Employees as Creditors: The Disciplinary Role of Pension Deficits in the Market for Corporate Control

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Abstract: This paper examines the disciplinary role of corporate pension deficits in the market for corporate control. We find that companies with larger pension deficits are less likely to engage in diversifying mergers, experience higher merger announcement returns, pay lower premiums to targets, and use a higher percentage of cash in their payment. These results are more evident for acquirers with pension plans that are dominated by actively working employees or collectively bargained by employees. Our findings indicate that corporate pension deficits provide employees with strong incentives to monitor managerial performance and influence managers to make value-enhancing investment decisions.

Keywords: Pension deficits; Market for corporate control; Mergers and acquisitions; Merger announcement returns; Disciplinary role

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"Employee engagement levels are linked to perceptions of a company's leadership and, more specifically, the extent to which members of the workforce believe senior management is committed to their well-being. Leading organizations understand that retirement programs and, more specifically, benefit security, can play a role in favorably influencing employee perceptions of senior management and the overall organization."

Towers Perrin (2009) "CFO guidebook: pensions and corporate financial performance - intricately linked"

1. Introduction

The extent of debt in a firm's capital structure can influence managerial decisions. For example, Jensen (1986) argues that while managers with large free cash flow have incentives to over-invest beyond the optimal level, high debt serves as a controlling mechanism that prevents managers from wasting free cash. Consistent with this argument, Kang(1993)shows that a bidder's announcement returns are positively related to the bidder's leverage. In addition, Berger, Ofek, and Yermack (1997) show that firms with entrenched managers maintain lower leverage. The banking literature further shows that the disciplinary role of debt is more evident for bank loans than for public debt (e.g., Diamond (1984), Fama (1985)).[1]

Although the studies above enhance our understanding of the disciplinary role of conventional debt, and of bank debt in particular, in corporate decisions, little is known about the disciplinary role of a firm's debt to its employees through its defined benefit (DB hereafter) pension plan.[2] This lack of evidence is surprising, given that the combined pension deficits (difference between the value of the projected pension benefit obligation and the fair value of plan assets) of all U.S. publicly listed companies reached $707 billion at the end of 2008.[3] To the extent that pension deficits create long-term unsecured debtors out of employees (Ippolito (1985a, 1985b)), employees thus can be considered as one of the most important debtholders when DB sponsoring firms have large pension deficits.[4] In this paper we focus on the role of employees as informed large unsecured debtholders and examine how corporate pension deficits owed to employees, a type of “inside” debt, influence managerial incentives to make value-enhancing investment decisions. Specifically, using a large sample of mergers over the 1981 to 2008 period, we examine whether pension deficits affect a firm’s decision to undertake diversifying mergers and how pension deficits affect acquirer announcement returns, value-weighted portfolio returns of the acquirer and the target, premiums paid to targets, and the choice of payment methods.

We argue that pension deficits are likely to serve as an important control mechanism that limits managers’ discretionary power, for several reasons. First, underfunded pension plans have an adverse effect on welfare and job security of employees. Munnell, Aubry, and Muldoon (2008) argue that large pension deficits force sponsoring firms to go bankrupt, lay off their employees, or freeze/terminate their DB pension plans, which would result in fewer retirement benefits for their employees than those anticipated. Thus, to minimize the adverse effects of pension deficits on their welfare, employees are expected to have strong incentives to monitor sponsoring firms’ investment decisions. Second, pension deficits account for a significant portion of total debt. Moreover, firms with large pension deficits are required by law to make periodic mandatory contributions and pay a high insurance premium to the Pension Benefit Guaranty Corporation (hereafter, PBGC). These characteristics of pension deficits help reduce the free cash flow available to managers. Third, as insiders, employees are expected to possess material private information about their employers, which gives them a competitive advantage in collecting information and thus enhanced capabilities to monitor managers. Finally, previous studies show that pension deficits increase employees’ perceived risk of sponsoring firms (Ippolito (1997)), reduce employee incentives to work hard (Hanka (1998)), increase the cost of capital (Cardinale (2007), Campbell, Dhaliwal, and Schwartz (2011)), and lower the credit rating of sponsoring firms (Rauh (2007)). As a result, pension deficits can incentivize managers to make value-enhancing investment decisions because value-destroying investments will only further exacerbate the negative effect of pension deficits on employee behavior and firm value (Ippolito (1997)). Taken together, these arguments suggest that pension deficits can serve as an effective control mechanism that mitigates managerial incentive problems.

Controlling for conventional leverage ratio and correcting for the nonrandom choice of whether to sponsor a DB plan and the endogeneity of the size of pension deficits, we find results that are largely consistent with the view that pension deficits serve as an important disciplinary mechanism. Specifically, we find that conditional on engaging in acquisitions, firms with larger pension deficits are less likely to make diversifying mergers: a one-standard deviation increase in the ratio of pension deficits to total assets is associated with a 3.4% decrease in the likelihood of making diversifying mergers. Moreover, bidder announcement returns and announcement returns of the value-weighted portfolio of the bidder and the target increase significantly with a bidder’s pension deficits. Economically, a one-standard deviation increase in a bidder’s pension deficits over total assets translates into a 0.34% increase in the bidder’s announcement returns over the three-day event window around the announcement date and a 0.91% increase in the combined portfolio’s announcement returns over the same event window. Thus, the disciplinary effect of pension deficits appears to be both statistically and economically significant. Finally, we find that acquirers' pension deficits are negatively associated with the premiums paid to targets but positively related to the percentage of cash used in payment to the targets: a one- standard deviation increase in an acquirer's pension deficits over total assets results in a 2.9% decrease in takeover premium and a 16.3% increase in the fraction of cash payment. These results suggest that pension deficits not only limit managerial hubris in overpaying for targets (Roll (1986)) but also influence managers to choose undervalued targets, thereby creating higher value gains in mergers (Shleifer and Vishny (2003), Dong et al. (2006)).

Although we find that the estimated coefficients on conventional leverage are also significant and have the same signs as those on pension deficits in most regressions, the economic impact of pension deficits is more pronounced than that of conventional debt, suggesting that pension deficits impose more binding constraints than conventional debt in disciplining managers.

We conduct a battery of additional tests to ensure that our results are robust to alternative empirical specifications and variable definitions. In addition, we examine several alternative arguments that might explain our results. For example, under a financial constraint argument, financially constrained firms (e.g., firms with pension deficits) undertake only profitable investments due to lack of their internal funds. It is also possible that the cash drain caused by mandatory pension contributions influences managers of firms with pension deficits to engage in more profitable mergers. We test these alternative explanations in the robustness section and find little evidence to support them. We also examine whether our main results are driven by post-acquisition pension fund revisions or synergistic gains created by merging the acquirers with an underfunded plan with the targets with an overfunded plan and find little support for these alternative explanations.

Finally, to further provide evidence on the disciplinary role of pension deficits and to exploit the cross-sectional heterogeneity across firms, we conduct several subsample analyses, which allows us to address the potential omitted variable problems effectively. In particular, we investigate how demographic characteristics of a pension plan influence the control function of pension deficits. We find that our results are mainly driven by subsamples of firms with a shorter plan age and those with a higher fraction of actively working employees (i.e., fewer retired employees). To the extent that young or current employees have stronger monitoring incentives than old or retired employees, these results provide further evidence in support of the control function of pension deficits. We also examine how employees' collective efforts to monitor managers affect our main results by partitioning our sample firms according to the collective bargaining status of a pension plan and the sample median of the industry unionization rate, respectively. Since collective bargaining agreement and labor unions give employees a collective voice through which their actions can be better coordinated and thus enhance their collective actions (Chen, Kacperczyk, and Ortiz-Molina (2010), Hochberg and Lindsey (2010), Comprix and Muller (2011)), we expect the control function of pension deficits to be more effective for firms whose pension plans are collectively bargained or those in more unionized industries. Consistent with this view, we find that the impact of pension deficits on several aspects of takeover decisions is more pronounced for such firms. In addition, using the pension law change in 1987 as a quasi-exogenous event, which requires higher contributions to severely underfunded pension plan, we find that the effect of pension deficits on acquisition outcomes is more statistically significant in the post-1987 period than in the pre-1987 period.

Our work is related to several recent studies that examine the impact of pension funding on corporate investments. For example, by exploiting nonlinear funding rules for DB pension plans, Rauh (2006a) documents a causal negative effect of mandatory pension contributions on corporate investments. Bergstresser, Desai, and Rauh (2006) argue that managers adjust assumed rates of return on pension assets in managing earnings and show that managers adjust these rates of return more aggressively prior to acquiring other firms. Finally, Franzoni (2009) investigates stock price reactions to the payment of mandatory pension contributions to a firm's DB pension plan. He finds that the price decline is more severe for financially constrained firms and interprets this result as evidence of a negative effect of financing frictions on investment.

Although our paper also examines the effect of pension funding status on corporate investments, it is distinct from prior studies in at least two important ways. First, while previous papers explore a firm's underinvestment in capital expenditures in the presence of financial constraints, highlighting the costs of raising external funds, our paper focuses on the disciplinary role of pension deficits by investigating how pension deficits serve as a control mechanism that limits managers' ability to engage in overinvestment and incentivizes managers to improve investment quality. We examine this issue by using an integrated approach that jointly examines several outcomes of takeover decisions, such as the frequency of diversifying mergers, their quality (i.e., valuation effects of mergers), and the choice of payment methods (i.e., cash versus stock payments). This approach, particularly using the merger announcement returns in the analysis, allows us to determine how pension deficits affect the quality of corporate investments, which is difficult to address using routine capital expenditure investments. Second, unlike prior papers that focus on capital expenditure decisions, we use takeover bids as our experimental setting, which provide a natural experiment for providing evidence on the disciplinary role of pension deficits because takeover decisions typically represent large and discrete investment choices. Furthermore, because managers can pursue private objectives at the expense of shareholder wealth during acquisitions (Jensen and Ruback (1983)), informed stakeholders such as large shareholders, creditors, and workers can each have an important impact on managerial decisions with respect to takeover bids. In addition, since mergers and acquisitions (M&A) frequently involve large-scale employee restructuring for cost savings and lead to the changes in the definition of employee jobs (Pagano and Volpin (2005), Rauh (2006b)), employees are more likely to pay attention to these events than to capital expenditure investments.

Our paper is also related to several recent papers that examine the role of the workforce in corporate decisions. Pagano and Volpin (2005) argue that managers who want to enjoy higher private benefits are more likely to guard against takeover threats by offering long-term contracts to workers who, to keep such contracts, are likely to resist hostile takeovers.[4] Supporting this argument, Rauh (2006b) shows that large employee stock holdings in their own companies form a takeover defense that entrenched managers can use to insulate themselves from market discipline. Similarly, Faleye, Mehrotra, and Morck (2006) investigate the role of labor in corporate governance and find that labor's voice in corporate governance is associated with lower equity value, sales growth, and job creation. Unlike these papers that focus on employees' negative role in influencing corporate decisions, our paper emphasizes their positive role in monitoring managerial behavior and shows that pension deficits influence managers to make value-enhancing investment decisions.

