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The Takeover Deterrent Effect of Open Market Share Repurchases

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ABSTRACT

This paper examines whether open market share repurchases deter takeovers. We model pre-repurchase takeover probability as a latent variable and examine its impact on the firm's decision to repurchase shares. Given specification tests reject the Tobit model, we turn to the censored quantile regression method of Powell (1986, *Journal of Econometrics* 32, 143–155). We find a significantly positive relation between open market share repurchases and takeover probability, and we reconcile empirical findings in previous studies that contradict predictions. Repurchase activity is inversely related to firm size, consistent with smaller firms having greater information asymmetry, and is related to temporary, but not permanent, cash flows.

OPEN MARKET SHARE REPURCHASE ACTIVITY has risen since the mid 1980s to become the most common method for a firm to repurchase its stock.[1] Commonly cited motives for this development include distributing free cash flows, signaling undervaluation, funding executive/employee stock option plans, adjusting capital structure, and defending against unwanted takeovers.[2] However, existing empirical studies fail to find evidence that takeover deterrence is a significant motive for open market share repurchases (Dittmar (2000)). This is perhaps surprising, given the evidence indicates that other methods of repurchases, such as fixed price tender offers and Dutch auction tender offers, are often used to thwart unwanted takeover attempts.

One reason for this discrepancy could be that the two repurchase methods (tender offer vs. open market repurchase) affect the takeover process differently: While tender offers have been shown to act as an effective defense in the midst of takeover battles, open market repurchases may deter unwanted bids, pre-empting would-be acquirers from bidding in the first place. On the

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[1] Jagannathan, Stephens, and Weisbach (2000) report the number and volume of fixed price, Dutch auction, and open market share repurchases from 1985 through 1996. There are a total of 660 fixed price tender offers for $67 billion worth of stock, 120 Dutch auctions for $27 billion, and 4,753 open market announcements for $471 billion.

[2] See for example, Brav et al. (2004), Chan, Ikenberry, and Lee (2004), Dittmar (2000), Grullon and Michaely (2002, 2004), Jagannathan, Stephens, and Weisbach (2000), Ikenberry, Lakonishok, and Vermaelen (1995), and Kahle (2002).

one hand, repurchases may reduce agency costs and make the company a less attractive target by reducing the gains from a disciplinary takeover. Alternatively, Bagwell (1991a) shows that if investors have heterogeneous valuations, repurchases increase the cost of a takeover by affecting the supply of shares available to a would-be acquirer. Bagwell cites the case of Sears, which announced a plan to conduct an open market repurchase of 10% of its shares in response to rumors that the company would soon become a target.[3] The preemptive repurchase appears to have been successful since no formal bid was made.

In this paper we revisit the motives of open market repurchases and hypothesize that firms may conduct open market repurchases in anticipation of possible takeover attempts. Specifically, we test whether firms repurchase shares when the probability of becoming a target is high, effectively deterring takeover bids from materializing, rather than in response to receiving an actual bid. Because previous studies examine the takeover hypothesis by regressing open market share repurchase activity on a dummy variable indicating whether a firm receives a takeover bid, they are not designed to capture this deterrent effect. To the extent that repurchases deter bids, we expect open market share repurchases to be associated with fewer ex post takeover bids, and the coefficient of the dummy variable to be biased downward.

We hypothesize that the ex ante rather than the ex post takeover probability captures the takeover deterrent effect, so we measure the perceived takeover probability *prior* to the repurchase decision. We first model takeover probability as a latent variable using prerepurchase firm characteristics. We then regress share repurchase activity on the perceived threat of a takeover. In contrast to previous studies, we find a statistically and economically significant coefficient on the takeover probability variable. This suggests that takeover deterrence plays a more important role for open market repurchases than previous studies suggest.[4]

Following existing studies, we initially examine open market repurchase activity by estimating Tobit models (Dittmar (2000), Fenn and Liang (2001), Stephens and Weisbach (1998)). However, we find that specification tests of the Tobit model soundly reject its two key assumptions of normally distributed and homoskedastic error terms.[5] Since the Tobit estimates are unreliable, we turn

[3] For a discussion of Sears's repurchase program, see "Sears to Sell Tower, Buy 10% of Stock Back—Restructuring also entails retail-strategy shift; Price of shares declines" (*Wall Street Journal*, November 1, 1988). For a discussion of the defensive nature of the repurchase, see "Sears Kicks Off Share Buy-Back As Stock Rises $3" (*Wall Street Journal*, November 4, 1988).

[4] When we replicate previous studies using a dummy variable for a takeover bid, we find no evidence of a takeover deterrent effect. Thus, the latent variable treatment of takeover probability appears to drive our result, rather than differences in sample or variable construction.

[5] Either heteroskedasticity or nonnormality of the error terms can lead to inconsistent estimates in the Tobit model. Chay and Powell (2001) summarize both the problem and the use of semiparametric censored regression as a solution. Greene (1997) summarizes the concerns of Tobit estimation in censored data, stating that many studies express "varying degrees of pessimism regarding how inconsistent the maximum likelihood estimator will be when heteroskedasticity occurs" (p. 967).

to the censored quantile regression (CQR) model proposed by Powell (1986).[6] The CQR model is a weighted version of the censored least absolute deviations (CLAD) method of Powell (1984) and is more appropriate when the data are highly censored (as is the case here; only 11% of our firm-year observations conduct open market share repurchases). More importantly, given the nature of our repurchase data, the CQR estimator places no requirement on the distribution of the errors and produces consistent estimates in the presence of heteroskedastic errors.

We find that estimates from the CQR model suggest a strong takeover deterrent effect. In particular, a firm that moves from the 5^{th} percentile value of takeover probability to the 95^{th} percentile will repurchase an additional 15.58% of outstanding shares. Thus, takeover deterrence appears to be a significant motive for open market share repurchases.

We also find that a number of results from the CQR model contradict those from the Tobit model. In general, the CQR results are more consistent with existing theory and empirical evidence, and they resolve difficulties in interpreting some of the Tobit results. For example, firm size has been argued to be inversely related to the degree of asymmetric information about firm value. However, in the Tobit estimation the coefficient on firm size is positive and statistically significant at the 1% level. This puzzling result suggests that large firms are more likely to be misvalued (Dittmar (2000)), contradicting the well documented result that abnormal returns to repurchase announcements are significantly negatively correlated with firm size (Ikenberry, Lakonishok, and Vermaelen (1995)). The result from the CQR model reconciles this contradiction: The coefficient on size is significantly negative.

Moreover, in contrast to the Tobit estimation results, the CQR estimation suggests that firms repurchase to distribute "temporary" nonoperating cash flows, rather than "permanent" operating cash flows. These results support the findings of Jagannathan, Stephens, and Weisbach (2000) who show that firms with more volatile cash flows prefer the financial flexibility of open market repurchases over dividends. Finally, we find that the time period dummy variable estimates in the Tobit model contradict the clear time pattern of share repurchase activity, whereas those in the CQR regressions change signs and are consistent with the time trend of share repurchases.

The specification test suggests the CQR model is correctly specified, that is, the test results show that we cannot reject the null hypothesis that the model is correctly specified. Greene (1997) points out that the assumptions of the Tobit model are often violated in practice. Our results indicate that, indeed, one needs to be careful when using Tobit models.

The remainder of the article is organized as follows. In Section I, we motivate the relationship between share repurchase and takeover activity. We present

[6] At the time Powell's CQR estimator was developed, computational difficulties made it less practical. However, advances in computer technology have overcome these difficulties. See Buchinsky (1998) for a guideline on applying these techniques and Brière et al. (2002) and Hochguertel (2003) for examples of the use of CQR models and CLAD estimators.

the data and sample in Section II. Section III describes our methodology, and Section IV reports estimation results. The conclusions follow in Section V.

