects	Yes	Yes
R^2	0.271	0.257
Number of obs.	2,854	2,919

Table 6
Probit Model for termination fee choice

The sample consists of 14,545 mergers announced between 1986 and 2018. When the sample is required to have non-missing covariates, the sample size is 12,519. The merger data is from Refinitiv SDC Platinum (formerly Thomson Reuters SDC Platinum). The dependent variable is the termination fee dummy. The coefficients are estimated using Probit and standard errors are robust to heteroskedasticity. Parentheses contain z-values. All variables are defined in Appendix Table A1.

	(1)	(2)	(3)	(4)
Param dummy	0.946 (8.83)	1.377 (5.22)	0.902 (3.63)	1.544 (3.62)
Param × Deal size		0.287 (4.31)	0.260 (3.79)	0.286 (4.13)
Param × Cashonly		−1.118 (−6.10)	−1.304 (−6.74)	−1.183 (−5.89)
Param × Bidder size		−0.279 (−4.57)	−0.264 (−4.32)	−0.274 (−4.54)
Param × Bidder lev.		0.746 (1.83)	0.661 (1.58)	0.701 (1.71)
Param × Runup		−0.573 (−2.24)	−0.512 (−2.06)	−0.646 (−2.36)
Inst. Ownership	0.104 (1.10)	0.112 (1.17)	0.085 (0.88)	0.103 (1.06)
Breadth of Ownership	−1.043 (−2.40)	−1.106 (−2.51)	−0.357 (−0.81)	−0.291 (−0.65)
Ln Deal Value	0.560 (26.21)	0.298 (4.67)	0.323 (4.89)	0.305 (4.60)
Cashonly	−0.305 (−5.43)	0.723 (4.13)	0.833 (4.50)	0.711 (3.68)
Unsolicited	−0.855 (−6.99)	−0.838 (−6.76)	−0.870 (−6.94)	−0.893 (−7.10)
Target is Public	0.607 (11.57)	0.599 (11.17)	0.648 (11.70)	0.645 (11.54)
Ln Market Cap	−0.494 (−15.48)	−0.239 (−3.74)	−0.233 (−3.62)	−0.248 (−3.84)
Market to Book	−0.004 (−0.64)	−0.004 (−0.53)	0.001 (0.16)	0.002 (0.25)
Leverage	0.543 (4.18)	−0.160 (−0.40)	−0.049 (−0.12)	−0.082 (−0.20)
Past one-year CAR	−0.047 (−1.43)	−0.047 (−1.42)	−0.056 (−1.68)	−0.042 (−1.27)
Volatility	−0.140 (−1.21)	−0.154 (−1.29)	−0.021 (−0.18)	−0.149 (−1.12)
Lagged price inf.	−0.047 (−2.15)	−0.046 (−2.11)	0.001 (0.05)	0.006 (0.24)
Normal Volume	0.057 (2.49)	0.059 (2.59)	0.009 (0.38)	0.022 (0.91)
Runup	−0.121 (−1.07)	0.419 (1.85)	0.361 (1.67)	0.493 (2.01)
Announcement AR	−0.130 (−0.32)	−0.026 (−0.06)	−0.199 (−0.48)	−0.112 (−0.27)
Same Industry	0.085 (1.55)	0.091 (1.66)	0.097 (1.73)	0.089 (1.58)
Trend			0.076 (2.55)	
Trend squared			−0.001 (−1.20)	
constant	−3.162 (−12.14)	−3.602 (−10.14)	−3.798 (−9.94)	−3.546 (−7.66)
Year FE	No	No	2012-2018	Yes
Industry FE	Yes	Yes	Yes	Yes
Pseudo R^2	0.297	0.306	0.321	0.324
N	12,519	12,519	12,519	12,274

Table 7
Price informativeness using the Paramount dummy and interaction terms as instruments for acquirer termination fee: First stage regression

The sample consists of 14,545 mergers announced between 1986 and 2018. When the sample is required to have non-missing covariates, the sample size is 12,519. The merger data is from Refinitiv SDC Platinum (formerly Thomson Reuters SDC Platinum). The dependent variable is the termination fee dummy. $\Phi(\hat{\theta}_0 + x_i\hat{\theta} + z_i\hat{\theta}_z)$ is the predicted probability of using a termination fee using the estimates of the respective probit models in Table 6. The coefficients are estimated using 2SLS and standard errors are robust to heteroskedasticity. Parentheses contain t-values. The row "5% critical value" contains the 5% critical values for a relative bias (IV bias divide by worst case OLS bias) of 5%. The weak IV tests are performed using the Stata function weakivtest. All variables are defined in Appendix Table A1.