By examining the effect of pension deficits on takeover decisions, we extend the existing literature in two important ways. First, our paper sheds light on the governance role of pension deficits. Previous studies show that debt serves as an important mechanism to control managerial discretion, but no study investigates how pension deficits as inside debt affect managerial behavior and firm value. We show that pension deficits play an important disciplinary role in a firm's M&A decisions, influencing managers to make value-enhancing decisions. Second, our study extends the literature on the stakeholder theory of the firm by showing that employees have strong incentives to exert pressure on managerial behavior when their claims on firm value are at stake. Further, we show that the interests of employees and shareholders are more closely aligned when a large portion of workers' retirement claims is tied to managers' investment quality, thereby identifying an important channel through which pension plan funding status is linked to shareholder wealth. Our results therefore are consistent with Acharya, Myers, and Rajan (2011), who argue that employees can act as an internal governance mechanism for the management. They also complement the findings of Cronqvist et al. (2009), who show that entrenched CEOs are willing to pay employees more to enjoy private benefits.

The rest of the paper is organized as follows. Section 2 briefly reviews the institutional background of U.S. pension plans and develops the paper's main hypotheses. In Section 3, we describe the data and report summary statistics. Section 4 outlines our empirical methodology and presents the empirical results. Section 5 discusses robustness tests. Finally, we present summary and concluding remarks in Section 6.

2. Institutional Background of Corporate Pension Plans and Main Hypotheses

2.1. Types of Pension Plans and Laws Related to DB Pension Plans

There are two basic types of retirement plans in the U.S., namely, defined contribution (DC hereafter) and DB pension plans. A DC plan is similar to a savings account. It requires that employers, and possibly employees, make regular contributions each year to employees' pension accounts. The employees' final retirement benefits hinge upon the total contributions and the investment performance of pension assets. Employees have discretion over the assets into which they invest and bear all of the shortfall risk upon retirement.

A firm sponsoring a DB plan, on the other hand, has an obligation to retirees and current employees that amounts to the present value of the future payments estimated based on various actuarial assumptions concerning mortality rates, discount rates, etc. To meet a stream of future committed payments, the firm makes periodic tax-deductible contributions to a pension fund. If the value of pension assets is insufficient to pay the promised benefits, the firm is responsible for the shortfall (i.e., pension deficit). The pension funding status of a DB plan is considered underfunded (fully funded or overfunded) if the present value of the pension liabilities is more than (equal to or less than) the fair value of the pension assets.

Pension plan sponsors are required by law to make mandatory contributions to their underfunded pension plans. These mandatory pension contributions are mainly determined by the extent of pension deficits and funding rules established by the Internal Revenue Code and several pension protection acts that went through the changes over time, such as the Employment Retirement Income Security Act (ERISA) of 1974, the Pension Protection Act of 1987, the Retirement Protection Act of 1994, and the Pension Protection Act of 2006.

The PBGC, created by the ERISA of 1974, insures the benefits of DB plan participants and serves as a statutory trustee of terminated pension plans when a sponsoring firm has insufficient assets to pay the benefits that participants are owed. PBGC insurance coverage of benefits under the DB plan is limited to a certain extent and far from comprehensive from the standpoint of employees.[5] In addition, because of limitation provisions set by the ERISA, large pension plans and participants with benefits exceeding the maximum limit are not fully protected by the PBGC.[6] Thus, in general, pension deficits are considered to be sponsoring firms' unsecured senior inside debt owed to employees, which are partly insured by the PBGC.

2.2. Main Hypotheses

To provide theoretical guidance to our empirical tests, in this subsection we discuss several rationales for the disciplinary role of pension deficits in limiting managers' discretionary power.

First, underfunded plans can have a significant bearing on welfare and job security of employees. For example, Bulow (1982) argues that employees of DB plan sponsoring firms sacrifice high wages for the stable pension income and that large pension deficits impose tremendous pressure on sponsors and expose employees to considerable risk of losing their jobs. Similarly, Munnell, Aubry, and Muldoon (2008) suggest that large pension deficits may force stressed firms to lay off workers, financially unhealthy firms to go bankrupt, or compel healthy firms to freeze their DB plans. They also show that in these events, employees would end up with significantly lower retirement incomes than those they had anticipated, despite the protection provided by the PBGC, because retirement benefits are usually computed based on wages at the time of the layoff, bankruptcy, or freeze, instead of at retirement. Moreover,their analysis suggests that compared to retirees and employees who are about to reach full retirement age, active and young employees are more adversely affected by the risk caused by pension deficits. Thus, to the extent that employees, especially active and young ones, have strong incentives to minimize these potential costs imposed by pension deficits and care about the long-term viability of the firm, they are expected to play an important role in monitoring firms' investment decisions when their pension plans are significantly underfunded.

Second, Jensen (1986) argues that managers with large free cash flow have incentives to overinvest beyond the optimal level, but large debt obligations allow managers to effectively bond their promise to pay out future cash flows, thus reducing the agency costs of free cash flow. Paying a stream of committed pension obligations to retirees and making mandatory periodic contributions to meet future retirement payments for current employees can significantly reduce the free cash flow available to managers. Furthermore, the PBGC charges high insurance premiums to firms whose pension plans are severely underfunded.[7] To avoid these high insurance premiums, firms with underfunded pension plans usually use their internal cash flow to accelerate their contributions, which further reduces the cash flow available for managers to spend at their discretion.[8]

Third, as Fama (1985) argues, inside debtholders have access to private information about borrowers that is not easily available to other debtholders, which provides them a significant advantage in monitoring their borrowers. Since employees participate in a firm's daily operations and are able to directly observe daily management decisions, they can be considered as important inside debtholders. Moreover, compared to public debtholders, employees are likely to spend less time and effort collecting information about their employer since they are on- the-spot. These information advantages are therefore expected to provide employees with enhanced monitoring capabilities and in turn stronger incentives to monitor their employer.

Finally, Hanka (1998) shows that higher debt is associated with more layoffs, greater reliance on part-time and seasonal employees, and lower wages. To the extent that these adverse effects of high debt on employees weaken the incentives of employees to work hard and lead to higher employee turnover, pension deficits, which account for a significant portion of a firm's debt, can have similarly negative effects on employee behavior. In addition, previous studies show that debt rating agencies take pension deficits into account when evaluating a firm's credit rating, with large pension deficits leading to a high cost of debt (Cardinale (2007), Rauh (2007)). Thus, managers of firms with large pension deficits are expected to make value-enhancing investment decisions to mitigate the adverse effects of pension deficits on employee behavior and the cost of capital.

We empirically evaluate the above arguments for the disciplinary role of pension deficits as follows.

First, we investigate whether the likelihood of engaging in diversifying mergers is lower for DB firms with a larger pension deficit than for other firms. According to Lang and Stulz (1994), the value-reducing consequences of investments are larger for diversified acquisitions than for non-diversified acquisitions. We therefore expect that all else being equal, DB firms with a larger pension deficit are less likely to engage in diversifying mergers than other firms.

Second, we examine the relation between the extent of an acquirer's pension deficits and the announcement returns for both the acquirer and the value-weighted portfolio of the acquirer and the target. If large pension deficits allow the acquirer to overcome free cash flow problems and the reduction in agency problems translates into better acquirer performance, we expect the abnormal returns for both the acquirer and the value-weighted portfolio of the acquirer and the target to be higher when the acquirer has larger pension deficits.

Third, we examine whether the takeover premiums paid by DB acquirers with a larger pension deficit are different from those paid by other acquirers. Our arguments above suggest that offer prices are affected by the disciplinary role of pension deficits, as pension deficits help reduce the extent of managerial overconfidence (Roll (1986)). Since the interests of bidding firms' employees and shareholders are likely to be more closely aligned when bidding firms have larger pension deficits, we expect DB acquirers with a large pension deficit to pay smaller takeover premiums to their targets than other acquirers.

Finally, we examine whether the extent of an acquirer's pension deficits affects its methods of payment in acquisitions. Harford, Klasa, and Walcott (2009) find that acquirers use a smaller fraction of cash in paying for acquisitions when they are highly leveraged, while Shleifer and Vishny (2003) and Dong et al. (2006) show that acquirers are more likely to use cash as a method of payment when they acquire undervalued targets. Thus, if a firm's pension deficit is simply a part of its debt, then an acquirer's pension deficit is expected to be negatively related to the fraction of cash used in the payment for a target. However, if pension deficits serve as an effective controlling mechanism and thus influence managers to make better acquisition decisions (i.e., to choose undervalued targets), then an acquirer's pension deficit is expected to be positively related to the fraction of cash used in acquisitions.

3. Data

3.1. Sample

Our sample consists of U.S. acquiring firms (both DB and non-DB firms) in mergers between 1981 and 2008. The initial sample of mergers comes from Thomson Financial's Security Data Corporation (SDC) Platinum database.[9] Our final sample includes all completed mergers that meet the following selection criteria:

(1) The acquiring firm is publicly traded and owns less than 50% of the target's shares before the announcement date and controls 100% of the target's shares after acquisition.

(2) The deal value disclosed in SDC exceeds 1% of the acquiring firm's market value of equity at the end of the fiscal year prior to the acquisition announcement.

(3) There are no missing data on the book value of assets, stock returns, and the control variables used in the regression analysis.

These sample criteria yield a final sample of 26,325 mergers. To mitigate the impact of outliers or misrecorded data on the results, all firm characteristics are winsorized at the 0.5% level at both tails of the distribution. All dollar values are converted into 2000 constant dollars using the GDP deflator. The accounting items related to DB pension plans and financial statement data are drawn from Compustat.[10] Stock price data come from the Center for Research in Security Prices (CRSP) files.

Firms are defined to have DB plans if both the fair value of pension plan assets (*PPA*) and the projected pension benefit obligation (*PBO*) are available in Compustat. The key variable of our interest, the pension deficit (*DEF*), is defined as the difference between *PBO* and *PPA*. Thus, a positive (negative) *DEF* indicates that the firm's DB pension plan is underfunded (overfunded). A positive pension deficit represents a true liability for the sponsoring company even if it does not appear on the balance sheet. The Online Appendix briefly reviews the history of pension accounting in the U.S. and describes how *PPA* and *PBO* are calculated using Compustat data. Following Jin, Merton, and Bodie (2006) and Shivdasani and Stefanescu (2010), we normalize each of the pension variables (*PBO*, *PPA*, and *DEF*) and control variables used in subsequent regressions using the end-of-period adjusted book value of assets (*Assets*). *Assets* is measured by adjusting the reported book value of assets for pension-related items, such as prepaid pension ost, accrued pension liabilities, and additional minimum liability.[11] Robustness checks, tabulated in the Online Appendix, show that using alternative measures of pension deficits does not affect our main results.