I. Repurchase and Takeover Probability

Numerous studies show how firms use share repurchases to deter unwanted takeover attempts. Harris and Raviv (1988) and Stulz (1988) model how firms issue debt and use the proceeds to repurchase shares to deter takeovers. Bagwell (1991a) shows that when a firm's shareholders have heterogeneous valuations, the shareholders who sell shares in a repurchase have relatively low valuations, but the remaining shareholders have a relatively high valuation, which increases the cost to any would-be acquirer. In Bagnoli and Lipman (1989), stock repurchases serve as a defense against takeover by signaling management's private information about firm value. While a large body of empirical evidence supports these theories in the case of tender offers, there is little evidence in relation to open market repurchases (Bagwell (1991b, 1992), Brown and Ryngaert (1992), Comment and Jarrell (1991), Denis (1990), Hodrick (1999)).[7]

In contrast to tender offers, which are a common defense in the midst of a hostile takeover attempt, open market repurchases may actually deter potential bidders. This deterrent effect of open market repurchases could arise for a number of reasons. First, open market repurchases may increase the cost to potential bidders by reducing the number of shares held by those who value control the least (Bagwell (1991a)).[8] This would increase both the cost of gaining a toehold and the price of the actual offer.

Moreover, open market repurchases may alleviate agency concerns and their associated costs, reducing potential gains from disciplinary takeovers. Hirshleifer and Thakor (1992) develop a model in which managers take actions in anticipation of possible takeover attempts. Specifically, they consider a reputation equilibrium and analyze the use of debt to deter a takeover attempt. In their model, if takeover targets are firms whose managers are perceived to be doing poorly, then managers will increase leverage during periods of high takeover activity. Applying their rationale to open market repurchases, repurchases can be used to increase leverage in anticipation of a takeover attempt. Moreover, repurchases may result in less free cash and increased insider ownership, characteristics that may diminish the potential gains to a would-be acquirer.

Finally, open market repurchase plans give the firm greater flexibility to repurchase at any time. This flexibility may act as deterrent to would-be acquirers, since the potential target can quickly respond by repurchasing shares.

[7] While most of these studies concentrate on tender offers, Denis (1990) takes a different approach. He identifies a sample of defensive share repurchases and special dividends during the course of a corporate control contest or in response to the threat of a change in control, during the period 1980 to 1987. Of a total of 49 cases, 10 are open market repurchases, 16 are tender offers, 7 are exchange offers, 7 are unknown repurchase types, and 9 are special dividends.

[8] As we mention in the introduction, Bagwell (1991a) cites the example of Sears using an open market repurchase to thwart rumored bidders from actually making a bid.

II. Data and Sample

Our sample begins with all the exchange-listed firms (NYSE, Amex, and NASDAQ) with data necessary to construct our variables available from CRSP and Compustat between 1985 and 1996. We start with 1985, since we use data from Compustat to measure the amount of shares that firms actually repurchased in the market.[9] We exclude financial firms, utilities, and firms with a share price below five dollars or with total assets of less than five million.[10]

We obtain open market share repurchase announcements from Securities Data Corporation (SDC). We begin with an initial sample of all open market repurchase announcements between 1985 and 1996. We then exclude programs that are withdrawn, privately negotiated, or announced in the fourth quarter of 1987. Following Stephens and Weisbach (1998), who find that the majority of repurchase programs last approximately 3 years, and that the initial repurchase announcement may be followed by subsequent announcements about the ongoing (previously announced) repurchase program, we also disregard repurchase announcements made within 3 years of a previous repurchase announcement.[11]

Next, we follow Jagannathan, Stephens, and Weisbach (2000) and identify our sample firms' gross repurchases each year for the 3 years subsequent to the announcement of the repurchase program from Compustat (Data item #115, Purchases of Common and Preferred Stock). This item minus the amount of privately negotiated, Dutch auction, and fixed price repurchases yields our estimated amount of actual open market repurchases. We use the actual repurchase amount scaled by the firm's prior year-end market value of equity to construct the variable *REP* as follows:

Repurchase (REP) is the annual amount of actual open market repurchases divided by the market value of common equity.

We also construct variables identified as important explanatory variables in previous studies of takeover probability (Palepu (1986), Comment and Schwert (1995), Billett (1996)) and open market repurchases (Stephens and Weisbach (1998), Dittmar (2000), Jagannathan, Stephens, and Weisbach (2000), Fenn and Liang (2001)). Using data from SDC's Mergers and Acquisitions database, we construct the following takeover variables:

Industry takeover dummy (ITODUM) equals one if at least one firm in the same industry (same four-digit SIC code) is a takeover target in the previous year, and zero otherwise.

Takeover dummy (TODUM) equals one if the firm is a takeover target in a given year, and zero otherwise (this definition is used to estimate the ex ante takeover probability, which we will discuss in Sections III and IV.A).

[9] Since 1984, firms have been required to report the value of repurchased shares on the Statement of Cash Flows, reported by Compustat as "Purchases of Common and Preferred Stock."

[10] Specifically, we exclude financials and utilities by eliminating firms with SIC codes from 4,800 through 4,829, 4,910 through 4,949, and 6,000 through 6,999.

[11] Our results are robust to whether we include the repurchase announcements that occur within 3 years of a previous announcement.

Targeted dummy (TARGET) equals one if the firm is a takeover target in either the year of or the year prior to the repurchase decision (this definition follows Dittmar (2000) and is used later in Section IV.B to replicate her results).

In addition, we construct the following control variables from Compustat and CRSP.

Cash dividends (CASHDIV) is cash dividends divided by net income.

Free cash flow (FCF) is operating income before depreciation minus interest expenses, the sum of preferred and common dividends, and income taxes (excluding deferred taxes), all scaled by total assets.

Excess return (EXRET) is the annual cumulative return on the firm's stock in the previous calendar year minus the annual cumulative return on the CRSP value-weighted market portfolio.

Fixed assets (*NPPE)* is net plant, property, and equipment scaled by total assets.

Industry-adjusted book value of leverage (LEVBIA) is the book value of debt divided by total assets, minus the median ratio for firms with the same two-digit SIC code.

Industry-adjusted market value of leverage (LEVIA) is the book value of debt divided by the sum of the book value of debt and the market value of equity, minus the median ratio for firms with the same two-digit SIC code.

Industry-adjusted return on assets (ROAIA) is operating income before depreciation divided by total assets, minus the median ratio for firms with the same two-digit SIC code.

Market-to-book ratio (MKBK) is the market value of common stock divided by the book value of common equity.

Nonoperating income (NONOPI) is nonoperating income scaled by total assets.

Return on equity (ROE) is net income divided by the book value of equity.

Sales growth (SALEGR) is the natural log of the ratio of sales over the sales of the previous year.

Size of assets (SIZEASST) is the natural log of total assets, where total assets is inflated to 1996 dollars using the consumer price index (CPI).

Size of equity (SIZEEQ) is the natural log of common equity, which is calculated as the number of shares outstanding times the previous year-end price, where price is inflated to 1996 dollars using the CPI.

Shares reserved for option exercise (OPTIONS) is common shares reserved for conversion of stock options divided by total shares outstanding.

Time trend dummy variables D1, D2, D3, and D4 equal one if the observation comes from 1985 to 1987, 1988 to 1990, 1991 to 1993, and 1994 to 1996, respectively, and zero otherwise.

We winsorize these variables at the 1% and 99% percentiles (with the exception of the dummy variables, *SIZEASST, SIZEEQ*, and *REP*) to avoid the influence of extreme observations. Our final sample spans 1985 through 1996 and consists of 23,208 firm-year observations.

We provide summary statistics in Table I. Panel A reports descriptive statistics. The mean of *TODUM* suggests that on average 6% of firms receive a takeover bid in a given year. Out of the 23,208 observations only 2,544 announce and conduct open market repurchase plans. In fact, both the mean and median values of *REP* are 0.00. Note that *REP* is highly censored. Panel B reports *REP* by year and across subperiods. We find the same time pattern in *REP* as documented by Allen and Michaely (2003): *REP* is relatively high in the mid 1980s, gradually declines in the early 1990s, and then levels off. This pattern is also apparent in the median values of *REP*. The variable *OPTIONS* is missing for 7.5% of our sample, so we estimate specifications both with and without this variable.