	(1)	(2)	(3)	(4)
$\Phi(\hat{\theta}_0 + x_i\hat{\theta} + z_i\hat{\theta}_z)$	1.067 (20.27)	1.020 (18.43)	1.024 (20.28)	1.031 (20.85)
Param dummy	−0.002 (−0.51)	−0.007 (−0.60)	−0.007 (−0.67)	−0.014 (−0.53)
Param × Deal size		−0.001 (−0.37)	−0.002 (−0.48)	−0.002 (−0.56)
Param × Cashonly		−0.001 (−0.08)	0.000 (−0.01)	0.000 (0.01)
Param × Bidder size		0.002 (0.66)	0.002 (0.73)	0.002 (0.83)
Param × Bidder lev.		−0.009 (−0.54)	−0.008 (−0.49)	−0.007 (−0.41)
Param × Runup		0.000 (0.02)	0.000 (−0.02)	0.000 (0.01)
Inst. Ownership	0.002 (0.20)	0.003 (0.26)	0.004 (0.39)	0.004 (0.41)
Breadth of Ownership	0.011 (0.27)	0.003 (0.06)	0.003 (0.08)	0.001 (0.03)
Ln Deal Value	−0.004 (−1.60)	0.000 (−0.12)	0.000 (−0.14)	0.000 (−0.19)
Cashonly	−0.001 (−0.31)	−0.001 (−0.12)	−0.001 (−0.10)	−0.001 (−0.08)
Unsolicited	0.004 (0.31)	−0.002 (−0.20)	−0.001 (−0.10)	0.000 (−0.04)
Target is Public	−0.004 (−0.74)	−0.001 (−0.11)	0.000 (−0.04)	0.000 (−0.04)
Ln Market Cap	0.003 (1.05)	0.000 (−0.06)	0.000 (−0.07)	0.000 (−0.01)
Market to Book	0.000 (0.12)	0.000 (0.03)	0.000 (0.02)	0.000 (−0.03)
Leverage	−0.007 (−0.58)	0.002 (0.19)	0.002 (0.13)	0.001 (0.05)
Past one-year CAR	0.000 (0.14)	0.000 (0.05)	0.000 (0.13)	0.000 (0.09)
Volatility	0.000 (0.04)	0.000 (−0.01)	0.000 (0.02)	0.001 (0.07)
Lagged price inf.	0.000 (0.07)	0.000 (0.02)	0.000 (0.05)	0.000 (0.08)
Normal Volume	0.000 (0.12)	0.000 (0.15)	0.000 (0.10)	0.000 (0.09)
Runup	−0.002 (−0.22)	−0.002 (−0.25)	−0.002 (−0.28)	−0.003 (−0.27)
Announcement AR	0.007 (0.12)	0.002 (0.03)	0.002 (0.04)	0.006 (0.09)
Same Industry	−0.001 (−0.15)	0.000 (−0.06)	0.000 (0.00)	0.000 (−0.08)
Trend			0.000 (0.21)	
Trend squared			0.000 (−0.23)	
constant	−0.008 (−0.38)	−0.004 (−0.16)	−0.004 (−0.19)	−0.003 (−0.11)
Year FE	No	No	2012-2018	Yes
Industry FE	Yes	Yes	Yes	Yes
R^2	0.208	0.212	0.227	0.233
F-stat weak instr	369.6	231.7	256.1	285.2
5% critical value	28.210	31.23	30.30	30.57
N	12519	12519	12519	12274

Table 8
Price informativeness using the Paramount dummy and interaction terms as instruments for acquirer termination fee: Second stage regression

The sample consists of 14,545 mergers announced between 1986 and 2018. When the sample is required to have non-missing covariates, the sample size is 12,519. The merger data is from Refinitiv SDC Platinum (formerly Thomson Reuters SDC Platinum). The dependent variable is price informativeness. The coefficients are estimated using 2SLS and standard errors are robust to heteroskedasticity. Parentheses in the columns contain z-values.. All variables are defined in Appendix Table A1.

	(1)	(2)	(3)	(4)
Term fee D	0.329 (2.61)	0.322 (2.60)	0.426 (3.61)	0.448 (3.86)
Inst. Ownership	−0.307 (−7.41)	−0.307 (−7.41)	−0.216 (−5.22)	−0.233 (−5.61)
Breadth of Ownership	1.274 (7.24)	1.272 (7.23)	0.531 (2.95)	0.490 (2.71)
Ln Deal Value	0.004 (0.41)	0.005 (0.45)	0.017 (1.69)	0.018 (1.78)
Cashonly	−0.050 (−2.33)	−0.050 (−2.34)	0.044 (2.02)	0.051 (2.36)
Unsolicited	0.087 (1.80)	0.086 (1.78)	0.097 (2.04)	0.102 (2.15)
Target is Public	−0.019 (−0.91)	−0.019 (−0.90)	−0.063 (−2.99)	−0.061 (−2.88)
Ln Market Cap	−0.087 (−6.31)	−0.088 (−6.34)	−0.119 (−8.75)	−0.120 (−8.61)
Market to Book	0.001 (0.40)	0.001 (0.40)	−0.004 (−1.56)	−0.004 (−1.45)
Leverage	0.217 (4.15)	0.217 (4.15)	0.201 (3.85)	0.212 (4.03)
Past one-year CAR	−0.057 (−4.20)	−0.057 (−4.21)	−0.046 (−3.36)	−0.050 (−3.63)
Volatility	0.002 (0.04)	0.002 (0.04)	−0.125 (−2.61)	−0.137 (−2.73)
Lagged price inf.	0.246 (27.33)	0.246 (27.33)	0.195 (21.38)	0.191 (20.49)
Normal Volume	−0.093 (−9.61)	−0.093 (−9.61)	−0.044 (−4.58)	−0.042 (−4.29)
Runup	−0.208 (−4.24)	−0.208 (−4.24)	−0.198 (−4.08)	−0.202 (−4.16)
Announcement AR	0.185 (0.95)	0.184 (0.95)	0.239 (1.25)	0.199 (1.04)
Same Industry	−0.056 (−2.47)	−0.056 (−2.46)	−0.056 (−2.53)	−0.064 (−2.86)
Trend			0.045 (7.79)	
Trend squared			−0.003 (−12.76)	
constant	2.202 (23.03)	2.201 (23.03)	1.953 (20.50)	1.902 (17.18)
Year FE	No	No	2012-2018	Yes
Industry FE	Yes	Yes	Yes	Yes
R^2	0.252	0.252	0.275	0.287
N	12519	12519	12519	12274