3.2. Summary Statistics

Table 1 reports summary statistics for our sample mergers by year. The number of mergers increases until the late 1990s, peaking in 1998 and then dropping off significantly through the end of the sample period. Column (3) shows that during our sample period, 31.8% of sample firms sponsor DB pension plans. Column (3) also indicates that more than 50% of acquiring firms have DB plans in the early 1980s. However, the percentage of DB acquirers decreases thereafter, hitting the lowest level around the 2001 recession, although it rebounds slightly afterwards until the end of the sample period.

Columns (4)-(9) of Table 1 report the time series of annual mean and median values of *PBO/Assets*, *PPA/Assets*, and *DEF/Assets*, respectively.On average, DB pension plans of acquiring firms had been sufficiently funded until the 1990s,during which time the booming stock markets increased the value of pension assets. However, owing to both low interest rates and weak stock markets, the value of pension assets dropped significantly in the 2000s, leading to positive mean and median values of *DEF* between 2001 and 2008.

Panel A of Table 2 reports the mean firm characteristics for the full sample (26,325) as well as those for the subsample (8,382) of mergers involving acquirers with DB pension plans.[12] DB acquirers are further divided into companies with a pension deficit (*DEF* > 0) and those with a pension surplus (*DEF* ≤ 0).

Comparing DB acquirers with a pension surplus with those with a pension deficit (columns (4) and (5)), we find that the latter firms are on average younger and more leveraged, and they have fewer tangible assets than the former firms. In addition, underfunded acquiring firms exhibit higher growth opportunities and a higher likelihood of financial distress than overfunded firms. Panel B of Table 2 reports the mean deal characteristics for our M&A sample. We find that acquirers with underfunded DB plans (*DEF* > 0) make fewer diversifying mergers than those with overfunded DB plans (*DEF* ≤ 0). Acquirers with underfunded DB plans are also less likely to acquire publicly held firms, make hostile acquisitions, and participate in multiple bids.

Panel B of Table 2 also reports the cumulative abnormal return (*CAR*) for the acquirer from one day before to one day after the announcement date of the mergers (*CAR*(-1,1)), the market- capitalization weighted *CAR* of the acquirer and the target during the same event window (*WCAR*(-1,1)), and the premium paid to the target by the acquirer (*Premium*). We employ a standard event study methodology to measure abnormal announcement day returns. To obtain our estimates of the market model we use 200 trading days of return data, beginning 205 days before and ending 6 days before the M&A announcement date. We use as the market return the CRSP value-weighted return. Using the CRSP equally-weighted returns yields qualitatively similar results. Following Bradley, Desai, and Kim (1988), we use the market capitalizations of the bidder and the target two days prior to the announcement date as the weights in calculating *WCAR*(-1,1). Following Dong et al. (2006), we calculate *Premium* as the difference between the acquirer's offer price (total value of cash, stock, and other securities offered by the acquirer to the target) and the target's market value of equity 5 days prior to the M&A announcement date, scaled by the target's market value of equity on the same day.[13] To minimize the impact of extreme values and misrecorded data, we also follow Dong et al. (2006) and truncate *Premium* at 50% and 150%.

When comparing acquirers with underfunded DB plans to those with overfunded DB plans, we find that the former acquirers realize higher *CAR*(-1,1) and higher *WCAR*(-1,1) and use more cash as the method of payment. These results, together with those for the deal characteristics discussed above, suggest that acquirers with underfunded DB plans are more likely to undertake value-enhancing mergers than acquirers with overfunded DB plans.[14]

4. Empirical Methodology and Results

4.1. Methodology

To examine the impact of *DEF/Assets* on various aspects of acquisition activities, our empirical specification needs to address two types of endogeneity. The first type concerns the nonrandom choice of whether or not to be a DB plan sponsor. The pension deficit is only observable for DB plan sponsors. However, sponsors' choice between DB and non-DB pension plans is likely to be endogenous and some firms may self-select into becoming DB plan sponsors based on their firm characteristics. Second, conditional on a firm being a DB plan sponsor, the firm's pension funding status, i.e., the size of its pension deficit, may also be endogenously determined.

To address these two types of endogeneity concerns, we employ a three-step bootstrapping procedure that is similar to the approach used by Chang, Dasgupta, and Hilary (2009) and Shivdasani and Stefanescu (2010). Specifically, in the first step we model the likelihood of a firm being a DB plan sponsor. We use as an instrumental variable the industry level unionization rate (*Union*) in the first step.[15] Shivdasani and Stefanescu (2010) argue that the degree of unionization should be positively related to the labor force's negotiation power with respect to the adoption of DB pension plans, but orthogonal to the size of pension benefits. In the second step, conditional on a firm being a DB sponsor, we predict the size of its pension deficit using the age of the pension plan (*Plan Age*) as an instrumental variable. Petersen (1996) shows that pension fund managers become more sophisticated in managing their pension plan as the plan ages. Atanasova and Gatev (2010) document that plan age is positively associated with the return on pension assets. We therefore expect the age of a pension plan to have a negative impact on pension deficits. In the meantime, we have no *a priori* reason to expect plan age to have direct impact on acquisition outcomes. To obtain consistent and asymptotically efficient coefficient estimates, we jointly estimate the regressions in the first and second steps using Heckman's (1979) maximum likelihood estimator. This procedure yields the predicted pension deficit $\overline{DEF / Asset}$. In the final step we use $\overline{DEF / Asset}$ and its interaction with DB (a dummy variable that equals one if the firm chooses a DB pension plan and zero otherwise) as the key independent variables in regressions to investigate the impact of pension deficits on various aspects of acquisition activities. Although the last step is the focus of our empirical analysis, the first and second steps are included to control for self-selection bias and the endogeneity of the size of pension deficits, respectively. To further mitigate the latter effect, we estimate the size of pension deficits lagged one period based on firm-specific variables lagged two periods. We bootstrap the three-step system 500 times to obtain consistent standard errors and report the coefficients' 95% confidence interval estimates in the tables. The Online Appendix provides a detailed description of this three-step procedure.

4.2. Determinants of Pension Plan Choice and Size of Pension Deficit

Table 3 reports the results obtained by jointly estimating the first and second steps using Heckman's (1979) maximum likelihood estimator. We use the full sample of Compustat firms (13,569 unique firms and 115,960 firm-year observations) to estimate the regressions. Column (1) presents the first-step estimates of the pension selection equation, which concerns the decision of whether to sponsor a DB pension plan (i.e., choice of pension plan). The second-step estimates, in which *DEF/Assets* is the dependent variable, are reported in column (2).

Column (1) of Table 3 shows that larger firms, older firms, and firms with more employees, lower growth (measured as the market-to-book ratio), higher *ROA*, lower earnings volatility, higher asset tangibility (measured as the ratio of PPE to total assets), and higher book-simulated marginal tax rate (*MTR*) (Graham and Mills (2008)), are more likely to sponsor DB pension plans. We also find that employee tenure (*Tenure*), defined as the median employee tenure for two-digit SIC industry firms obtained from the Employee Benefits Survey provided by the Bureau of Labor Statistics, is negatively associated with the decision to adopt DB pension plans (*z*-statistic = -4.6). More importantly, from an identification perspective, we find that the impact of the unionization rate (*Union*) on the incidence of DB pension plans is positive and highly significant (*z*-statistic = 23.2), suggesting that firms are more likely to adopt DB pension plans when the labor force is organized by unions.

Column (2) indicates that the size of a firm's pension deficit is large for small DB plan sponsors, young DB plan sponsors, and DB plan sponsors with a large number of employees, high growth potential, a high likelihood of financial distress as measured by *Z-Score*, low *ROA*, high earnings volatility, low asset tangibility, low marginal tax rate, and low employee tenure. In addition, we find that firms with negative book value of equity and high leverage are more underfunded. In contrast, firms with rated debt have better pension funding status, possibly due to easier access to the corporate bond market. The size of a firm's pension deficit is found to decrease with the level of interest rates and stock market returns. We also find that the age of a firm's pension plan has a negative effect on the size of its pension deficit (*z*-statistic = -2.5), suggesting that a firm's funding status generally improves as the plan ages.

The λ coefficient, a statistic for the selectivity effect, in our Heckman estimation is -0.004 and is significant at the 5% level, indicating that self-selection is indeed a concern. The estimated correlation between the error terms in the equations in the first and second steps is -0.10 and significant at the 5% level.

4.3. Effects of Pension Deficits on M&A Activities

In this subsection, using multivariate regression models, we examine how pension deficits affect various aspects of M&A activities, such as the decision to engage in diversifying mergers, acquirer returns and portfolio returns of the acquirer and the target around the M&A announcement dates, takeover premiums paid by acquirers, and the choice of payment methods in acquisitions. Our tests include both DB and non-DB acquirers and compare the marginal effects that pension plan type has on various aspects of M&A activities. The key variable of interest is the interaction between *DB* and *DEF/Assets*, which captures the impact of DB acquirers' pension deficits. Throughout our regression analysis, we include two-digit SIC industry indicators to control for industry-specific merger waves. We also include year indicators to account for merger waves over time. Untabulated tests show that industry and year indicators are jointly significant, indicating the importance of merger waves across industries and over time in M&A activities.[16]

4.3.1. Pension Deficits and the Decision to Engage in Diversifying Mergers

To examine the effect of an acquiring firm's DB pension deficits on its decision to engage in diversifying mergers, we use a probit regression. Specifically, using a sample of 26,325 mergers (13,764 diversifying mergers and 12,561 nondiversifying mergers, we estimate the following probit model to predict which firms make diversifying mergers:[17]

$$P[Diversify = 1] = F(d_1 + d_2 DB + d_3 DB \times DEF / Assets + d_4 C + \varepsilon), \qquad (1)$$

where *P* stands for the probability of acquirers engaging in diversifying mergers, *Diversify* is a dummy variable that takes a value of one if the acquirer and the target are in the same 3-digit SIC industry and zero otherwise, *F* denotes the normal cumulative distribution function. The coefficient on *DB*, d_2, reflects the difference in the likelihood of acquirers making diversifying mergers between DB plan sponsors and non-DB firms. To the extent that pension deficits reduce the incentives of DB plan sponsors to engage in diversifying mergers, the coefficient on the interaction term between *DB* and *DEF/Assets*, d_3, is expected to be negative. In equation (1), *C* denotes a set of control variables shown by prior literature to influence a firm's decision to acquire. Following Harford (1999), we include as control variables *Leverage*, *Ln*(*Assets*), *Ln*(*Age*), *PPE/Assets*, *ROA*, *Earnings Volatility*, *M/B*, *Sales Growth*, and *Stock Return* (the compounded monthly stock returns over the fiscal year prior to the acquisition). As a key control variable in regressions, *Leverage* is measured as the ratio of the sum of short-term and long-term debt to *Assets*. We include as additional control variables *Cash/Assets* and *Cash Flows/Assets* to account for the effect of internal funds on investment decisions. Since dividend payments can reduce the cash available to managers for new acquisitions, we also add *Dividend/Assets* as a control variable. To mitigate endogeneity concerns, all independent variables in equation (1) are pre-determined (one-period lagged) except *Cash Flows/Assets*, which is calculated using contemporaneous cash flows (Rauh (2006a)).