III. Estimation Methods

We develop a system of equations to analyze the effect of prerepurchase takeover probability on the repurchase decision. Our first equation explains the takeover probability process, and our second equation explains repurchase activity. The timing of our takeover measure relative to the repurchase decision is critical to our tests. We hypothesize that managers repurchase when they perceive the threat of a possible takeover attempt. To test this hypothesis, we need the perceived takeover probability prior to the repurchase. Such an ex ante measure of takeover probability is predicted to be positively correlated with future repurchase decisions under the deterrent hypothesis. In contrast, the deterrent hypothesis predicts a negative correlation between ex post measures of takeover probability and repurchases, given that fewer takeover bids should be observed ex post.

Dittmar (2000) uses a dummy variable indicating whether or not a takeover occurs around the repurchase. For our purposes, such a measure is problematic for two reasons. First, it contains information about the postrepurchase takeover status, which leads to the timing problem described above and imparts a downward bias on the resulting coefficient. Second, a dummy variable based on takeover status measures ex ante takeover probability with error. Thus, under the takeover deterrent hypothesis, using the dummy variable as a proxy (for ex ante takeover probability) in the repurchase model would introduce measurement error and lead to inconsistent estimates.[12]

To address these concerns, we adopt a procedure analogous to Hodrick (1999).[13] Hodrick investigates whether the choice of repurchase method (in her case fixed price versus Dutch auction tender offers) is related to a firm's expected inverse price elasticity. Observing that using the ex post realized elasticity as a proxy for the expected elasticity may introduce measurement error, Hodrick alleviates this concern by adopting a two-stage instrumental variables approach. In the first stage, she regresses realized inverse price elasticity on instrumental variables and firm characteristics to obtain predicted

[12] These problems may explain why Dittmar concludes that "increases in the probability of a takeover attempt have little effect on the probability of a stock repurchase" (p. 352).

[13] We thank the referee for pointing this out.

Table I

Descriptive Statistics

This table reports descriptive statistics for the firm-year sample as well as time trends of repurchases. *FCF* is operating income before depreciation minus interest expenses, the sum of preferred and common dividends, and income taxes (excluding deferred taxes), all scaled by total assets; *ROAIA* is operating income before depreciation divided by total assets, minus the median ratio for all firms within the same two-digit SIC code; *ROE* is the net income divided by common equity; *SIZEEQ* is the natural log of common equity, which is the number of shares outstanding times the previous year-end price inflated to 1996 dollars using the CPI; *SIZEASST* is the natural log of total assets, which is inflated to 1996 dollar using CPI; *LEVBIA* is debt divided by assets, minus the median ratio for all firms within the same two-digit SIC code; *LEVIA* is the book value of debt divided by the sum of the book value of debt and the market value of equity minus the median ratio for all firms within the same two-digit SIC code; *NPPE* is net plant, property, and equipment scaled by total assets; *SALEGR* is natural log of the ratio of sales over the sales of the previous year; *NONOPI* is nonoperating income scaled by total assets; *MKBK* is the market value of common stock to book value of equity; *CASHDIV* is the ratio of cash dividends paid to net income; *EXRET* is the stock's cumulative return in the previous calendar year minus the value-weighted market return; *ITODUM* equals one if at least one firm in the same industry (same four-digit SIC code) is a takeover target in the previous year, and zero otherwise; *TODUM* equals one if the firm is a takeover target in year t, and zero otherwise; *OPTIONS* is common shares reserved for conversion of stock options divided by total shares outstanding; Repurchase (*REP*) is the amount of actual open market repurchases divided by the prior year-end market value of common equity; Time trend dummy variables *D1*, *D2*, *D3*, and *D4* equal one if the observation comes from the 1985–1987, 1988–1990, 1991–1993, and 1994–1996 periods, respectively, and zero otherwise; and all accounting variables except *SIZEEQ*, *SIZEASST*, and *REP* are winsorized at 1% and 99% percentiles.

Panel A. All Variables

Variable	*N*	Mean	Median	Min	Max
FCF	23,208	0.10	0.11	−0.32	0.32
ROAIA	23,208	0.00	0.00	−0.45	0.26
ROE	23,208	0.10	0.12	−1.05	0.55
NONOPI	23,208	0.01	0.01	−0.02	0.08
MKBK	23,208	2.63	1.91	0.51	15.37
NPPE	23,208	0.33	0.28	0.02	0.88
LEVBIA	23,208	0.01	−0.01	−0.42	0.56
LEVIA	23,208	0.02	−0.01	−0.32	0.54
SALEGR	23,208	0.23	0.13	−0.42	2.63
SIZEEQ	23,208	5.35	5.08	0.86	11.81
SIZEASST	23,208	5.38	5.13	1.64	12.49
CASHDIV	23,208	0.16	0.00	0.00	0.87
EXRET	23,208	0.09	−0.01	−0.74	2.73
ITODUM	23,208	0.07	0.00	0.00	1.00
OPTIONS	21,473	0.09	0.07	0.00	0.38
TODUM	23,208	0.06	0.00	0.00	1.00
REP	23,208	0.00	0.00	0.00	0.66

Panel B. Open Market Share Repurchases (*REP*) by Year and Subperiod

Year or Subsample	*N*	Mean (%)	Median (%)
1985	38	7.35	3.83
1986	86	5.52	2.41
1987	97	5.95	3.33
1988	105	6.40	3.47
1989	164	4.53	3.29
1990	282	3.52	2.22
1991	252	3.35	1.31
1992	224	2.55	1.46
1993	213	2.51	1.42
1994	267	2.68	1.58
1995	347	3.31	2.08
1996	469	3.01	1.91
1985–1996	2,544	3.52	1.90
1985–1987	221	6.02	3.06
1988–1990	551	4.37	2.68
1991–1993	689	2.83	1.39
1994–1996	1,083	3.02	1.87

elasticity. She then uses this predicted elasticity in the second stage estimation of the repurchase choice. Here, we follow this two-stage approach by first estimating the ex ante repurchase takeover probability, and then using this predicted probability in the estimation of subsequent repurchase activity.

Specifically, we begin by modeling the latent takeover process (TO_{it}^*) as a linear relation:

$$TO_{it}^* = z_{it-1}\beta_1 + u_i, \quad u_i \sim N(0, \sigma_1^2), \tag{1}$$

where z includes a constant and firm characteristics. Note that TO_{it}^* is unobservable in practice. However, we do observe whether a firm receives a bid or not. Thus, defining the dummy variable *TODUM* as

$$TODUM_{it} = \begin{cases} 1 & \text{if } TO_{it}^* > 0 \\ 0 & \text{otherwise,} \end{cases} \tag{2}$$

we can model the ex ante probability that a firm will become a takeover target using a binary probit model as follows:

$$TOPROB_{it} = Prob(TODUM_{it} = 1) = \Phi(z_{it-1}\beta_1), \tag{3}$$

where $\Phi(.)$ is the cumulative distribution function of a standard normal distribution.

We then model the relation between takeover probability and subsequent repurchase activity. Since the repurchase data are censored in nature, we use a censored regression technique to estimate[14]

$$REP_{it}^* = \gamma TOPROB_{it} + x_{it-1}\beta + \varepsilon_i, \tag{4}$$

where

$$REP_{it} = \begin{cases} REP_{it}^* & \text{if } REP_{it}^* > 0 \\ 0 & \text{otherwise} \end{cases}$$

and x includes a constant, firm characteristics, and time effects that existing studies show affect firms' repurchase decisions. We specify *TOPROB* separately from x to emphasize that we explicitly model the role of the unobservable takeover probability in the repurchase decision equation.