Table 9
Price informativeness using the Paramount dummy and interaction terms as instruments for acquirer termination fee using sub-samples selected on acquirer announcement day abnormal returns below −0.02 and above 0.02

The sample consists of 14,545 mergers announced between 1986 and 2018. When the sample is required to have non-missing covariates, the sample size is 12,519. The merger data is from Refinitiv SDC Platinum (formerly Thomson Reuters SDC Platinum). The columns named "Negative signal" contain all observations with acquirer Announcement AR below −0.02. The columns named "Positive signal" contain all observations with acquirer Announcement AR above 0.02. The coefficients are estimated using 2SLS and standard errors are robust to heteroskedasticity. Parentheses in the columns named "First stage" contain t-values. Parentheses in the columns named "Probit" and "Second stage" contain z-values. The rows labeled "F-stat weak instr" contains the F-statistic for the Montiel-Pflueger robust weak instrument test. The row "5% critical value" contains the 5% critical values for a relative bias (IV bias divided by worst case OLS bias) of 5%. The weak IV tests are performed using the Stata function weakivtest. All variables are defined in Appendix Table A1.

	Negative signal			Positive signal		
	Probit	1st stage	2nd stage	Probit	1st stage	2nd stage
Term fee D			0.508 (2.82)			−0.109 (−0.51)
$\Phi(\hat{\theta}_0 + x_i\hat{\theta} + z_i\hat{\theta}_z)$		1.087 (14.68)			0.979 (12.11)	
Other Controls	Yes	Yes	Yes	Yes	Yes	Yes
Year Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Industry FE	Yes	Yes	Yes	Yes	Yes	Yes
pseudo R^2/R^2	0.335	0.290	0.261	0.324	0.268	0.264
F-stat weak instr		100.900			77.42	
5% critical value		29.520			30.220	
Observations	2757	2757	2757	2729	2729	2729

Table 10
Endogenous switching model

The sample consists of 14,545 mergers announced between 1986 and 2018. The merger data is from Refinitiv SDC Platinum (formerly Thomson Reuters SDC Platinum). The estimated model is equations (4), (5), and (11). Parentheses contain z-values. Square brackets contain 95% confidence intervals. All variables are defined in Appendix Table A1.

	Quadratic trend		Year fixed effects	
Variable	Selection: Termination Fee D	Outcome: Price informativ	Selection: Termination Fee D	Outcome: Price informativ
A. Treatment effect model				
Term fee D		0.531 (3.500)		0.557 (3.860)
Param dummy	0.938 (3.900)		1.426 (3.100)	
Param × Deal size	0.268 (4.020)		0.287 (4.160)	
Param × Cashonly	−1.165 (−5.430)		−1.042 (−4.710)	
Param × Bidder size	−0.278 (−5.020)		−0.281 (−4.940)	
Param × Bidder leverage	0.690 (1.650)		0.780 (1.790)	
Param × Runup	−0.410 (−1.370)		−0.736 (−2.630)	
Other Controls	Yes	Yes	Yes	Yes
Year fixed effects	No	No	Yes	Yes
Industry fixed effects	Yes	Yes	Yes	Yes
Number of observations	12,519	12,519	12,519	12,519
B. Model error correlations				
ρ_0		−0.435 [−0.703, −0.059]		−0.469 [−0.733, −0.082]
ρ_1		−0.219 [−0.370, −0.057]		−0.223 [−0.366, −0.070]
σ_0		0.969 [0.946, 0.992]		0.960 [0.936, 0.984]
σ_1		0.984 [0.929, 1.043]		0.973 [0.917, 1.032]
$\sigma_{\epsilon 0}$		−0.421 [−0.749, −0.093]		−0.450 [−0.779, −0.121]
$\sigma_{\epsilon 1}$		−0.216 [−0.374, −0.058]		−0.217 [−0.365, −0.069]
Wald test: $\rho_0 = \rho_1 = 0$		$Pr(\chi^2_2 > \hat{\chi}^2_2 = 8.040) = 0.018$		$Pr(\chi^2_2 > \hat{\chi}^2_2 = 9.290) = 0.010$

Table 11
Endogenous switching model using the year fixed effects specification and sub-samples selected on acquirer announcement day abnormal returns below −0.02 and above 0.02

The sample consists of 14,545 mergers announced between 1986 and 2018. The merger data is from Refinitiv SDC Platinum (formerly Thomson Reuters SDC Platinum). The columns named "Negative signal" contain all observations with acquirer Announcement AR below −0.02. The columns named "Positive signal" contain all observations with acquirer Announcement AR above 0.02. The estimated model is equations (4), (5), and (11). Parentheses contain z-values. Square brackets contain 95% confidence intervals. All variables are defined in Appendix Table A1.