Columns (1) and (2) of Table 4 report the estimates from the probit regressions. We report the marginal effects that measure the effect of a one unit change in the continuous explanatory variables (moving from zero to one for dummy variables) on the dependent variable. Column (1) presents estimates obtained by using a simple maximum likelihood approach and the actual value of *DEF/Assets* in the probit regression. For non-DB firms, we set their *DEF* equal to zero. We allow for clustering of firm observations to adjust the standard errors for serial correlation and also correct standard errors for heteroskedasticity. In column (2), we report the results obtained using the three-step bootstrapping procedure.

We find that the results using these two different estimation procedures are qualitatively similar. In column (1), the coefficient estimate on *DB×DEF/Assets* is negative and significant at the 1% level in column (1).[18] The estimated coefficient (-0.835) suggests that increasing *DEF/Assets* by a one standard deviation lowers the probability of making diversifying mergers by 3.4% (= -0.835 × 0.041) for DB plan sponsors. In contrast, conventional debt ratio, *Leverage*, has a statistically insignificant coefficient of 0.035 (*z*-statistic = 1.3). The result using the bootstrapped approach in column (2) shows that the estimated coefficient on *DB×DEF/Assets* is also negative and significant at the 5% level.[19] Taken together, our findings in Table 4 suggest that acquiring firms with underfunded DB pension plans are less likely to make diversifying mergers if they decide to enter the market for corporate control. These results are consistent with our hypothesis that pension deficits, as inside debt owed to employees, limit managerial incentives to spend resources on empire building.

4.3.2. Pension Deficits and M&A Announcement Effects

To examine the impact of pension deficits on the quality of investment, we regress acquirers' *CAR*(-1,1) on *DB*, *DB×DEF/Assets*, and acquirer and deal characteristics.[20] As acquirer characteristics, we include *Leverage*, *Ln*(*Assets*), *M/B*, and *Cash Flows/Assets*. To control for deal characteristics, we include *Relative Size*, *Multiple Bids*, *Diversify*, *High Tech* (a dummy variable equal to one if a deal is made between two high tech firms as defined by Loughran and Ritter (2004) and zero otherwise), *Hostility*, *Public Target*, and *Industry M&A* (for each year and each of the three-digit SIC industries, the value of all SDC acquisition deals in the industry divided by total book value of assets of Compustat firms in the same industry). We include *Relative Size* and *Multiple Bids* because Moeller, Schlingemann, and Stulz (2004) show that bidder announcement returns increase with relative deal size and decrease when there are multiple bidders. We include *Diversify* since Morck, Shleifer, and Vishny (1990) show that bidders earn negative returns when making unrelated acquisitions. *High Tech* is included as Loughran and Ritter (2004) document that when both the acquirer and the target are in high-tech industries, the acquirer is more likely to underestimate the costs and overestimate the synergies of the combination. Schwert (2000) finds that acquirers realize lower abnormal returns in hostile takeovers, so we include *Hostility*, which equals one if the SDC classifies the acquisition as a hostile takeover and zero if the SDC classifies the acquisition as a friendly takeover. In addition, Fuller, Netter, and Stegemoller (2002) document that acquirers experience significantly negative abnormal returns when they buy publicly held targets and significantly positive abnormal returns when they buy privately held targets or subsidiaries, and thus we include *Public Target* as a control variable. Since cash (stock) is more likely to be used as a method of payment when there is low (high) valuation uncertainty in the acquisition (Loughran and Vijh (1997)), we also control for *PureCash* in the regression. Finally, we include Moeller, Schlingemann, and Stulz's (2004) industry M&A activity measure (i.e., *Industry M&A*) to control for the intensity of acquisition activity in the target industry. The regression also controls for the industry and year fixed effects.

Columns (1) and (2) of Table 5 present the regression results. Column (1) presents the estimates from ordinary least squares (OLS) regressions. The *t*-statistics are calculated using Huber/White/Sandwich heteroskedasticity-consistent errors, which are corrected for serial correlation across observations for a given firm. We find that the coefficient estimate on *DB×DEF/Assets* is positive and significant at the 1% level (*t*-statistic = 4.2). Furthermore, the effect of pension deficits on bidder returns is economically large and significant: a one-standard deviation increase in *DEF/Assets* results in an approximately 0.34% (= 0.083*0.041) increase in bidder returns. In comparison, we find that a one-standard deviation increase in leverage (0.20) is associated with an increase in bidder returns of only 0.16% (= 0.008*0.2).[21] Thus, pension deficits are more than twice as likely to have a positive effect on bidder returns as is conventional debt. In column (2), we report the results obtained using the three-step bootstrapping procedure. We find that the effect of *DEF/Assets* on bidder returns is positive and significant at less than the 5% level. Taken as a whole, these results support the disciplinary role of corporate pension deficits. We also find that the signs of the coefficients on control variables in both regressions are generally consistent with prior studies: bidder announcement returns are positively associated with *Relative Size* and *PureCash*, but are negatively related to *Ln*(*Assets*), *M/B*, and *Public Target*.[22]

To further explore whether acquirers with a pension deficit make better acquisition decisions than those with a pension surplus, we examine whether larger pension deficits are associated with higher portfolio synergistic gains between the acquirer and the target (*WCAR*(-1,1)) and lower premiums paid to targets (*Premium*). The regression results are reported in Table 6.

In columns (1) and (2) of Table 6, we use *WCAR*(-1,1) as the dependent variable. In addition to including the independent variables used in Table 5, to be consistent with Moeller, Schlingemann, and Stulz (2004), Dong et al. (2006), and Wang and Xie (2009), we include target firm characteristics(*Leverage*^{Target}, *Ln*(*Assets*)^{Target}, and *M/B*^{Target})as additional control variables.[23] We find that in column (1), which presents the estimates from the OLS regression, the coefficient estimate on *DB×DEF/Assets* is positive and significant at the 1% level (*t*-statistic= 3.3). The coefficient of 0.221 suggests that a one-standard deviation increase in *DEF/Assets* is associated with a 0.91% (= 0.221* 0.041) increase in *WCAR*(1,1). In comparison, the corresponding effect of *Leverage* on *WCAR*(1,1) is only 0.36% (= 0.018*0.2). Thus, the economic significance of pension deficits on portfolio returns is more than twice as large as that of conventional debt. The results using the three-step bootstrapping procedure in column (2) confirm those in column (1).

In columns (3) and (4) of Table 6, we use *Premium* as the dependent variable. Consistent with the prediction that larger pension deficits limit managerial hubris to overpay, we find that the coefficient on *DB×DEF/Assets* in column (3) is negative and significant at the 5% level. The coefficient of -0.705 indicates that a one-standard deviation increase in *DEF/Assets* lowers the premiums paid to targets by 2.9% (= -0.705* 0.041). The coefficient on *Leverage*, however, is not significant. The analysis using the three-step bootstrapping procedure in column (4) suggests that the bootstrapped coefficient of *DB×DEF/Assets* is negative (-1.03), but not significant at the 5% level.

Overall, the results in Tables 5 and 6 suggest that market participants value the governance role of pension deficits in disciplining managers, and hence the stock prices of acquiring firms react more favorably to acquisitions by DB sponsors with larger pension deficits. Moreover, DB sponsors with larger pension deficits pay smaller premiums to their targets than other acquirers, further supporting the disciplinary role of pension deficits.

4.3.3. Pension Deficits and Payment Methods

In this subsection we examine the disciplinary role of pension deficits by investigating whether the level of an acquirer's pension deficits affects its choice of payment methods in mergers. To the extent that acquirers are more likely to use cash as a method of payment when they acquire undervalued targets (Shleifer and Vishny (2003), Dong et al. (2006)), we expect the level of an acquirer's pension deficits to be positively related to the use of cash as the method of payment in acquisitions.

In columns (1) and (2) of Table 7, we report the estimates from tobit regressions in which the dependent variable is the cash payment as a fraction of the acquisition's transaction value (*Pct_Cash*). We use as explanatory variables those used in Dong et al. (2006). Consistent with our prediction, we find that the coefficient estimate on *DB×DEF/Assets* in column (1) is positive and significant at the 1% level. The coefficient of 3.965 suggests that a one-standard deviation increase in *DEF/Assets* leads to a 16.3% increase in the fraction of cash payment. The bootstrapped results in column (2) also indicate that this effect is significant at less than the 5% level.[24] These results suggest that larger pension deficits influence managers to make better acquisition decisions by choosing undervalued targets.

However, it is possible that acquirers with large pension deficits prefer cash mergers to stock mergers simply because they are undervalued in the market (Shleifer and Vishny (2003)).[25] It is also possible that these acquirers find it difficult to sell their equity due to a high adverse selection problem in the market. Although we control for *M/B* in regressions, to further ensure that our findings are not driven by these alternative explanations, we include past one-year stock returns as an additional control variable. Untabulated results show that the significance of the coefficient estimates on *DB×DEF/Assets* remains the same.

5. Additional Tests

To check the robustness of our main results, we conduct a battery of additional tests and report the results in Tables 8 and 9. Although the dependent and control variables used in these two tables are the same as those used in Tables 4 through 7, to save space, we only report the coefficient estimates on pension-related variables. Also, we only report results obtained using the plain-vanilla pooled OLS, probit, and tobit models. Similar results (untabulated) are obtained using the three-step bootstrapping procedure.

5.1. Self-Selection Bias

To ensure that our results are not simply driven by the inclusion of non-DB firms in our regression analysis, we reestimate the regressions in Tables 4 through 7 using DB firms only. Consistent with the results in Tables 4 through 7, we find that the coefficient estimates on *DEF/Assets* are significant in all regressions with the expected signs, suggesting that our key results are not driven by including non-DB firms in the analyses.

To further address the self-selection bias, we utilize the fact that the pension-plan choice is sticky. Specifically, we reestimate the regressions in Tables 4 through 7 using only firms that have sponsored DB plans for longer than three years. If a DB plan has been in place for several years, then it is unlikely that the choice of a DB plan is a byproduct of the firm's *current* financial characteristics and investment decisions. Thus, the degree of self-selection bias is likely to be smaller in this subsample. We find that all of the estimated coefficients on *DEF/Assets* remain statistically significant with expected signs. These results are reported in the Online Appendix.

5.2. Alternative Explanations

Faleye, Mehrotra, and Morck (2006) argue that employees can influence firms' governance through their ownership in firms. Thus, it is possible that our results are driven by employees' holdings of their company's stocks in pension assets, not necessarily by the disciplinary effect of pension deficits. To rule out this alternative explanation, using the employee ownership data from the IRS Form 5500 provided by the Centre for Retirement Research at the Boston College, we exclude DB firms with pension assets invested in their stocks from the analysis. The results tabulated in the Online Appendix show that our findings are not affected by excluding these firms.