To estimate the above system of equations, we adopt a two-stage instrumental variable approach like that in Hodrick (1999). In the first stage, we estimate the probit model given in equations (1)–(3) to obtain the prerepurchase takeover probability. The dummy variable *TODUM* that we use in equation (2) is the observed event of whether a takeover occurs in year t. Our explanatory variables z_{it-1} in equation (1) consist of the firm characteristics shown by existing studies as determinants of takeovers. These firm characteristics are measured as of the end of year $t-1$ to ensure that we measure the takeover probability

[14] We use two different censored regression techniques, the Tobit and the CQR. We discuss the technical details in Section IV.

prior to the repurchase decision. In other words, if we use firm characteristics that are influenced by repurchase activity (i.e., had we used z_{it} rather than z_{it-1}), the results will likely be biased. Thus, *TOPROB* as estimated by equation (3) measures the ex ante takeover probability perceived *at the beginning of* year t, without influence from the firms' repurchase activities during year t.

In our second step, we estimate equation (4) using the predicted *TOPROB* estimated from the first stage. Our instrumental variables consist of firm characteristics that previous studies document to be determinants of takeover probability, but that do not appear to be proxies of motives for share repurchases. We include these firm characteristics, specifically, fixed assets (*NPPE*) and sales growth (*SALEGR*), in the takeover equation but not the repurchase equation (i.e., these two characteristics serve as instruments).

IV. Estimation Results

We estimate three primary sets of results. First, we estimate the takeover probability using a probit model. Second, we estimate open market repurchase activity using a Tobit model and conduct specification tests. Third, faced with the rejection of the Tobit model, we estimate open market repurchase activity using a CQR technique.

A. Probit Model of Takeover Probability

We estimate a heteroskedasticity-corrected probit model of takeover probability by maximum likelihood (ML).[15] In the probit model, we define a takeover attempt based on whether a firm receives a takeover bid as reported by SDC in a given year t. We use the following explanatory variables, measured as of the end of the prior year (i.e., $t-1$): *ROAIA*, *SIZEEQ*, *LEVBIA*, *MKBK*, *SALEGR*, *NPPE*, and *ITODUM*. We also include time dummy variables. The estimation results are reported in Table II.

Table II reports our estimation results. In general, our results are similar to those reported in previous studies. Namely, we find that takeover probability decreases in firm size (*SIZEEQ*), industry-adjusted leverage (*LEVBIA*), and market-to-book (*MKBK*), and increases in industry-adjusted return on assets (*ROAIA*). Further, the model is quite significant, with a p-value of less than 0.01 on the χ^2 statistics.

[15] Yatchew and Griliches (1984) show that heteroskedastic disturbances in a binary choice model may lead to inconsistent estimates. We adopt the Lagrangian multiplier (LM) test specific to the probit model, as described by Greene (1997, Section 19.4.1), to test for heteroskedasticity of a homoskedastic probit model. We find that the test rejects the null of homoskedastic disturbances. We then reestimate a heteroskedastic probit model in which the disturbances are a function of *ROAIA, SIZEEQ, LEVBIA, MKBK, SALEGR,* and *NPPE* (these variables are found to be related to the form of heteroskedasticity). The test of this heteroskedastic probit suggests the model is well specified.

Table II

Takeover Probability Estimation

This table reports the estimation results of the probability that a firm will become a takeover target, by estimating a probit model. We use the firm's characteristics as of the end of year $t-1$ and the observation of whether a firm becomes a target of a takeover attempt in year t to estimate the firm's takeover probability (*TOPROB*) at the beginning of year t. The sample is from 1985 through 1996. The left-hand side variable equals one if the firm is a takeover target, and zero otherwise; *ROAIA* is operating income before depreciation divided by total assets, minus the median ratio for all firms within the same two-digit SIC code; *SIZEEQ* is the natural log of common equity, which is the number of shares outstanding times the year-end price and price is inflated to 1996 dollars using the CPI; *LEVBIA* is debt divided by assets, minus the median ratio for all firms within the same two-digit SIC code; *NPPE* is net plant, property, and equipment scaled by total assets; *SALEGR* is natural log of the ratio of sales over the sales of the previous year; *NONOPI* is nonoperating income scaled by total assets; *MKBK* is the market value of common stock to the book value of equity; *ITODUM* equals one if at least one firm in the same industry (same four-digit SIC code) is a takeover target in the previous year, and zero otherwise; Time trend dummy variables *D1*, *D2*, *D3*, and *D4* equal one if the observation comes from the 1985–1987, 1988–1990, 1991–1993, and 1994–1996 periods, respectively, and zero otherwise; standard errors are Huber-White quasi-maximum likelihood standard errors; and z-statistics are in parentheses.

Intercept	−1.207
	(−5.63)[c]
ROAIA	1.892
	(2.04)[b]
SIZEEQ	−0.163
	(−1.83)[a]
LEVBIA	−0.813
	(−1.86)[a]
MKBK	−0.379
	(5.58)[c]
SALEGR	−0.014
	(−0.06)
NPPE	0.284
	(0.56)
ITODUM	−0.047
	(−0.53)
D2	0.096
	(1.44)
D3	−0.498
	(−4.87)[c]
D4	0.020
	(0.31)
N	23,208
N (*TODUM* = 1)	1,419
Wald χ^2 (d.f. = 10)	79.62[c]
Log likelihood	−5,236.00

[a,b,c] Significantly different from zero at the 10%, 5%, and 1% level, respectively.

We plot percentile values of the estimated takeover probability in Figure 1. The mean (median) probability is 6.12% (6.28%) and the range is from 0.47% to 16.57%. The 5$^{\text{th}}$ and 95$^{\text{th}}$ percentiles are 2.47% and 9.50%, respectively. Overall, we observe substantial variation in takeover probability.

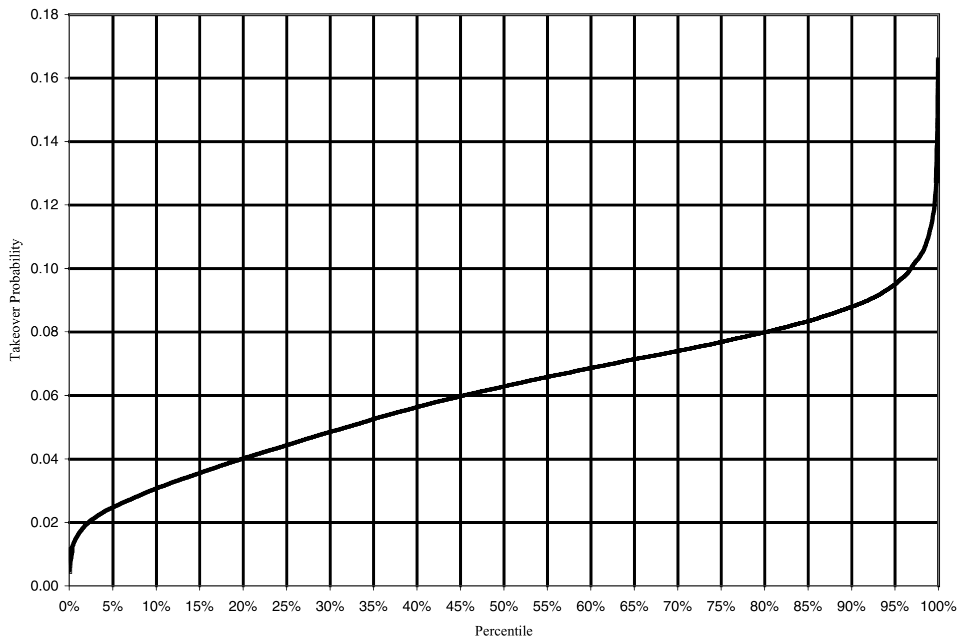

Figure 1. Percentiles of the estimated takeover probability. This figure plots the estimated takeover probability against its associated percentile ranking. The takeover probability is estimated from a probit model.

B. Tobit Model of Open Market Repurchases

We estimate equation (4) as a Tobit model using maximum likelihood (ML). The following subsection compares the results of this model with prior studies. We then examine specification tests of the Tobit model.