	Negative signal		Positive signal	
Variable	Selection: Termination Fee D	Outcome: Price informativ	Selection: Termination Fee D	Outcome: Price informativ
A. Treatment effect model				
Term fee D		0.832 (4.010)		−0.114 (−0.430)
Param dummy	1.179 (1.590)		6.983 (5.760)	
Param × Deal size	0.347 (2.930)		0.080 (0.440)	
Param × Cashonly	−0.883 (−2.190)		−1.700 (−3.370)	
Param × Bidder size	−0.394 (−4.210)		−0.241 (−1.720)	
Param × Bidder leverage	0.248 (0.300)		4.410 (2.060)	
Param × Runup	−1.129 (−2.430)		−1.112 (−1.840)	
Other Controls	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes
Industry fixed effects	Yes	Yes	Yes	Yes
Number of observations	2,854	2,854	2,919	2,919
B. Model error correlations				
ρ_0		−0.623 [−0.851, −0.197]		0.204 [−0.094, 0.468]
ρ_1		−0.366 [−0.567, −0.123]		0.0991 [−0.256, 0.431]
σ_0		0.956 [0.909, 1.006]		0.959 [0.932, 0.988]
σ_1		1.001 [0.897, 1.118]		0.966 [0.864, 1.080]
$\sigma_{\epsilon 0}$		−0.596 [−0.930, −0.261]		0.196 [−0.082, 0.473]
$\sigma_{\epsilon 1}$		−0.366 [−0.613, −0.120]		0.0957 [−0.253, 0.445]
Wald test: $\rho_0 = \rho_1 = 0$		$Pr(\chi^2_2 > \hat{\chi}^2_2 = 11.690) = 0.003$		$Pr(\chi^2_2 > \hat{\chi}^2_2 = 1.84) = 0.398$

Table 12
Price informativeness and acquirer termination fees in a subsample of mergers with non-listed and listed targets

The merger data is from Refinitiv SDC Platinum (formerly Thomson Reuters SDC Platinum). Coefficients are estimated using OLS, 2SLS, and the endogenous switching model. Standard errors are robust to heteroskedasticity. Parentheses contain T-statistics in the case of OLS and z-statistics in the case of IV and the endogenous switching model. All variables are defined in Appendix Table A1.

	Dependent variable: Price informativeness			
	Non-listed targets		Listed targets	
A. OLS				
Term fee D	0.131 (1.6)	0.141 (1.7)	0.135 (2.82)	0.146 (3.06)
Control variables	Yes	Yes	Yes	Yes
Quadratic trend	Yes	No	Yes	No
Year fixed effects	2012-2018	Yes	2012-2018	Yes
Industry fixed effects	Yes	Yes	Yes	Yes
R^2	0.272	0.273	0.293	0.305
Number of observations	6,772	6,772	3,897	3,897
B. IV				
Term fee D	0.916 (2.65)	0.676 (1.94)	0.650 (3.31)	0.742 (3.81)
Control variables	Yes	Yes	Yes	Yes
Quadratic trend	Yes	No	Yes	No
Year fixed effects	2012-2018	Yes	2012-2018	Yes
Industry fixed effects	Yes	Yes	Yes	Yes
R^2	0.262	0.295	0.272	0.278
Number of observations	6,772	6,772	3,897	3,897
C. Endogenous Switching				
Term fee D	1.304 (3.34)	1.246 (3.24)	0.654 (3.16)	0.783 (4.35)
Control variables	Yes	Yes	Yes	Yes
Quadratic trend	Yes	No	Yes	No
Year fixed effects	2012-2018	Yes	2012-2018	Yes
Industry fixed effects	Yes	Yes	Yes	Yes
Number of observations	6,772	6,772	3,897	3,897