Although our results above support the view that pension deficits play an important role in reducing inefficient investment, some of them are also consistent with the alternative explanation that firms with larger pension deficits are those with higher financial constraints.[26] For example, Rauh (2006a) shows that firms that make mandatory pension contributions, especially those that are financially constrained, reduce capital expenditures (including acquisitions) because they experience a lack of internal funds.

To see if the financial constraint explanation may be behind our results, we divide our sample acquirers into two subsamples according to Whited and Wu's (2003) financial constraint index (*WW* index) and reestimate the regressions in Tables 4 through 7 separately for these two subsamples. The financial constraint explanation would suggest that the impact of *DEF/Assets* on several aspects of takeover decisions is more pronounced for financially constrained (*FC*) firms than for financially unconstrained (*FUC*) firms. In contrast, our hypothesis predicts that the effect of *DEF/Assets* on these aspects is more evident for *FUC* firms than for *FC* firms.

The regression results for these subsamples are reported in Panel A of Table 8. We find that our key results from the previous tables (i.e., the significance of the coefficient estimates on *DB*×*DEF/Assets*) are mainly driven by *FUC* firms, suggesting that the disciplinary effect of pension deficits we have identified is not merely a reflection of a firm's financial constraints. The differences in coefficient estimates on *DB*×*DEF/Assets* between *FUC* and *FC* firms, however, are not statistically significant in most regressions (not reported). As robustness checks, in unreported tests we also experiment with other alternative measures of financial constraints, such as Hadlock and Pierce's (2010) financial constraint index, dividend-payer indicator (Fazzari, Hubbard, and Petersen (1988)), firm size (*Ln*(*Assets*)), Altman's *Z* score, and bond ratings, and find that the results are similar to those reported in Panel A of Table 8.

To further ensure that the channels through which pension deficits affect investment decisions are not merely through the cash drain caused by financial constraints, we augment the regressions in Tables 4 through 7 by including a proxy for mandatory pension contribution (*MC*), which is usually triggered by large pension deficits. Following Campbell, Dhaliwal, and Schwartz (2011), we measure *MC* as the ratio of pension expenses, as recorded in Compustat, to total assets if a firm's pension plan is underfunded, and zero if a firm's pension plan is fully funded or overfunded. Since *MC* can be measured only for DB firms, we use only these firms in the analysis. The results are reported in Panel B of Table 8. We find that the estimated coefficients on *DEF/Assets* remain economically and statistically significant even after controlling for mandatory pension contributions.

5.3. Regulatory Change as a Quasi-Exogenous Event

Pension Protection Act of 1987 requires a "deficit reduction contribution" in addition to the "minimum funding contribution" after 1987.[27] This regulatory change can serve as a quasi- exogenous event that helps us deal with omitted variables and endogeneity and thus allows us to examine the effect of pension deficits on acquisition outcomes before and after the regulation change. The requirement of the act suggests that the disciplinary role of pension deficits should be more pronounced in the post-1987 period than in the pre-1987 period. The results reported in Panel C of Table 8 suggest that the disciplinary effect of pension deficits indeed mainly come from the post-1987 period.

5.4. Importance of Pension Plan Characteristics

In this subsection, using a sample of acquiring firms sponsoring a DB plan, we examine how pension plan characteristics affect the disciplinary role of pension deficits in mergers. The first characteristic we consider is the fraction of actively working employees who are covered by DB pension plans. While DB plans cover both active and retired employees, active employees are the ones who participate in firms' daily operations and directly observe daily management decisions since they are on-the-spot. Moreover, according to PBGC's Guarantee Limits – an Update (2008), the average loss in benefits computed based on the PBGC maximum insurance limitation is twice as large for active participants as for retired participants. Thus, active employees of firms with underfunded pension plans should have stronger capabilities and incentives to monitor managers than retirees, suggesting that the control function of pension deficits is more effective when the plans cover a larger faction of active employees.

We use two pension demographic characteristics to address this issue. The first variable is the age of a pension plan (*Plan Age*), which is measured as the number of years since a firm reported pension data in Compustat. Atanasova and Gatev (2010) use this variable as a measure of pension plan maturity under the assumption that older pension plans cover relatively older workforce. The second variable is the faction of active employees, which is computed as the ratio of the number of active employees to the sum of the numbers of active and retired employees in the DB plan.[28] The regression results are reported in Table 9. In Panels A1 and A2, we divide firms into active-employee dominated (*Active*) and retiree dominated (*Retired*) firms according to the sample medians of *Plan Age* and the fraction of active employees in the pension plan, respectively, and reestimate the regressions in Tables 4 through 7 separately for these two subsamples.[29] We find that the coefficient estimate on *DEF/Assets* is more significant in regressions using an *Active* subsample than using a *Retired* subsample, suggesting that the disciplinary role of pension deficits is indeed more salient for plans with greater shares of active employees.

The second characteristic we consider is whether the pension plans are collectively bargained or whether firms are in more unionized industries. Employees' incentives and abilities to exert influence over managers become stronger if they are able to coordinate their actions and organize their bargaining power in a more systematic way. Supporting this view, Comprix and Muller (2011) argue that collective bargaining agreement and union give employees a collective voice through which they can represent their preference and dissatisfaction and thus enhance employees' incentives to exert efforts. Thus, we expect that the control function of pension deficits, if it exists, should be more effective for firms whose pension plans are collectively bargained or those that operate in industries where labor force is well organized.

To examine this conjecture, we partition our sample acquiring firms according to whether their pension plans are collectively bargained and then reestimate the regressions in Tables 4 through 7 separately for these two subsamples. The information on whether the pension plans are collectively bargained is obtained from Form 5500. The results are reported in Panel B of Table 9. We find that the significant relation between pension deficits and various acquisition outcomes is evident only for firms whose pension plans are collectively bargained, supporting the view that collective bargaining provides employees with stronger incentives to monitor.

To future explore the impact of bargaining power on employee monitoring, we bifurcate our sample according to the sample median industry unionization rate (*Union*) and repeat our analysis in Tables 4 through 7 separately for firms with high and low unionization rates. The results reported in Panel C of Table 9 show that the impact of pension deficits on several aspects of takeover decisions is indeed more pronounced for firms in more unionized industries than for firms in less unionized industries, confirming the argument of Chen, Kacperczyk, and Ortiz- Molina (2010).

5.5. Target Firms' Pension Deficits and Takeover Gains

In this subsection we examine the role of target firms' pension deficit in explaining these gains. Pontiff, Shleifer, and Weisbach (1990) argue that firms have incentives to acquire targets with overfunded pension plans to engage in potential post-acquisition pension fund reversions.[30] To the extent that target shareholders require higher takeover premiums for mergers motivated by pension fund reversions and acquirers' incentives for pension fund reversions are greater when they have larger pension deficits, the reversion explanation predicts that targets' pension deficits are *negatively* related to takeover premiums.

To test this prediction, in untabulated tests, we augment our takeover premium model in column (3) of Table 6 by including pension deficit measures for both acquirers and targets ($DEF^{Acquirer}/Assets^{Acquirer}$ and $DEF^{Target}/Assets^{Target}$). We find that the coefficient estimate on our main variable of interest, $DEF^{Acquirer}/Assets^{Acquirer}$, remains negative and significant at the 1% level (coefficient = -0.763). However, the oefficient estimate on $DEF^{Target}//Assets^{Target}$ is *positive* and significant at the 1% level (coefficient = 0.758), which does not support the pension assets reversion explanation of takeover premiums.[31] However, our result that there is a positive relation between targets' pension deficits and takeover premiums is consistent with the view that pension deficits are an important control mechanism for the targets as well. The result also supports the view that targets' pension deficits are an important anti-takeover device that deters potential buyers from purchasing target shares. Using UK data, Cocco and Volpin (2010) document that firms sponsoring DB pension plans, especially those having large pension deficits, are less likely to be targeted and acquired. Since employees in targets with underfunded pension plans are concerned about their post-acquisition retirement benefits, to reduce employee resistance in target firms, acquirers may have to pay higher acquisition premiums, which can be used to compensate target employees for their underfunded pension plans. Consistent with the view that anti-takeover devices increase acquisition premiums, Heron and Lie (2006) show that conditional on attempted takeovers, anti-takeover devices such as poison pill enable target management to negotiate for a higher takeover premium.

Finally, we examine whether high announcement returns for the acquirer and the value- weighted portfolio of the acquirer and the target are due to synergistic gains created by merging the acquirers with an underfunded plan with the targets with an overfunded plan. Mergers between these types of acquirers and targets may benefit both parties because otherwise excess assets in the overfunded pension plan cannot be fully and directly reverted to sponsors through standard terminations or reversions given high excise tax rates on asset reversions. For example, in 1984, Gulf Oil agreed to a friendly merger with Chevron Oil Corp in an attempt to recover its excess pension assets. As a result, approximately $550million excess pension assets accumulated in the Gulf Oil pension plan were merged into the Chevron pension plan, creating a single overfunded pension plan after the merger. To examine this synergy argument as a potential explanation for acquirer returns and value-weighted portfolio returns of the acquirer and the target, we add $DEF^{Target}/Assets^{Target}$ and its interaction with $DEF^{Acquirer}/Assets^{Acquirer}$ in regression models reported in Tables 5 and 6 and reestimate these regressions. Untabulated results show that in all regressions, while the coefficient estimates on $DEF^{Acquirer}/Assets^{Acquirer}$ remain statistically significant, the coefficient estimates on $DEF^{Acquirer}/Assets^{Acquirer} \times DEF^{Target}/Assets^{Target}$ are not significant, suggesting that synergistic gains created by combining two firms with different pension funding status, if they exist, are dominated by the control function of pension deficits in acquiring firms.

6. Summary and Conclusion

Despite the fact that pension deficits have increased significantly during the past two decades, few empirical studies examine how these claims owed to employees affect firms' investment decisions and performance. In this paper, we investigate this little-explored issue using a large sample of mergers during the 1981 to 2008 period and show that pension deficits play an important role in disciplining managerial discretion in M&A decisions. Specifically, we find that firms with larger pension deficits are less likely to engage in diversifying mergers. We also find that acquirers' pension deficits have a significant positive effect on their announcement returns as well as the announcement returns of value-weighted portfolio of the acquirer and the target. Further, compared to acquirers with smaller pension deficits, those with larger pension deficits pay lower takeover premiums to their targets and use more cash in payment to their targets. These results are more pronounced for subsamples of acquirers whose pension plans are dominated by actively working employees and acquirers in more unionized industries. Finally, we examine several alternative hypotheses that might explain our results, such as financial constraints, cash drains caused by mandatory contributions, employee ownership in pension assets, post-acquisition pension fund revisions, and synergistic gains created by merging the acquirers with an underfunded plan with the targets with an overfunded plan, and find little support for these alternative explanations. Overall, our results suggest that corporate pension deficits serve as an effective monitoring mechanism that influences managers to make value- enhancing investment decisions.