B.1. Estimation Results

The estimation results reported in columns (1) and (2) of Table III proxy for the takeover motive using a dummy variable approach (*TARGET*). Following Dittmar (2000), *TARGET* equals one if a bid is received in either the year before or the year of the repurchase. Column (1) excludes the variable *OPTIONS* due to its detrimental impact on sample size; for robustness, column (2) reports the results including *OPTIONS*. Columns (3) and (4) report similar specifications using the estimated takeover probability (*TOPROB*) instead of *TARGET*. Following the open market repurchase literature, we include *FCF, ROE, SIZEASST, LEVIA, NONOPI, MKBK, CASHDIV, EXRET, OPTIONS* as our control variables, and we include the time dummy variables *D2, D3, D4*.

Similar to previous studies, all the estimated parameters are statistically significant at the 1% level in all four specifications with the exception of the

Table III

Repurchase Estimation: Tobit Model

This table reports the estimation results of the factors affecting firms' repurchases using Tobit models. Repurchase (*REP*) is the amount of actual open market repurchases divided by the prior year-end market value of common equity; Target dummy (*TARGET*) equals one if the firm is a takeover target in year t or $t-1$, and zero otherwise; Takeover probability (*TOPROB*) is the firm's estimated takeover probability as of the beginning of year t; *FCF* is operating income before depreciation minus interest expenses, the sum of preferred and common dividends, income taxes (excluding deferred taxes), all scaled by total assets; *ROE* is net income divided by common equity; *SIZEASST* is the natural log of total assets inflated to 1996 dollars using CPI; *LEVIA* is the book value of debt divided by the sum of the book value of debt and the market value of equity, minus the median ratio for all firms within the same two-digit SIC code; *NONOPI* is nonoperating income scaled by total assets; *MKBK* is the market value of common stock to the book value of equity; *CASHDIV* is the ratio of cash dividends paid to net income; *EXRET* is the stock's cumulative return in the previous calendar year minus the value-weighted market return; *OPTIONS* is common shares reserved for conversion of stock options divided by total shares outstanding; Time trend dummy variables *D1*, *D2*, *D3*, and *D4* equal one if the observation comes from the 1985–1987, 1988–1990, 1991–1993, and 1994–1996 periods, respectively, and zero otherwise; and t-statistics are in parentheses.

	(1)	(2)	(3)	(4)
Intercept	−0.204	−0.225	−0.221	−0.241
	(−35.31)[c]	(−34.17)[c]	(−30.80)[c]	(−29.83)[c]
TARGET	0.005	0.004		
	(1.59)	(1.12)		
TOPROB			0.472	0.421
			(4.37)[c]	(3.62)[c]
FCF	0.101	0.104	0.108	0.110
	(6.47)[c]	(6.34)[c]	(6.88)[c]	(6.64)[c]
ROE	0.068	0.077	0.067	0.076
	(7.74)[c]	(8.11)[c]	(7.70)[c]	(8.04)[c]
SIZEASST	0.006	0.008	0.003	0.005
	(9.63)[c]	(11.11)[c]	(3.45)[c]	(5.12)[c]
LEVIA	−0.034	−0.041	−0.053	−0.058
	(−5.41)[c]	(−6.17)[c]	(−6.88)[c]	(−7.08)[c]
NONOPI	0.268	0.288	0.280	0.300
	(3.77)[c]	(3.86)[c]	(3.95)[c]	(4.02)[c]
MKBK	−0.003	−0.004	−0.002	−0.003
	(−5.42)[c]	(−6.39)[c]	(−3.92)[c]	(−4.95)[c]
CASHDIV	0.020	0.028	0.020	0.028
	(4.17)[c]	(5.33)[c]	(4.02)[c]	(5.24)[c]
EXRET	−0.018	−0.019	−0.017	−0.018
	(−8.08)[c]	(−7.92)[c]	(−7.68)[c]	(−7.60)[c]
OPTIONS		0.152		0.153
		(10.37)[c]		(10.47)[c]
D2	0.038	0.036	0.035	0.033
	(11.27)[c]	(10.16)[c]	(10.06)[c]	(9.11)[c]
D3	0.051	0.048	0.064	0.060
	(14.96)[c]	(13.52)[c]	(13.91)[c]	(12.28)[c]
D4	0.058	0.055	0.058	0.055
	(17.67)[c]	(15.73)[c]	(17.70)[c]	(15.73)[c]
N	23,208	21,473	23,208	21,473
N(*REP* > 0)	2,544	2,283	2,544	2,283
Log likelihood	−1,888.58	−1,698.37	−1,880.45	−1,692.55
LR χ^2 (df = 12)	1,009.09[c]	1,042.57[c]	1,025.34[c]	1,054.20[c]

[a,b,c] Significantly different from zero at the 10%, 5%, and 1% level, respectively.

coefficient on *TARGET*.[16] In particular, repurchase activity increases in *FCF, ROE, SIZEASST, NONOPI, CASHDIV,* and *OPTIONS,* and decreases in *LEVIA, MKBK,* and *EXRET*. However, while the above results are consistent with prior empirical studies, a number of these results are inconsistent with theoretical predictions. For example, we would expect *SIZEASST* to be negatively related to *REP*, given firm size is a proxy for asymmetric information, and small firms are considered more opaque. Further, while Jagannathan, Stephens, and Weisbach (2000) argue the flexibility of open market repurchases make them a useful mechanism to pay out temporary rather than permanent cash flow, we find a positive coefficient for both *FCF* (a measure of permanent cash flow) and *NONOPI* (a measure of temporary cash flow), rather than simply on *NONOPI*. We also would expect negative coefficients on our time dummy variables, which are intended to capture differences between repurchase activity in the 1985 to 1987 period and that in later subperiods: Panel B of Table I shows that repurchase activity in the late 1980s to mid-1990s is below that in the mid-1980s, predicting a negative sign on *D2*, *D3*, and *D4*. However, we find positive and statistically significant coefficients on these time indicator variables.

With respect to the takeover deterrent results, the estimations using the takeover dummy variable in columns (1) and (2) replicate those of Dittmar (2000) reasonably well. Dittmar (2000) finds a statistically significant coefficient on *TARGET* in some years, with an average coefficient of 0.02.[17] The magnitude of the estimated coefficients leads her to conclude that takeover deterrence plays only a "minor" role. We find a statistically insignificant coefficient of *TARGET* equal to 0.005. Economically, the coefficient on *TARGET* suggests that a repurchasing firm with a takeover deterrent motive will repurchase an additional 0.5% of shares outstanding. The results are similar in column (2) when we include *OPTIONS*.

In columns (3) and (4), we use a latent variable to capture the takeover probability *TOPROB*. The results for (3) and (4) are similar. In column (3) the coefficient on *TOPROB* is 0.472 with a t-statistic of 4.37. A firm with a takeover probability of 2.47% (the 5th percentile value of *TOPROB*) will repurchase 1.17% of its shares, while a firm with a takeover probability of 9.50% (the 95th percentile of *TOPROB*) will repurchase 4.48% of its shares. Thus, firms with a relatively high takeover probability will repurchase an additional 3.31% of outstanding shares. There are a number of ways to interpret the economic significance of this 3.31% figure. On the one hand, the average value of *REP* for all repurchasers is 3.52%, making a deviation of 3.31% seem relatively large. On the other hand, from the perspective of the takeover market, it is not clear whether repurchasing an additional 3.31% of shares would serve as an economically meaningful deterrent. What is clear, however, is that from either the repurchase or the takeover perspective, the latent variable approach leads to more economically meaningful results compared to the dummy variable approach.

[16] The results of marginal effects are similar to those reported in Table III.

[17] Dittmar (2000) reports estimation results for each year in her sample. We compute the sample-size weighted average of her coefficients from 1985 to 1996.