A Tables

Table A1
Variable Definitions

Variable	Definition
A. Dependent variables	
Price informativeness	The natural logarithm of the ratio of variation in daily stock returns caused by firm specific information to the variation in stock returns caused by the market, the industry, the size factor and the book-to-market factor. Returns are measured over the period +2 through +22 relative to the event date.
Withdrawn	Dummy variable equal to one if SDC status is "Withdrawn" and equal to zero if SDC status is "Completed."
B. Covariates	
Announcement AR	Abnormal return on the day after the announcement date for the merger. The model for computing normal return is estimated over the period −294 through −43 with an minimum of 126 observations and using the Fama-French three-factor model. The AR is trimmed at the 1st and 99th percentile.
Breadth of Ownership	The number of bidder shareholders filing 13F forms prior to the event date divided by the total number of institutions filing 13F forms prior to the event date.
Cashonly	Equal to one if SDC variable "percent cash" equals 100% or SDC variable "Consideration structure" equals CASHO. Zero otherwise.
Deal Value	Deal value as reported by Refinitiv SDC Platinum.
Ind. fixed effects	Based on Fama-French 12 industry portfolios
Inst. Ownership	Aggregate bidder ownership by institutions filing 13F forms. Measured before but as close to the event date as possible.
Lagged price inf.	The natural logarithm of the ratio of variation in daily stock returns caused by firm specific information to the variation in stock returns caused by market and industry factors. Returns are measured over the period −84 through −43 relative to the event date.
Leverage	Total debt divided by the sum of total debt and Market Cap (defined below). Total debt is sum of Compustat items dltt and dlc.
Market Cap	Bidder market capitalization on day −42 relative to the event date.
Market to Book	The bidder Market Cap divided by the book value of equity. Book value is Compustat book value of stockholders equity, plus balance sheet deferred taxes and investment tax credits (if available), minus the book value of preferred stock. Depending on availability, we use the redemption, liquidation, or par value (in that order) to estimate the value of preferred stock.
Normal volume	Average traded bidder volume over the period −294 through −43 relative to the event date.
Paramount	Dummy variable equal to one for the period 1994 through 2018 and zero for the period 1986 through 1993. Instrument for bidder termination fee.
Past one-year CAR	Cumulative abnormal daily return over the period −294 through −43 relative to the event date. The model for computing normal return is estimated over the period −546 through −295 with an minimum of 126 observations and using the Fama-French three-factor model.

. . . continued on next page

... Table A1 continued

Variable	Definition
Runup	Cumulative abnormal daily return over the period −42 through −2 relative to the event date. The model for computing normal return is estimated over the period −294 through −43 with an minimum of 126 observations and using the Fama-French three-factor model.
Same industry	Acquirer and target is in the same industry based on Fama-French 12 industries.
Target is Public	The target is listed on a stock exchange. Classified by SDC.
Termination Fee D	Dummy variable equal to one if the merger agreement includes a bidder termination fee and the Termination Fee Size (defined below) is larger than its 25%-quantile (of all nonzero Termination Fee Size observations) and zero otherwise.
Termination Fee Size	The size of the fee to be paid by the bidder to the target if the bidder terminates the transaction divided by Market Cap (see above). This ratio is then multiplied by the Termination Fee Dummy (defined above).
Total Assets	Bidder total assets (Compustat variable at) from the last fiscal year prior to the event date.
Unsolicited	Equal to one if SDC has classified the deal as resisted by management and zero otherwise.
Volatility	Annualized daily volatility. Daily volatility computed over trading days −294 through −43 relative to the event date.

Table A2
Price informativeness and acquirer termination fees using subsamples selected on acquirer announcement day abnormal returns below −0.01 and above 0.01

The sample consists of 14,545 mergers announced between 1986 and 2018. When the sample is required to have non-missing covariates, the sample size is 12,519. The merger data is from Refinitiv SDC Platinum (formerly Thomson Reuters SDC Platinum). The column named "Negative signal" contains all observations with acquirer Announcement AR below −0.01. The column named "Positive signal" contains all observations with acquirer Announcement AR above 0.01. Coefficients are estimated using OLS and standard errors are robust to heteroskedasticity. T-statistics in parentheses. All variables are defined in Appendix Table A1.

	Dependent variable: Price informativeness			
	Negative signal	Positive signal	Negative signal	Positive signal
A. Termination Fee Dummy				
Term fee D	0.121 (2.05)	0.040 (0.62)	0.134 (2.28)	0.049 (0.75)
Inst. Ownership	−0.232 (−3.30)	−0.191 (−2.68)	−0.249 (−3.54)	−0.195 (−2.72)
Breadth of Ownersh.	0.716 (2.43)	0.727 (1.91)	0.617 (2.12)	0.767 (2.03)
Ln Deal Value	0.026 (1.84)	0.024 (1.78)	0.026 (1.85)	0.028 (2.03)
Cashonly	0.026 (0.71)	0.086 (2.28)	0.036 (0.97)	0.095 (2.53)
Unsolicited	0.060 (0.78)	0.065 (0.70)	0.074 (0.96)	0.067 (0.72)
Target is Public	−0.043 (−1.21)	−0.040 (−1.10)	−0.040 (−1.12)	−0.031 (−0.87)
Ln Firm Value	−0.148 (−6.62)	−0.113 (−5.14)	−0.144 (−6.22)	−0.118 (−5.24)
Market to book	0.000 (−0.06)	−0.008 (−1.96)	−0.001 (−0.24)	−0.008 (−1.87)
Leverage	0.278 (3.14)	0.098 (1.06)	0.275 (3.10)	0.127 (1.36)
Past one-year CAR	−0.034 (−1.56)	−0.049 (−2.22)	−0.037 (−1.66)	−0.053 (−2.43)
Volatility	−0.136 (−1.74)	−0.060 (−0.72)	−0.127 (−1.46)	−0.056 (−0.62)
Lagged price inf.	0.181 (11.57)	0.166 (10.33)	0.169 (10.57)	0.162 (9.87)
Normal volume	−0.030 (−1.79)	−0.066 (−4.14)	−0.030 (−1.73)	−0.065 (−4.05)
Runup	−0.109 (−1.35)	−0.268 (−3.45)	−0.109 (−1.36)	−0.275 (−3.57)
Announce CAR	0.095 (0.19)	0.717 (1.70)	−0.001 (0.00)	0.465 (1.12)
Same Industry	−0.029 (−0.76)	−0.056 (−1.54)	−0.033 (−0.86)	−0.050 (−1.36)
Trend	0.034 (3.40)	0.053 (4.79)		
Trend squared	−0.002 (−5.97)	−0.003 (−7.28)		
constant	1.964 (11.49)	2.065 (12.01)	1.918 (9.58)	2.142 (9.74)
Year fixed effects	2012-2018	2012-2018	Yes	Yes
Industry fixed effects	Yes	Yes	Yes	Yes
R-sq	0.264	0.254	0.273	0.267
Number of obs.	4,185	4,177	4,185	4,177
B. Termination Fee Size				
Termination Fee Size	5.296 (3.92)	1.606 (1.14)	5.524 (3.97)	1.455 (1.02)
Other Control var.	Yes	Yes	Yes	Yes
Year fixed effects	2012-2018	2012-2018	Yes	Yes
Industry fixed effects	Yes	Yes	Yes	Yes
R^2	0.266	0.255	0.275	0.267
Number of obs.	4,185	4,177	4,185	4,177