Although our tests identify merger decisions as the set of decisions over which employees have strong incentives to monitor managerial behavior when their claims on the firm's value are at stake, these incentives could also exist in other instances. For firms with large pension deficits, the questions of when the interests of employees are more likely to be closely aligned with those of shareholders and when employees exert a strong influence on management decisions represent useful areas for additional work.

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Table 1

Sample distribution and mean and median values of pension variables by year

The sample consists of firms with either a defined benefit (DB) or a non-DB pension plan acquiring the majority shares of target firms in the U.S. between 1981 and 2008, reported in the SDC's Mergers and Acquisitions database. Excluded are firms that have missing data on the book value of assets, stock returns, and variables used in the regression analysis. Firms are defined to have DB plans if both the fair value of plan assets (*PPA*) and the projected pension benefit obligation (*PBO*) are available in Compustat. The pension funding deficit (*DEF*) is defined as the difference between *PBO* and *PPA*. All pension items are deflated by the end-of-period book value of assets adjusted for pension items on the balance sheet (*Assets*) and are winsorized at the 0.5% level at both tails of the distribution.

Year	Number of total firms	% of DB plan firms	*PBO/A* (%)		*PPA/A* (%)		*DEF/A* (%)	
			Mean	Median	Mean	Median	Mean	Median
(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
1981	244	66.0	7.4	5.0	8.8	6.6	-1.4	-1.1
1982	390	59.2	8.4	5.3	9.9	7.1	-1.4	-0.9
1983	459	54.2	7.9	5.5	10.1	7.5	-2.2	-1.7
1984	305	67.9	9.7	6.2	12.9	9.0	-3.0	-2.5
1985	318	62.3	9.4	7.1	12.4	9.3	-2.9	-2.1
1986	348	54.0	11.1	8.5	13.0	10.2	-1.9	-1.5
1987	399	48.9	10.9	9.2	12.7	9.6	-1.6	-0.8
1988	460	43.3	10.7	8.6	11.7	8.8	-0.9	-0.3
1989	448	39.1	12.4	8.6	14.1	10.0	-1.4	-0.7
1990	498	29.5	11.6	8.1	12.3	8.5	-0.6	-0.2
1991	650	28.0	10.7	7.0	11.3	7.4	-0.5	-0.1

1992	952	27.1	12.4	8.4	12.6	8.2	-0.3	0.0
1993	1,262	32.4	10.4	4.7	10.2	3.7	0.3	0.0
1994	1,456	28.5	10.3	6.8	10.1	5.3	0.2	0.1
1995	1,716	25.2	11.2	6.2	11.1	5.8	0.1	0.2
1996	2,178	22.9	10.5	5.8	11.0	5.3	-0.4	0.0
1997	2,237	26.1	10.2	6.0	11.2	6.4	-0.9	-0.1
1998	1,786	28.6	12.4	7.7	12.8	7.1	-0.3	0.0
1999	1,537	24.8	12.0	7.4	13.5	7.6	-1.6	-0.4
2000	1,154	28.3	10.4	5.8	11.2	5.3	-0.8	-0.1
2001	974	27.6	12.0	7.1	11.7	6.6	0.6	0.3
2002	1,005	32.3	13.2	5.4	10.6	3.7	2.6	1.3
2003	1,139	31.0	12.4	6.0	10.4	4.7	2.0	1.1
2004	1,067	32.5	12.6	6.6	10.5	4.9	2.2	1.3
2005	1,063	31.8	15.0	7.3	12.8	5.8	2.4	1.5
2006	1,074	36.6	12.3	5.8	11.0	4.5	1.4	0.9
2007	767	35.1	10.2	5.3	9.4	4.1	0.8	0.4
2008	439	31.9	11.6	6.1	8.8	4.4	2.5	1.6
Mean	-	31.8	11.2	6.5	11.4	6.4	-0.1	0.0
Total	26,325	-	-	-	-	-	-	-

Table 2

Summary statistics for acquiring firms and M&A characteristics

The sample consists of firms with either a defined benefit (DB) or a non-DB pension plan acquiring the majority shares of target firms in the U.S. between 1981 and 2008, reported in the SDC's Mergers and Acquisitions database. Excluded are firms that have missing data on the book value of assets, stock returns, and variables used in the regression analysis. Firms are defined to have DB plans if both the fair value of plan assets (*PPA*) and the projected pension benefit obligation (*PBO*) are available in Compustat. The pension funding deficit (*DEF*) is defined as the difference between *PBO* and *PPA*. All pension items are deflated by the end-of-period book value of assets adjusted for pension funding status. *DEF* > 0 means that DB firms have pension deficits and *DEF* ≤ 0 means that DB firms have pension surpluses. Dollar values (in millions) are converted into 2000 constant dollars using the GDP deflator. All variable are winsorized at the 0.5% level at both tails of the distribution. The symbols ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively.

	Total firms *N = 26,325*		DB firms *N = 8,382*			
			All	*DEF* > 0	*DEF* ≤ 0	
	Mean	Standard deviation	Mean	Mean	Mean	Test of difference
	(1)	(2)	(3)	(4)	(5)	(4) - (5)
Panel A: Acquiring firm characteristics						
Assets: millions of dollar	3422.0	11852.0	8061.4	6692.5	9615.2	-2922.8***
Ln(Assets)	5.95	2.13	7.50	7.36	7.66	-0.30***
Age: year	14.88	12.85	24.91	23.40	26.63	-3.23***
# of employees: thousands	8.46	20.77	17.97	15.80	20.47	-4.67***
Market-to-Book (M/B)	2.22	2.47	1.58	1.64	1.52	0.11***
Leverage	0.22	0.20	0.25	0.26	0.25	0.01***
Sales Growth	0.49	1.38	0.19	0.19	0.18	0.02
Stock Return	0.31	0.75	0.23	0.22	0.24	-0.02
ROA	0.05	0.20	0.09	0.09	0.09	-0.00
Cash Flows/Assets	0.04	0.42	0.10	0.10	0.10	-0.00**
Cash/Assets	0.17	0.20	0.08	0.09	0.08	0.01***
Earnings Volatility	0.09	0.10	0.05	0.05	0.05	0.01***
PPE/Assets	0.25	0.24	0.29	0.28	0.31	-0.03***
DEF/Assets	0.00	0.02	0.00	0.02	-0.02	0.04***

Panel B: Deal characteristics						
Deal Value: millions of dollar	376.3	2780.8	743.9	694.3	800.3	-105.9
Relative Size	0.32	0.75	0.26	0.27	0.26	0.01
Public Target: indicator	0.19	0.39	0.28	0.25	0.31	-0.06***
Hostility: indicator	0.05	0.21	0.08	0.06	0.10	-0.04***
Multiple Bids: indicator	0.01	0.11	0.02	0.02	0.03	-0.01
CAR(-1,1): %	1.23	8.07	0.56	0.91	0.17	0.74***
WCAR(-1,1): %	1.84	7.64	2.03	2.45	1.69	0.75**
Diversify: indicator	0.52	0.50	0.60	0.57	0.63	-0.06***
Pct_Cash: %	0.61	0.44	0.66	0.72	0.59	0.13***
PureCash: indicator	0.47	0.50	0.55	0.58	0.50	0.08***
Premium: %	43.85	35.48	43.41	43.41	43.93	-0.52

Table 3

Determinants of the choice of pension plans and the size of pension deficits

The sample consists of all U.S. firms with either a defined benefit (DB) or a non-DB pension plan covered in the Compustat database between 1981 and 2008. The regression model with selection is fitted using Heckman's maximum likelihood estimator. Column(1) reports the first-step probit estimates of the selection equation, which investigates the determinants of the choice of pension plans. Firms are defined to have DB plans if both the fair value of plan assets (*PPA*) and the projected pension benefit obligation (*PBO*) are available in Compustat. Column (2) reports the second-step estimates of the regression in which the dependent variable is the pension funding deficit (*DEF/Assets*). *DEF/Assets* is estimated as the difference between *PBO* and *PPA,* deflated by the end-of-period book value of assets adjusted for pension items on the balance sheet. The *z*- statistics and *t*-statistics are reported in parentheses. The symbols ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively.

	Dependent variable: Choice of pension plans (= 1 if a firm chooses a *DB* pension plan)	Dependent variable: Size of pension deficits (*DEF/Assets*)
	(1)	(2)
Ln(Assets)	0.242***	-0.003***
	(45.0)	(-10.3)
Ln(Employees)	0.051***	0.003***
	(5.9)	(8.5)
Ln(Age)	0.610***	-0.004***
	(74.3)	(-6.3)
M/B	-0.087***	0.001***
	(-18.3)	(3.0)
Z-Score	-0.009**	0.000
	(-2.0)	(1.2)
ROA	0.569***	-0.000
	(8.8)	(-0.0)
Earnings Volatility	-2.703***	0.036***
	(-29.1)	(6.8)
PPE/Assets	0.298***	0.004***
	(10.1)	(4.1)
MTR	0.464***	-0.055***
	(4.5)	(-12.5)
Tenure	-0.024***	0.000
	(-4.6)	(1.0)
Union	0.017***	
	(23.2)	
Negative Equity		0.008***
		(6.5)
Cash Flows/Assets		-0.002
		(-1.0)

Leverage			0.017***
			(13.6)
Debt Rating			-0.000
			(-0.7)
Stock Market Return			-0.012***
			(-6.1)
Interest Rate			-0.007***
			(-31.6)
Plan Age			-0.001**
			(-2.5)
Industry Median DEF/Assets			0.669***
			(31.9)
Industry (2-digit SIC) dummies	YES		NO
Year dummies	YES		YES
Intercepts	YES		YES
Estimated Correlation Coefficient (ρ)		-0.10	
$\chi 2$ of the Likelihood Ratio Test for $\rho = 0$		5.74**	
Heckman's λ		-0.004**	
Sample Size / Pseudo R^2 (Probit)	115,960 / 0.41		115,960

Table 4

The effect of pension deficits on the likelihood of diversifying mergers

The sample consists of firms with either a defined benefit (DB) or a non-DB pension plan acquiring the majority shares of target firms in the U.S. between 1981 and 2008, reported in the SDC's Mergers and Acquisitions database. The dependent variable is a dummy variable that equals one if the acquirer and the target are in the same 3-digit SIC industry and zero otherwise (*Diversify*). Coefficient estimates reported are the marginal effects that measure the effect of a one unit change in continuous explanatory variables (moving from 0 to 1 for dummy variables) on the dependent variable. Firms are defined to have defined benefit (DB) plans if both the fair value of plan assets (*PPA*) and the projected pension benefit obligation (*PBO*) are available in Compustat. *Cash Flow/Assets* is measured over the fiscal year in which the acquisition is announced. All other explanatory variables are measured at the fiscal year-end that immediately precedes the announcement date of share acquisitions. Column (1) reports results from the probit regression that is estimated using the actual value of pension funding deficits (*DEF/Assets*), which is calculated as the difference between *PBO* and *PPA*, deflated by the end-of-period book value of assets adjusted for pension items. The *z*-statistics in parentheses are calculated from the Huber/White/Sandwich heteroskedastic consistent errors, which are also corrected for correlation across observations for a given firm. The symbols ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively. Column (2) reports results from the three-step bootstrapping procedure outlined in the Online Appendix and uses the predicted value of *DEF/Assets* obtained using Heckman's (1979) maximum likelihood estimator. The 95% confidence intervals in square brackets are calculated from 500 bootstrap replications of the three-step estimation based on resampling from the data set with replacement of clusters.