B.2. Specification Tests

It is well known in the literature (e.g., Greene (1997)) that the parametric censored linear regression model (Tobit) may render inconsistent estimates in the presence of either heteroskedasticity or non-normality in the error terms. Following Petersen and Waldman (1981), we conduct a likelihood ratio (LR) test of the null hypothesis of homoskedasticity for the Tobit models presented in Table III.[18] The test results show that the null hypothesis is rejected for all four specifications, suggesting our results in Table III may be inconsistent. We then conduct the conditional moment test of Pagan and Vella (1989) to test the null hypothesis of normally distributed disturbances. We find that the hypothesis of normally distributed errors is also soundly rejected for all four specifications.

To address these concerns, we estimate a heteroskedastic Tobit model. Unfortunately, even the heteroskedastic Tobit model can not pass either of the two specification tests. We also find that heteroskedasticity has a large effect on the estimates. To highlight the magnitude of this effect, we estimate a Tobit model specifying generalized linear model (GLM) standard errors.[19] We find that the log likelihood value is six times that of the homoskedastic and heteroskedastic Tobit estimations. In sum, prior studies would have likely documented a significant takeover deterrent effect had they used a latent variable for takeover probability. However, the Tobit model does not seem to be appropriate for repurchase data.[20] As a result, the estimates are not reliable, making economic inferences from these estimates suspect.

C. Censored Quantile Regression (CQR) Model of Open Market Repurchases

Given the specification concerns of the Tobit model, we apply the CQR estimator proposed by Powell (1986). This model only requires the distribution of the random component to satisfy regular conditions, allowing the distribution of the random component to be unknown. Powell (1986) shows that this estimator is consistent and asymptotically normal for a wide class of error distributions, and it is robust to heteroskedasticity.

The CQR model we estimate is:

$$REP_{it}^{*} = \gamma_{\theta} TOPROB_{it} + x_{it-1}\beta_{\theta} + \varepsilon_{\theta i}, \quad Quant_{\theta}(\varepsilon_{\theta i} \mid TOPROB_{i}, x_{i}) = 0, \tag{5}$$

where

$$REP_{it} = \begin{cases} REP_{it}^{*} & \text{if } REP_{it}^{*} > 0 \\ 0 & \text{otherwise,} \end{cases}$$

[18] Details on the specification tests of the Tobit model are available from the authors. Also, see Drukker (2002)

[19] We use the GLM standard errors to illustrate the dramatic effect of heteroskedasticity on the estimation results. However, imposing GLM standard errors has its own concerns since we lack theory to justify such a specification for our case.

[20] We also find that specification tests reject the Tobit model when we use data that are not winsorized. Thus, the error distribution issues in our data that lead to the rejection of the Tobit are not caused or alleviated by winsorizing the data.

and $Quant_\theta(\varepsilon_{\theta i} \mid .)$ is the θ^{th} conditional quantile of the error. Following Powell (1986), we estimate the CQR model by calculating the first-order conditions and the Hessian matrix. These details are presented in the Appendix.

Bandwidth selection is a critical issue in the nonparametric and semiparametric literature. We are unaware of any fully automatic bandwidth selection method for Powell's CQR model. Horowitz and Neumann (1987) apply Powell's CQR method to duration data, and Melenberg and Van Soest (1996) apply it to vacation expenditure data. Both articles find that the standard errors, and therefore inferences, are sensitive to bandwidth selection. Therefore, we estimate models with bandwidths of 0.10, 0.25, 0.50, 0.75, and 1.00 to verify our results' robustness.

We must also choose data quantiles for our estimation. If the dependent variable is highly censored, choosing low quantiles of the dependent variable will be uninformative about β. Intuitively, the censored observations are less informative than the uncensored observations, and thus, centering the distribution of the dependent variable at a high quantile, which increases the influence of uncensored observations, leads to more informative estimates. About 90% of our observations are censored. This implies, therefore, that quantiles below 90% will yield less information about the true coefficients than quantiles above 90%. However, while taking high quantiles helps ensure that coefficient parameter estimates are informative, such quantiles can be estimated less precisely. Given this trade-off, we choose the 95% quantile for our estimation. For robustness, we also run our estimation with the 90% quantile and compare the resulting parameters with those from the 95% quantile. If the 95% quantile is too high, then we would expect the results from the 90% quantile to be very different.

C.1. Estimation Results

We report the CQR results censored at the 95% quantile for specifications with and without *OPTIONS* in Table IV. Overall, the estimation results are similar regardless of whether we include *OPTIONS*. In addition, the standard errors are generally comparable across bandwidths, suggesting that inferences are consistent across bandwidths for the majority of the variables. The estimates from the 90% quantile are also similar to those of the 95% quantile, suggesting the 95% quantile is appropriate.[21]

Table IV indicates that the CQR coefficient on takeover probability (*TOPROB*) is positive and significant at the 1% level for all bandwidths. This coefficient is more than four times as large as that from the Tobit estimation.[22] This result is also economically significant: A firm moving from the 5^{th} percentile value of takeover probability to the 95^{th} percentile will repurchase an additional 15.58% of outstanding shares.[23] This suggests the takeover deterrent

[21] The results for the 90% quantile are available from the authors.

[22] The difference between the coefficients is significant at the 1% level.

[23] The 15.58% figure is equal to the coefficient on *TOPROB* multiplied by the difference between the values of *TOPROB* at the 5^{th} and 95^{th} percentiles: $2.216\ (0.0950 - 0.0247) = 0.1558$.

Table IV

Repurchase Estimation: Censored Quantile Regression (CQR) Model

This table reports the results of estimating Powell's (1986) CQR model for repurchases. The model is estimated semiparametrically using the 95% censored quantile for bandwidths of 0.10, 0.25, 0.50, 0.75, and 1.00. Repurchase (*REP*) is the amount of actual open market repurchases divided by the prior year-end market value of common equity; Takeover probability (*TOPROB*) is the firm's estimated takeover probability (*TOPROB*) as of the beginning of year *t*; *FCF* is operating income before depreciation minus interest expenses, the sum of preferred and common dividends, income taxes (excluding deferred taxes), all scaled by total assets; *ROE* is net income divided by common equity; *SIZEASST* is the natural log of total assets inflated to 1996 dollars using CPI; *LEVIA* is the book value of debt divided by the sum of the book value of debt and the market value of equity, minus the median ratio for all firms within the same two-digit SIC code; *NONOPI* is nonoperating income scaled by total assets; *MKBK* is the market value of common stock to the book value of equity; *CASHDIV* is the ratio of cash dividends paid to net income; *EXRET* is the stock's cumulative return in the previous calendar year minus the value-weighted market return; *OPTIONS* is common shares reserved for conversion of stock options divided by total shares outstanding; Time trend dummy variables *D1*, *D2*, *D3* and *D4* equal one if the observation comes from the 1985–1987, 1988–1990, 1991–1993, and 1994–1996 periods, respectively, and zero otherwise.