Table A3

Price informativeness using the Paramount dummy and interaction terms as instruments for acquirer termination fee using sub-samples selected on acquirer announcement day abnormal returns below -0.01 and above 0.01

The sample consists of 14,545 mergers announced between 1986 and 2018. When the sample is required to have non-missing covariates, the sample size is 12,519. The merger data is from Refinitiv SDC Platinum (formerly Thomson Reuters SDC Platinum). The columns named "Negative signal" contain all observations with acquirer Announcement AR below -0.01. The columns named "Positive signal" contain all observations with acquirer Announcement AR above 0.01. The coefficients are estimated using 2SLS and standard errors are robust to heteroskedasticity. Parentheses in the columns named "First stage" contain t-values. Parentheses in the columns named "Probit" and "Second stage" contain z-values. The row labeled "Effective F-statistic" contains the F-statistic for the Montiel-Pflueger robust weak instrument test. The row "5% critical value" contains the 5% critical values for a relative bias (IV bias divided by worst case OLS bias) of 5%. The weak IV tests are performed using the Stata function weakivtest. All variables are defined in Appendix Table A1.

	Negative signal			Positive signal		
	Probit	1st stage	2nd stage	Probit	1st stage	2nd stage
Term fee D			0.507 (3.13)			0.005 (0.02)
$\Phi(\hat{\theta}_0 + x_i\hat{\theta} + z_i\hat{\theta}_z)$		1.062 (16.26)			0.986 (13.27)	
Other Controls	Yes	Yes	Yes	Yes	Yes	Yes
Year Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Industry FE	Yes	Yes	Yes	Yes	Yes	Yes
pseudo R^2/R^2	0.351	0.285	0.267	0.334	0.246	0.274
Effect. F-stat.		138.800			99.84	
5% critical value		30.280			30.54	
Observations	4,041	4,041	4,041	3,995	3,995	3,995

Table A4

Endogenous switching model using the quadratic trend specification and sub-samples selected on acquirer announcement day abnormal returns below −0.02 and above 0.02

The sample consists of 14,545 mergers announced between 1986 and 2018. When the sample is required to have non-missing covariates, the sample size is 12,519. The merger data is from Refinitiv SDC Platinum (formerly Thomson Reuters SDC Platinum). The columns named "Negative signal" contain all observations with acquirer Announcement AR below −0.02. The columns named "Positive signal" contain all observations with acquirer Announcement AR above 0.02. The estimated model is equations (4), (5), and (11). Parentheses contain z-values. Square brackets contain 95% confidence intervals. All variables are defined in Appendix Table A1.

	Negative signal		Positive signal	
Variable	Selection: Termination Fee D	Outcome: Price informativ	Selection: Termination Fee D	Outcome: Price informativ
A. Treatment effect model				
Term fee D		1.175 (5.12)		−0.045 (−0.13)
Param dummy	0.956 (2.93)		1.118 (2.35)	
Param × Deal size	0.179 (2.05)		0.207 (1.66)	
Param × Cashonly	−0.578 (−1.57)		−1.039 (−2.21)	
Param × Bidder size	−0.235 (−2.57)		−0.252 (−1.50)	
Param × Bidder leverage	0.962 (1.32)		0.846 (0.97)	
Param × Runup	−1.101 (−1.44)		−0.593 (−1.42)	
Other Controls	Yes	Yes	Yes	Yes
Year fixed effects	No	No	No	No
Industry fixed effects	Yes	Yes	Yes	Yes
Number of observations	2,854	2,854	2,854	2,854
B. Model error correlations				
ρ_0		−0.742 [−0.847, −0.583]		0.150 [−0.238, 0.497]
ρ_1		−0.524 [−0.713, −0.263]		0.0560 [−0.320, 0.417]
σ_0		0.989 [0.948, 1.031]		0.972 [0.943, 1.002]
σ_1		1.100 [0.971, 1.246]		0.976 [0.862, 1.105]
$\sigma_{\epsilon 0}$		−0.734 [−0.885, −0.583]		0.146 [−0.229, 0.521]
$\sigma_{\epsilon 1}$		−0.577 [−0.875, −0.278]		0.0546 [−0.323, 0.433]
Wald test: $\rho_0 = \rho_1 = 0$		$Pr(\chi^2_2 > \hat{\chi}^2_2 = 49.55) = 0.000$		$Pr(\chi^2_2 > \hat{\chi}^2_2 = 0.57) = 0.751$