	Dependent variable: *Diversify* (= 1 if the acquirer and the target are in the same 3-digit SIC industry)	
	Probit	Three-step bootstrap
	(1)	(2)
DB	0.086***	0.114
	(6.8)	[0.084, 0.142]
DB×DEF/Assets	-0.722***	-1.197
	(-3.3)	[-2.099, -0.369]
Leverage	-0.034	-0.025
	(-1.3)	[-0.072, 0.022]
Ln(Assets)	-0.009***	-0.011
	(-3.0)	[-0.017, -0.005]
Ln(Age)	0.031***	0.032
	(4.0)	[0.016, 0.046]
PPE/Assets	-0.160***	-0.132
	(-5.2)	[-0.183, -0.078]
ROA	-0.055**	-0.073
	(-2.3)	[-0.122, -0.022]
Sales Growth	0.003	0.009
	(1.1)	[0.002, 0.017]
M/B	-0.008***	-0.008
	(-3.7)	[-0.012, -0.004]
Cash Flows/Assets	-0.019	0.012
	(-1.6)	[-0.028, 0.051]
Cash/Assets	-0.144***	-0.167
	(-5.3)	[-0.218, -0.112]
Stock Return	0.003	0.000
	(0.6)	[-0.010, 0.009]
Earnings Volatility	-0.030	-0.037
	(-0.5)	[-0.156, 0.079]
Dividend/Assets	0.329***	0.396
	(3.5)	[0.121, 0.659]
Industry dummies	YES	YES
Year dummies	YES	YES
Constant	YES	YES
Sample size / Pseudo R^2	26,325 / 0.09	500 replications

Table 5

Regression of cumulative abnormal returns (-1, 1) for bidders on explanatory variables

The sample consists of firms acquiring the majority shares of target firms in the U.S. between 1981 and 2008, reported in the SDC's Mergers and Acquisitions database. The abnormal returns are calculated using the market model. The market model is estimated by using 200 trading days of return data ending 6 days before the acquisition announcement. The CRSP equally weighted return is used as a proxy for the market return. The dependent variable is cumulative abnormal returns from one day before to one day after the announcement date of the M&A for acquirers (*CAR*(-1,1)). Firms are defined to have defined benefit (DB) plans if both the fair value of plan assets (*PPA*) and the projected pension benefit obligation (*PBO*) are available in Compustat. The pension funding deficit (*DEF/Assets*) is defined as the difference between *PBO* and *PPA*, deflated by the end-of-period book value of assets adjusted for pension items. All explanatory variables are measured at the fiscal year-end that immediately precedes the announcement date of share acquisitions. All variables are winsorized at the 0.5% level at both tails of the distribution. Column (1) uses the actual value of *DEF/Assets* in OLS regression. The *t*-statistics in parentheses are calculated from the Huber/White/Sandwich heteroskedastic consistent errors, which are also corrected for correlation across observations for a given firm. The symbols ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively. Column (2) reports results from the three-step bootstrapping procedure outlined in the Online Appendix and uses the predicted value of *DEF/Assets* obtained using Heckman's (1979) maximum likelihood estimator. The 95% confidence intervals in square brackets are calculated from 500 bootstrap replications of the three-step estimation based on resampling from the data set with replacement of clusters.

	Dependent variable: *CAR(-1,1)*	
	OLS	Three-step bootstrap
	(1)	(2)
DB	0.001	-0.001
	(0.5)	[-0.004, 0.002]
DB×*DEF/Assets*	0.083***	0.180
	(4.2)	[0.096, 0.271]
$Leverage^{Acquirer}$	0.008**	0.004
	(2.3)	[-0.002, 0.011]
$Ln(Assets)^{Acquirer}$	-0.004***	-0.004
	(-10.0)	[-0.004, -0.003]
Relative Size	0.013***	0.011
	(9.0)	[0.008, 0.014]
$(M/B)^{Acquirer}$	-0.001	-0.001
	(-1.3)	[-0.002, 0.000]
$(Cash\ Flows/Assets)^{Acquirer}$	-0.000	-0.006
	(-0.1)	[-0.012, 0.000]
Diversify	-0.000	-0.001
	(-0.1)	[-0.003, 0.002]
High Tech	-0.003*	-0.004
	(-1.7)	[-0.007, -0.001]
Public Target	-0.027***	-0.033
	(-13.5)	[-0.037, -0.029]
Hostility	0.018***	0.023
	(4.3)	[0.015, 0.030]
Multiple Bids	0.001	0.003
	(0.1)	[-0.014, 0.022]
PureCash	0.001	0.001
	(1.3)	[0.000, 0.003]
Industry M&A	0.000	0.000
	(0.2)	[0.000, 0.001]
Industry dummies	YES	YES
Year dummies	YES	YES
Constant	YES	YES
Sample size / Adjusted R^2	25,645 / 0.05	500 replications

Table 6

Regressions of value-weighted portfolio returns of the acquirer and the target and takeover premium on explanatory variables

The sample consists of firms acquiring the majority shares of publicly listed targets in the U.S. between 1981 and 2008, reported in the SDC's Mergers and Acquisitions database. The abnormal returns are calculated using the market model. The market model is estimated by using 200 trading days of return data ending 6 days before the acquisition announcement. The CRSP equally weighted return is used as a proxy for the market return. In columns (1) and (2), the dependent variable is market-capitalization weighted portfolio cumulative abnormal returns of the acquirer and the target from one day before to one day after the announcement date of the M&A (*WCAR*(*-1,1*)). In columns (3) and (4), the dependent variable is the premium paid to the target by the acquirer (*Premium*). *Premium* is measured by the difference between the acquirer's offer price (total value of cash, stock, and other securities offered by the acquirer to the target) and the target's market value of equity 5 days prior to the M&A announcement date, scaled by the target's market value of equity on the same day. Firms are defined to have defined benefit (DB) plans if both the fair value of plan assets (*PPA*) and the projected pension benefit obligation (*PBO*) are available in Compustat. The pension funding deficit (*DEF/Assets*) is defined as the difference between *PBO* and *PPA*, deflated by the end-of-period book value of assets adjusted for pension items. All explanatory variables are measured at the fiscal year-end that immediately precedes the announcement date of share acquisitions. All variables are winsorized at the 0.5% level at both tails of the distribution. Columns (1) and (3) use the actual value of *DEF/Assets* in OLS regressions. The *t*-statistics in parentheses are calculated from the Huber/White/Sandwich heteroskedastic consistent errors, which are also corrected for correlation across observations for a given firm. The symbols ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively. Columns (2) and (4) report results from the three-step bootstrapping procedure outlined in the Online Appendix and use the predicted value of *DEF/Assets* obtained using Heckman's (1979) maximum likelihood estimator. The 95% confidence intervals in square brackets are calculated from 500 bootstrap replications of the three-step estimation based on resampling from the data set with replacement of clusters.

	Dependent variable: *WCAR(-1,1)*		Dependent variable: *Premium*	
	OLS	Three-step bootstrap	OLS	Three-step bootstrap
	(1)	(2)	(3)	(4)
DB	-0.002	0.002	-0.010	-0.010
	(-0.5)	[-0.007, 0.010]	(-0.6)	[-0.044, 0.026]
DB×DEF/Assets	0.221***	0.535	-0.705**	-1.030
	(3.3)	[0.299, 0.827]	(-2.5)	[-2.301, 0.266]
$Leverage^{Acquirer}$	0.018*	0.015	-0.073	-0.042
	(1.9)	[-0.005, 0.034]	(-1.6)	[-0.132, 0.046]
$Ln(Assets)^{Acquirer}$	-0.005***	-0.004	0.033***	0.035
	(-3.5)	[-0.007, -0.001]	(5.0)	[0.022, 0.049]
Relative Size	0.010***	0.009	0.040***	0.035
	(4.4)	[0.004, 0.014]	(3.6)	[0.015, 0.057]
$M/B^{Acquirer}$	-0.001	-0.001	0.011***	0.015
	(-1.2)	[-0.004, 0.002]	(3.2)	[0.007, 0.021]
$Cash\ Flows/Assets^{Acquirer}$	-0.000	0.007	0.020	0.048
	(-0.0)	[-0.008, 0.023]	(0.8)	[-0.045, 0.163]
Diversify	-0.007**	-0.005	-0.012	-0.001
	(-2.5)	[-0.011, 0.001]	(-0.9)	[-0.029, 0.027]
High Tech	-0.012**	-0.014	0.027	0.017
	(-2.4)	[-0.022, -0.006]	(1.1)	[-0.025, 0.059]
Public Target	0.010	0.008	0.066	0.073
	(1.3)	[-0.005, 0.021]	(1.5)	[-0.004, 0.148]
Hostility	0.028***	0.031	0.099***	0.093
	(7.2)	[0.025, 0.038]	(5.3)	[0.060, 0.127]
Multiple Bids	-0.012**	-0.015	0.006	-0.002
	(-2.1)	[-0.025, -0.005]	(0.2)	[-0.053, 0.042]
PureCash	0.022***	0.017	-0.078***	-0.093
	(7.2)	[0.011, 0.023]	(-5.0)	[-0.124, -0.060]
Industry M&A	0.001	0.001	0.006	0.008
	(0.9)	[-0.001, 0.004]	(1.2)	[-0.001, 0.017]
$Ln(Assets)^{Target}$	0.004***	0.003	-0.036***	-0.033
	(3.3)	[0.000, 0.005]	(-5.6)	[-0.047, -0.020]
M/B^{Target}	-0.001	-0.002	-0.018***	-0.023
	(-0.7)	[-0.004, -0.001]	(-4.5)	[-0.031, -0.015]
$Leverage^{Target}$	-0.008	-0.006	0.314***	0.351
	(-1.1)	[-0.019, 0.008]	(7.0)	[0.268, 0.435]

Industry dummies	YES	YES	YES	YES
Year dummies	YES	YES	YES	YES
Constant	YES	YES	YES	YES
Sample size / Adjusted R^2	3,515 / 0.13	500 replications	3,301 / 0.12	500 replications

Table 7

Regression of the fraction of cash payment on explanatory variables

The sample consists of firms acquiring the majority shares of publicly listed targets in the U.S. between 1981 and 2008, reported in the SDC's Mergers and Acquisitions database. In columns (1) and (2), the dependent variable is the fraction of cash payment in the transaction value (*Pct_Cash*). Firms are defined to have defined benefit (DB) plans if both the fair value of plan assets (*PPA*) and the projected pension benefit obligation (*PBO*) are available in Compustat. The pension funding deficit (*DEF/Assets*) is defined as the difference between *PBO* and *PPA*, deflated by the end-of-period book value of assets adjusted for pension items on the balance sheet. All explanatory variables are measured at the fiscal year-end that immediately precedes the announcement date of share acquisitions. All variables are winsorized at the 0.5% level at both tails of the distribution. Column (1) uses the actual value of *DEF/Assets* in the tobit regression. The *z*-statistics in parentheses are calculated from the Huber/White/Sandwich heteroskedastic consistent errors, which are also corrected for correlation across observations for a given firm. The symbols ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively. Column (2) reports results from the three-step bootstrapping procedure outlined in the Online Appendix and uses the predicted value of *DEF/Assets* obtained using Heckman's (1979) maximum likelihood estimator. The 95% confidence intervals in square brackets are calculated from 500 bootstrap replications of the three-step estimation based on resampling from the data set with replacement of clusters.