		t-Statistics by Bandwidth				
	Parameter	0.10	0.25	0.50	0.75	1.00
Intercept	0.173	8.58[c]	21.65[c]	10.97[c]	7.31[c]	5.48[c]
TOPROB	2.216	16.04[c]	16.54[c]	8.41[c]	5.60[c]	4.20[c]
FCF	0.009	0.65	0.47	0.24	0.16	0.12
ROE	0.029	4.40[c]	2.83[c]	1.42	0.95	0.71
SIZEASST	−0.018	−15.11[c]	−16.05[c]	−8.20[c]	−5.46[c]	−4.10[c]
LEVIA	−0.046	−5.10[c]	−4.89[c]	−2.48[b]	−1.65[a]	−1.24
NONOPI	0.446	5.16[c]	5.14[c]	2.64[c]	1.76[a]	1.32
MKBK	−0.003	−7.69[c]	−5.27[c]	−2.68[c]	−1.79[a]	−1.34
CASHDIV	0.001	0.12	0.10	0.05	0.03	0.02
EXRET	−0.002	−0.93	−0.65	−0.33	−0.22	−0.16
D2	−0.091	−4.72[c]	−19.58[c]	−10.26[c]	−6.84[c]	−5.13[c]
D3	−0.053	−2.65[c]	−9.08[c]	−4.64[c]	−3.09[c]	−2.32[b]
D4	−0.097	−5.01[c]	−22.36[c]	−11.72[c]	−7.81[c]	−5.86[c]
		t-Statistics by Bandwidth				
	Parameter	0.10	0.25	0.50	0.75	1.00
Intercept	0.173	6.95[c]	19.90[c]	10.09[c]	6.73[c]	5.05[c]
TOPROB	2.078	15.27[c]	14.64[c]	7.40[c]	4.93[c]	3.70[c]
FCF	−0.045	−3.04[c]	−2.29[b]	−1.15	−0.77	−0.58
ROE	0.025	3.38[c]	2.30[b]	1.15	0.77	0.58
SIZEASST	−0.017	−13.69[c]	−13.89[c]	−7.04[c]	−4.69[c]	−3.52[c]
LEVIA	−0.055	−6.37[c]	−5.59[c]	−2.84[c]	−1.89[a]	−1.42
NONOPI	0.512	5.36[c]	5.66[c]	2.93[c]	1.95	1.47
MKBK	−0.004	−7.73[c]	−4.89[c]	−2.45[c]	−1.63	−1.23
CASHDIV	0.002	0.42	0.33	0.17	0.11	0.08
EXRET	0.001	0.55	0.41	0.21	0.14	0.10
OPTIONS	0.072	4.83[c]	4.22[c]	2.12	1.41	1.06
D2	−0.094	−3.86[c]	−19.58[c]	−10.10[c]	−6.73[c]	−5.05[c]
D3	−0.064	−2.55[c]	−10.38[c]	−5.29[c]	−3.52[c]	−2.64[c]
D4	−0.100	−4.09[c]	−22.20[c]	−11.44[c]	−7.63[c]	−5.72[c]

[a,b,c] Significantly different from zero at the 10%, 5%, and 1% level, respectively.

effect of open market repurchases is much more economically meaningful than previous studies document.

Comparing the CQR estimation results to those from the Tobit estimation in Table III, we see that a number of parameter estimates differ in both sign and significance across the two models. With respect to cash flows, free cash flow (*FCF*) is significant at the 1% level in the Tobit estimation, but the coefficient is much smaller and not statistically significant in the CQR. While the coefficient of nonoperating cash flow (*NONOPI*) remains statistically significant for most bandwidths, the point estimate from the CQR estimation is much larger. Recall that *FCF* is a proxy for permanent cash flows, whereas *NONOPI* measures temporary cash flows. In contrast to the Tobit results, the CQR results distinguish between these two variables and give a more sensible interpretation. In particular, these results support the empirical evidence presented by Jagannathan, Stephens, and Weisbach (2000), who find that firms use open market repurchases to pay out "temporary" nonoperating cash flows, while dividends represent an ongoing commitment and are used to distribute "permanent" operating cash flows.

Similarly, the coefficient on *CASHDIV* is positive and significant in the Tobit estimation, suggesting a complementary relationship between dividends and repurchases, whereas theory typically suggests that dividends and repurchases are substitutes. In the CQR estimation, the coefficient of *CASHDIV* is not significant, suggesting that repurchases and dividends are neither complements nor substitutes, but rather uncorrelated.

The coefficient on *SIZEASST* is significantly positive in the Tobit estimation, but negative and significant (at the 1% level) in the CQR model. The latter result makes intuitive sense. Firm size is commonly used as a proxy for information asymmetry in empirical studies: The bigger the firm, the less likely it will be misvalued. In this sense, a positive estimate of *SIZEASST* contradicts the well-documented result that the positive abnormal repurchase announcement returns are statistically negatively correlated with firm size (Ikenberry, Lakonishok, and Vermaelen (1995)). Note that both Dittmar (2000) and Fenn and Liang (2001) find a positive and significant coefficient on size in their Tobit regressions; Dittmar (2000) points out that the Tobit's positive coefficient on *SIZEASST* is puzzling.

The estimates of the three time period dummy variables (*D2*, *D3*, and *D4*) change from significantly positive in the Tobit estimation to significantly negative in the CQR. While the aggregate amount of open market repurchases has generally increased since the mid-1980s, Allen and Michaely's (2003) Figure 1 shows that the repurchase percentage, that is, the repurchase amount scaled by the market value of equity, is higher in the 1985 to 1987 subperiod (i.e., our *D1*) than in any other subperiod over 1972 to 1998. This pattern is consistent with our results reported in Panel B of Table I, where both the mean and the median of the repurchase percentage *REP* are higher in the 1985 to 1987 subperiod than in any of the other three subperiods. The negative coefficients on *D2*, *D3*, and *D4* in the CQR results are consistent with this pattern.

In summary, the Tobit model leads to inconsistent estimates. The CQR model, together with a semiparametric estimation, gives more sensible interpretations. Given the censored data problem is often encountered in empirical work, our results show that one needs to be careful in using the Tobit model.

C.2. Specification Tests

Powell (1986) shows that the CQR model generates consistent estimates under the assumption that the errors (with unknown distribution) are independent of the regressors. If the assumption of homoskedastic errors is violated, the CQR estimator is still consistent under the hypothesis of conditional symmetry of the error distribution. We test for homoskedasticity of the random errors in the CQR estimation. Powell (1986) suggests using a test with a histogram-type density estimator; and Horowitz and Neumann (1987) offer a modified version of this test. We adopt the χ^2 statistic of Horowitz and Neumann (1987). The details of the test statistic are presented in the Appendix. Intuitively, the test statistic is based on the difference between the estimated slope parameters using different quantiles. If the difference is large, then the null hypothesis of homoskedastic errors is rejected.

As we discuss above, estimation with low quantiles will be uninformative about the parameters since our repurchase data are highly censored. We therefore form the test statistic based on the difference between the 90% and 95% quantile estimates. The test results, reported in Table V, show that we cannot reject the null hypothesis of homoskedasticity. This suggests that the CQR model is appropriate for our sample.

V. Conclusions

We take a new approach to examine the takeover deterrent role of open market share repurchase. Existing studies investigate this question by looking at whether corporate control events are followed by repurchases. However, if a repurchase deters a takeover attempt from occurring in the first place, then this methodology will fail to show the effectiveness of open market repurchases in deterring takeovers.

To capture this deterrent effect, we model takeover probability as a latent variable and examine how the perceived threat affects the firm's decision to repurchase. We find a statistically and economically large impact. In particular, firms' repurchase activity increases when they face a high takeover probability. We believe this result sheds new light on the takeover deterrent motivation for open market share repurchases, extending the literature on other well recognized motivations for share repurchases.

We also contribute to the literature by examining the appropriateness of the use of Tobit models in explaining open market repurchase activity. We find that the Tobit model suffers from heteroskedasticity and non-normality in the error term, resulting in unreliable estimates. We therefore adopt Powell's CQR method, together with semiparametric estimation, as in theory, this estimator

Table V

Repurchase Estimation: Homoskedasticity Test of the Censored Quantile Regression (CQR) Model

This table reports the homoskedasticity test of the CQR model. Let z be a vector of J ($J \geq 1$) distinct quantile estimators, that is,

$$z_N = [\beta_N(\theta_1), \ldots, \beta_N(\theta_J)]',$$

and let ζ be the corresponding vector of true regression quantiles. Powell (1986) shows that

$$N^{\frac{1}{2}}(z_N - \zeta) \overset{A}{\sim} N(0, V_Q),$$

where the asymptotical covariance matrix can be written in partitioned form as follows:

$$V_Q = \left[\omega_{jk} D_l^{-1}\right], \quad l = \max(j,k), j,k = 1, \ldots, J$$

$$\omega_{jk} \equiv \frac{\min(\theta_j, \theta_k) - \theta_j, \theta_k}{f(u_{\theta_j}) f(u_{\theta_k})},$$

$$D_l \equiv E\left\{I\left(x'\hat{\beta}(\theta_l) > 0\right) x'x\right\}.$$

The test statistic follows Horowitz and Neumann (1987). Let T be the matrix of zeros and ones that transform z_N into the vector of differences of the nonintercept coefficients

$$Tz_N = \left[\alpha_N^*(\theta_1) - \alpha_N^*(\theta_2), \ldots, \alpha_N^*(\theta_1) - \alpha_N^*(\theta_J)\right]'.$$

We then have the χ^2 test statistic

$$z_N' T' V_z^{-1} T z_N \overset{A}{\sim} \chi^2(rank\,(V_z)), \text{ where } V_z = TV_Q T'.$$

We report the test results by using the 95% and 90% quantiles with bandwidths of 0.10, 0.25, 0.50, 0.75, and 1.00.