Table A5
Average treatment effects

The sample consists of 14,545 mergers announced between 1986 and 2018. The merger data is from Refinitiv SDC Platinum (formerly Thomson Reuters SDC Platinum). Parentheses contain T-statistics (OLS) or z-values (IV, endogenous switching model).

Model	Number of obs.	ATE	ATT
A. OLS			
Full Sample	12,519	0.123 (3.30)	0.123 (3.30)
Negative Signal	2,854	0.151 (2.33)	0.151 (2.33)
Positive Signal	2,919	0.067 (0.91)	0.067 (0.91)
B. IV			
Full Sample	12519	0.448 (3.86)	0.448 (3.86)
Negative Signal	2757	0.508 (2.82)	0.508 (2.82)
Positive Signal	2729	−0.109 (−0.51)	−0.109 (−0.51)
C. Endogenous Switching Model			
Full Sample	12519	0.557 (3.861)	0.872 (3.041)
Negative Signal	2854	0.832 (4.012)	1.099 (3.831)
Positive Signal	2919	−0.114 (−0.431)	−0.236 (−0.999)

B OLS And IV Bias

First, we derive the bias of the OLS estimates of equation (7) if the true model is the endogenous switching model of equations (4), (5), and (11).

We employ the finding that the linear projection of y on (x, t) is the same as the linear projection $E[y|t, x, z]$ on (x, t). x is the vector of outcome equation covariates, t is the termination fee dummy, and z is the vector of instruments based on the Paramount ruling. Equation (12) showed that the conditional expectation, $E[y|t, x, z]$, is linear in the coefficients:

$$E[y|t, x, z] = \alpha_0 + x\beta_0 + ATE\ t - (1 - t)\sigma_{0\epsilon}\frac{\phi(q\theta)}{1 - \Phi(q\theta)} + t\sigma_{1\epsilon}\frac{\phi(q\theta)}{\Phi(q\theta)}$$

where ATE is $(\alpha_1 - \alpha_0)$ and q is the vector of exogenous variables $q = (x, z)$.

In the following, we group the intercept and the coefficients of the covariate vector x in the vector $\beta = (\alpha_0, \beta_0)'$. The population equivalent to our sample OLS regression (7) is the linear projection of y on the vector (x, t), whose coefficients we denote by $(\ddot{\beta}, A\ddot{T}E)$.

$$\begin{aligned}
\begin{pmatrix} \hat{\beta} \\ \widehat{\text{ATE}} \end{pmatrix} &= E[(x,t)'(x,t)]^{-1}E[(x,t)'E[y|t,x,z]] \\
&= \begin{pmatrix} \beta \\ \text{ATE} \end{pmatrix} + E[(x,t)(x,t)']^{-1}E\left[(x,t)'(-(1-t)\sigma_{0\epsilon}M_0(q\theta) + t\sigma_{1\epsilon}M_1(q\theta))\right] \\
&= \begin{pmatrix} \beta \\ \text{ATE} \end{pmatrix} + E[(x,t)'(x,t)]^{-1} \begin{pmatrix} E\left[x'\phi(q\theta)\right](\sigma_{1\epsilon} - \sigma_{0\epsilon}) \\ \sigma_{1\epsilon}E\left[\phi(q\theta)\right] \end{pmatrix}
\end{aligned}$$

Unless $\sigma_{0\epsilon}$ and $\sigma_{1\epsilon}$ are zero, the OLS regression won't identify β and ATE. To establish the sign of the bias, we evaluate the inverse of the partitioned matrix $E[(x,t)'(x,t)]$: [24]

$$E\left[(x,t)'(x,t)\right] = \begin{pmatrix} E[x'x] & E[x't] \\ E[tx] & E[t] \end{pmatrix} = \begin{pmatrix} Q_{xx} & Q_{xt} \\ Q_{tx} & Q_{tt} \end{pmatrix}$$
$$E\left[(x,t)'(x,t)\right]^{-1} = \begin{pmatrix} Q_{xx\cdot t}^{-1} & -Q_{xx\cdot t}^{-1}Q_{xt}Q_{tt}^{-1} \\ Q_{tt\cdot x}^{-1}Q_{tx}Q_{xx}^{-1} & Q_{tt\cdot x}^{-1} \end{pmatrix}$$

where $Q_{tt\cdot x} = E[(t - x\pi)^2]$ is the residual variance of the projection of t on x, denoted by $L(t|x) = x\pi$. This projection is the linear probability model for $P(t = 1|x)$. The matrix multiplication for the row representing $\widehat{\text{ATE}}$ yields:

$$\begin{aligned} \widehat{\text{ATE}} &= \text{ATE} + \frac{-E[tx]E[x'x]^{-1}E[x'\phi(q\theta)](\sigma_{1\epsilon} - \sigma_{0\epsilon}) + E[\phi(q\theta)]\sigma_{1\epsilon}}{Q_{tt\cdot x}} \\ &= \text{ATE} + \frac{-\pi' E[x'\phi(q\theta)](\sigma_{1\epsilon} - \sigma_{0\epsilon}) + E[\phi(q\theta)]\sigma_{1\epsilon}}{Q_{tt\cdot x}} \\ &= \text{ATE} + \frac{1}{Q_{tt\cdot x}} E\left[\left((1 - \pi' x')\sigma_{1\epsilon} + \pi' x' \sigma_{0\epsilon}\right)\phi(q\theta)\right] \end{aligned} \tag{13}$$

Since $\phi(q\theta) > 0$, the sign of the bias depends on the linear combination of $\sigma_{0\epsilon}$ and $\sigma_{1\epsilon}$. The weight of each covariance for a given realization x is given by the linear probability model $x\pi$. [25]

Next, we derive the bias of the estimates resulting of the IV approach depicted in equations (8), (9), and (10), if the true model is the endogenous switching model of equations (4), (5), and (11). The second stage regression of the IV approach has the following population regression equivalent:

[24] See, e.g., Hansen (2022) for the inverse of a partitioned matrix.

[25] Our result is the population regression counterpart to the bias created by being not able to match on the right control variables derived in Heckman and Navarro-Lozano (2004).

$$
\begin{pmatrix} \hat{\beta} \\ \widehat{\text{ATE}} \end{pmatrix} = E\left[(x,t^*)'(x,t)\right]^{-1} E\left[(x,t^*)' E[y|x,t]\right]
$$

$$
= \begin{pmatrix} \beta \\ \text{ATE} \end{pmatrix} + E\left[(x,t^*)'(x,t)\right]^{-1} E\left[(x,t^*)'\left(-(1-t)M_0(q\theta), tM_1(q\theta)\right)\right] \begin{pmatrix} \sigma_{0\epsilon} \\ \sigma_{1\epsilon} \end{pmatrix}
$$

$$
= \begin{pmatrix} \beta \\ \text{ATE} \end{pmatrix} + E\left[(x,t^*)'(x,t)\right]^{-1} \begin{pmatrix} E\left[x'\phi(q\theta)\right](\sigma_{1\epsilon} - \sigma_{0\epsilon}) \\ E\left[t^*\phi(q\theta)\right](\sigma_{1\epsilon} - \sigma_{0\epsilon}) \end{pmatrix}
$$

Using a similar partitioning of the matrix $E\left[(x,t^*)'(x,t)\right]$, we obtain its inverse as:

$$
E\left[(x,t^*)'(x,t)\right] = \begin{pmatrix} E[x'x] & E[x't] \\ E[t^*x] & E[t^*t] \end{pmatrix} = \begin{pmatrix} Q_{xx} & Q_{xt} \\ Q_{t^*x} & Q_{t^*t} \end{pmatrix}
$$

$$
E\left[(x,t^*)'(x,t)\right]^{-1} = \begin{pmatrix} Q_{xx\cdot t}^{-1} & -Q_{xx\cdot t}^{-1} Q_{xt} Q_{t^*t}^{-1} \\ Q_{t^*t\cdot x}^{-1} Q_{t^*x} Q_{xx}^{-1} & Q_{t^*t\cdot x}^{-1} \end{pmatrix}
$$

where $Q_{t^*t\cdot x} = E[t^*t] - E[t^*x]E[x'x]^{-1}E[x't]$ and represents the variance of the residuals from linear projection of t^* on x, denoted by $x\pi$.

The resulting large-sample bias of the IV estimator can be written as:

$$
\widehat{\text{ATE}} = \text{ATE} + \frac{\left(-E[t^*x]E[x'x]^{-1}E[x'\phi(q\theta)] + E[t^*\phi(q\theta)]\right)(\sigma_{1\epsilon} - \sigma_{0\epsilon})}{Q_{t^*t\cdot x}}
$$

$$
= \text{ATE} + \frac{E[-\pi'x'\phi(q\theta) + t^*\phi(q\theta)](\sigma_{1\epsilon} - \sigma_{0\epsilon})}{Q_{t^*t\cdot x}}
$$

$$
= \text{ATE} + \frac{1}{Q_{t^*t\cdot x}} E\left[(t^* - \pi'x')\phi(q\theta)\right](\sigma_{1\epsilon} - \sigma_{0\epsilon}) \tag{14}
$$

The sign of the IV bias depends on the signs of $(\sigma_{1\epsilon} - \sigma_{0\epsilon})$ and $E\left[(t^* - \pi'x')\phi(q\theta)\right]$. As discussed in section 4.3, this bias vanishes, if u_0 and u_1 have the same distribution, which implies identical covariances $\sigma_{0\epsilon}$ and $\sigma_{1\epsilon}$.