	Bandwidth				
	0.10	0.25	0.50	0.75	1.00
χ^2 statistics (d.f. = 12)*	13.33	12.11	3.12	1.38	0.78

*The one-sided 10%, 5%, and 1% χ^2 (d.f. = 12) critical values are 18.55, 21.03, and 26.22.

is consistent under regular conditions. We find the Tobit and CQR yield similar conclusions with regard to takeover deterrence. However, the CQR method generates a number of new results that contradict both our Tobit results and the Tobit results of previous studies. Specifically, the CQR model supports the asymmetric information hypothesis and the motive of distributing temporary but not permanent cash flows. Our results highlight the need to verify the specification tests of the Tobit model in analyzing censored data in corporate finance.

Appendix: Computational Details of Censored Quantile Regression (CQR)

The censored quantile regression (CQR) model can be written as:

$$y_i^* = x_i'\beta_\theta + u_{\theta i}, \quad Quant_\theta(u_{\theta i} \mid x_i) = 0, \tag{A1}$$

where

$$y_i = \begin{cases} y_i^* & \text{if } y_i^* > 0 \\ 0 & \text{otherwise,} \end{cases}$$

and $Quant_\theta(u_{\theta i} \mid x_i)$ is the θ^{th} conditional quantile of the error. Following Powell (1986), we estimate the coefficients β_θ as a solution to

$$\arg\min_\beta \frac{1}{N}\sum_{i=1}^{N} \rho_\theta(y_i - \max(x_i'\beta, 0)), \tag{A2}$$

where $\rho_\theta(\lambda) = (\theta - I(\lambda < 0))^* \lambda$ and $I(.)$ is the usual indicator function.

Then, the first-order condition is

$$\frac{1}{N}\sum_{i=1}^{N} I(x_i'\hat{\beta}_\theta > 0)(\theta - 0.5 + 0.5\text{sign}(y_i - x_i'\hat{\beta}_\theta))x_i. \tag{A3}$$

Note that in estimating β_θ only the observations for which $x_i'\hat{\beta}_\theta > 0$ are used.

Following Powell (1986), it can be shown that

$$\sqrt{N}(\hat{\beta}_\theta - \beta_\theta) \xrightarrow{D} N\big(0, \theta(1-\theta)A^{-1}E\{xx'I(x'\beta_\theta > 0)\}A^{-1}\big), \tag{A4}$$

where $A \equiv E\{f_{u_\theta}(0 \mid x)I(x'\beta_\theta > 0)xx'\}$ and $f_{u_\theta}(0 \mid x)$ denotes the density of u_θ given x. If we assume u is independent of x, then $f_{u_\theta}(0 \mid x) = f_{u_\theta}(0)$. Therefore, the variance simplifies to

$$\frac{\theta(1-\theta)}{f_{u_\theta}^2(0)}(E\{xx'I(x'\beta_\theta > 0)\})^{-1} \tag{A5}$$

and

$$Quant_\theta(y^* \mid x_i) = (\beta_0 + Quant_\theta(u)) + x_1\beta_1 + \cdots + x_k\beta_k. \tag{A6}$$

The intercept term β_0 cannot be identified, and thus cannot be estimated consistently. The other parameters are identifiable for any quantile θ under the assumption that u is independent from x, while this is valid only for the θ satisfying $Quant_\theta(u \mid x) = 0$ if independence does not hold (in this case, one needs to resort to the hypothesis of conditional symmetry of the error distribution; see Powell (1986) for details).

A. Nonparametric Estimation of the Density Function

Density $f_{u_\theta}(0)$ is estimated nonparametrically. We choose the trimmed quadratic kernel which has bounded support

$$bs[j] = \frac{1}{N}\left(\prod_{i=1}^{k}\frac{1}{h\sigma_i}\sum_{l=1}^{N}\frac{3}{4}\left[1-\frac{(x_i - x_i[j])^2}{(h\sigma_i)^2}\right]I_i(.)\right)(I(y_j > 0) - I(y_j \leq 0)), \tag{A7}$$

where j stands for the j^{th} observation, N is the number of total observations, k is the number of slope parameters to be estimated (excluding the intercept as it is not identifiable), x_i is the i^{th} explanatory variable (again, not including the intercept term), $x_i[j]$ is its j^{th} observation, h stands for bandwidth, σ_i is the standard deviation of the i^{th} variable x_i, $I(.)$ is the usual indicator function, and

$$I_i(.) = I\left(\frac{\text{abs}\left(x_i - x_i[j]\right)}{h\sigma_i} < 1\right), \tag{A8}$$

where abs(.) means absolute value.

Theory tells us that the optimal bandwidth is $cN^{-0.2}$, where c is some constant and N is the number of observations. For our sample $N = 23{,}208$ and we choose $c = 4$. We therefore have $cN^{-0.2} = 0.5$.

B. Testing for Heteroskedasticity in the CQR Model

Let z be a vector of J ($J \geq 1$) distinct quantile estimators, that is,

$$z_N = [\beta_N(\theta_1), \ldots, \beta_N(\theta_J)]', \tag{A9}$$

and let ζ be the corresponding vector of true regression quantiles. Powell (1986) shows that

$$N^{\frac{1}{2}}(z_N - \zeta) \xrightarrow{A} N(0, V_Q), \tag{A10}$$

where the asymptotical covariance matrix can be written in partitioned form as

$$V_Q = \left[\omega_{jk} D_l^{-1}\right], \quad l = \max(j,k), \quad j,k = 1,\ldots,J \tag{A11}$$

$$\omega_{jk} \equiv \frac{\min(\theta_j,\theta_k) - \theta_j,\theta_k}{f(u_{\theta_j})f(u_{\theta_k})} \tag{A12}$$

$$D_l \equiv E\{I(x'\hat{\beta}(\theta_l) > 0)x'x\}, \tag{A13}$$

where $f(.)$ stands for the probability density function. Powell (1986) suggests using a histogram-type density estimator for $f(.)$:

$$f_N(u_\theta) = \left[\sum_{i=1}^{N} I(x_i'\hat{\beta}(\theta) > 0)\right]^{-1} h_N^{-1}\left[\sum_{i=1}^{N} I(x_i'\hat{\beta}(\theta) > 0) I(0 \leq \hat{u}(\theta) \leq h_N)\right]. \tag{A14}$$

That is, the density function is estimated by the fraction of residuals (with positive regression functions) in the interval $[0, h_N]$ divided by the width of this interval.

We follow Horowitz and Neumann (1987) to test for homoskedasticity. Let T be the matrix of zeros and ones that transform z_N into the vector of differences of the nonintercept coefficients:

$$Tz_N = \left[\beta_N^*(\theta_1) - \beta_N^*(\theta_2), \ldots, \beta_N^*(\theta_1) - \beta_N^*(\theta_J)\right]', \tag{A15}$$

then we have χ^2 test statistic:

$$z_N' T' V_z^{-1} T z_N \xrightarrow{A} \chi^2(\text{rank}(V_z)), \quad \text{where} \quad V_z = T V_Q T'. \tag{A16}$$

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