	Dependent variable: *Pct_Cash*	
	Tobit	Three-step bootstrap
	(1)	(2)
DB	-0.023	0.018
	(-0.3)	[-0.176, 0.202]
DB×DEF/Assets	3.965***	16.386
	(3.0)	[12.004, 21.640]
$Leverage^{Acquirer}$	-0.218	-0.488
	(-1.2)	[-0.838, -0.143]
$Ln(Assets)^{Acquirer}$	0.167***	0.205
	(5.3)	[0.146, 0.265]
Relative Size	-0.016	-0.019
	(-0.4)	[-0.095, 0.062]
$M/B^{Acquirer}$	-0.124***	-0.115
	(-5.0)	[-0.158, -0.070]
$Cash\ Flows/Assets^{Acquirer}$	0.853***	0.549
	(5.1)	[0.182, 0.875]
Diversify	-0.012	-0.053
	(-0.2)	[-0.165, 0.066]
High Tech	0.014	0.032
	(0.1)	[-0.135, 0.192]
Public Target	1.125***	1.056
	(4.3)	[0.578, 1.524]
Hostility	1.668***	1.646
	(19.4)	[1.496, 1.798]
Multiple Bids	0.470***	0.343
	(4.3)	[0.168, 0.523]

Industry M&A	0.016	0.033
	(0.7)	[-0.004, 0.076]
Ln(Assets)^{Target}	-0.202***	-0.195
	(-6.4)	[-0.251, -0.140]
M/B^{Target}	-0.103***	-0.101
	(-4.1)	[-0.147, -0.056]
Leverage^{Target}	-0.396**	-0.486
	(-2.5)	[-0.784, -0.180]
Industry dummies	YES	YES
Year dummies	YES	YES
Constant	YES	YES
Sample size / Pseudo R^2	3,815 / 0.20	500 replications

Table 8

Reestimation of regressions in Tables 4 through 7 for subsamples of firms classified according to firms' financial constraints, and subperiods, and by including mandatory pension contribution as an additional control for financial constraints

The sample consists of firms acquiring the majority shares of publicly listed targets in the U.S. between 1981 and 2008, reported in the SDC's Mergers and Acquisitions database. Firms are defined to have defined benefit (DB) plans if both the fair value of plan assets (*PPA*) and the projected pension benefit obligation (*PBO*) are available in Compustat. The pension funding deficit (*DEF*) is defined as the difference between *PBO* and *PPA*, deflated by the end-of-period book value of assets adjusted for pension funding status. All explanatory variables are measured at the fiscal year-end that immediately precedes the announcement date of share acquisitions. All variables are winsorized at the 0.5% level at both tails of the distribution. In Panel A, firms are divided into financially constrained (*FU*) and unconstrained (*FUC*) firms according to the median values of the Whited and Wu (2006) index. In Panel B, mandatory pension contribution (*MC*) is included in the regression to control for financial constraints. *MC* is measured as the ratio of pension expenses, as recorded in Compustat, to total assets if a firm's pension plan is underfunded, and zero if a firm's pension plan is fully funded or overfunded. Only *DB* firms are used in estimating the regressions in Panel B. In Panel C, the sample period is divided into pre- and post-1987 subperiods. Intercept and control variables are included in the estimation of regressions, but not reported in the table. The *t*- and *z*-statistics in parentheses are calculated from the Huber/White/Sandwich heteroskedastic consistent errors, which are also corrected for correlation across observations for a given firm. The symbols ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively.

Dependent variables	*Diversify* Probit (1)		*CAR(-1,1)* OLS (2)		*WCAR(-1,1)* OLS (3)		*Premium* OLS (4)		*Pct_Cash* Tobit (5)	
Panel A: Partitioning sample according to a firm's White and Wu (WW) index										
	FUC	FC	FUC	FC	FUC	FC	FUC	FC	FUC	FC
DB	0.039**	0.083***	0.000	0.001	-0.002	0.006	-0.009	-0.026	-0.100	0.161
	(2.3)	(4.8)	(0.1)	(0.6)	(-0.6)	(0.7)	(-0.5)	(-0.8)	(-1.2)	(1.3)
DB×DEF/Assets	-1.115***	0.038	0.054***	0.173***	0.205***	0.067	-0.846***	1.018	2.917**	3.156
	(-4.2)	(0.1)	(2.6)	(3.3)	(2.8)	(0.3)	(-2.9)	(1.4)	(2.1)	(0.8)
N	11,690	14,635	11,457	14,188	2,438	1,050	2,290	985	2,603	1,179
Panel B: Including mandatory pension contributions (MC)										
MC/Assets	4.705***		-0.151		0.113		0.402		-25.287**	
	(2.8)		(-0.9)		(0.2)		(0.2)		(-2.4)	
DEF/Assets	-1.290***		0.070***		0.175**		-0.793**		4.595***	
	(-4.7)		(2.8)		(2.3)		(-2.4)		(2.7)	
N	8,257		8,103		1,695		1,613		1,824	
Panel C: Subperiod analysis using Pension Protection Act of 1987 as a quasi-exogenous event										
	Pre-1987	Post-1987	Pre-1987	Post-1987	Pre-1987	Post-1987	Pre-1987	Post-1987	Pre-1987	Post-1987
DB	0.087**	0.077***	0.004	0.001	0.002	-0.001	-0.063	-0.010	-1.266**	0.080
	(2.6)	(5.8)	(0.9)	(0.7)	(0.1)	(-0.3)	(-1.1)	(-0.6)	(-2.5)	(1.2)
DB×DEF/Assets	-0.807	-0.487**	0.086	0.067***	-0.065	0.278***	-1.704**	-0.813**	-5.767	4.958***
	(-1.6)	(-2.0)	(1.4)	(3.3)	(-0.5)	(3.9)	(-2.6)	(-2.5)	(-0.9)	(3.5)
N	2,016	24,242	2,034	23,611	387	3,128	417	2,884	453	3,362

Table 9

Reestimation of regressions in Tables 4 through 7 for subsamples of defined benefit (DB) plan firms classified according to their collective bargaining status, industry unionization rate, and pension demographic characteristics

The sample consists of DB plan firms acquiring the majority shares of publicly listed targets in the U.S. between 1981 and 2008, covered in the SDC's Mergers and Acquisitions and IRS Form 5500 databases. Firms are defined to have DB plans if both the fair value of plan assets (*PPA*) and the projected pension benefit obligation (*PBO*) are available in Compustat. The pension funding deficit (*DEF*) is defined as the difference between *PBO* and *PPA*, deflated by the end-of-period book value of assets adjusted for pension funding status. All explanatory variables are measured at the fiscal year-end that immediately precedes the announcement date of share acquisitions. All variables are winsorized at the 0.5% level at both tails of the distribution. In Panels A1 and A2, firms are divided into active-employee dominated (*Active*) and retiree dominated (*Retired*) firms according to a firm's pension plan age and the faction of actively working employees in the pension plan, respectively. In Panel A1, a firm is classified as *Active* if its plan age is below the sample median, and *Retired* otherwise. In Panel A2, a firm is classified as *Active* if the fraction of actively working employees is above the sample median, and *Retired* otherwise. In Panel B, firms are divided into two subsamples according to whether their pension plans are collectively bargained as reported in the IRS Form 5500. In Panel C, firms are divided into two subsamples according to the sample median of industry unionization rate (*Union*). Intercept and control variables are included in the estimation of regressions, but not reported in the table. The *t*- and *z*-statistics in parentheses are calculated from the Huber/White/Sandwich heteroskedastic consistent errors, which are also corrected for correlation across observations for a given firm. The symbols ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively.

Dependent variables	*Diversify* Probit (1)		*CAR(-1,1)* OLS (2)		*WCAR(-1,1)* OLS (3)		*Premium* OLS (4)		*Pct_Cash* Tobit (5)	
	Panel A: Partitioning the sample according to pension demographic characteristics									
	Active	Retired	Active	Retired	Active	Retired	Active	Retired	Active	Retired
	Panel A1: Partitioning the sample according to the age of pension plans									
DEF/Assets	-0.987**	-0.462	0.071*	0.046*	0.546**	0.091	-0.530	-0.652	7.360**	-0.375
	(-2.3)	(-1.5)	(1.7)	(1.7)	(2.4)	(1.4)	(-0.8)	(-1.4)	(2.4)	(-0.3)
N	4,122	3,173	4,065	3,217	834	642	796	578	888	683
	Panel A2: Partitioning the sample according to the percentage of active employees in the pension plan									
DEF/Assets	-0.830*	-0.262	0.095**	0.040	0.190*	0.076	-1.727**	-0.974	3.564*	1.379
	(-1.7)	(-0.6)	(2.4)	(0.9)	(1.8)	(0.6)	(-2.4)	(-1.3)	(1.7)	(0.6)
N	1,690	1,815	1,714	1,838	414	363	392	335	428	381
	Panel B: Partitioning the sample according to whether a firm sponsors a collectively bargained DB plan									
	Yes	No	Yes	No	Yes	No	Yes	No	Yes	No
DEF/Assets	-0.878*	0.179	0.068*	0.017	0.121	0.009	-1.750**	-0.918	3.321**	2.132
	(-1.8)	(0.4)	(1.7)	(0.4)	(1.0)	(0.1)	(-2.3)	(-1.0)	(2.3)	(0.7)
N	1,385	2,121	1,413	2,140	267	511	239	488	276	534
	Panel C: Partitioning the sample according to median unionization rate (Union)									
	High	Low	High	Low	High	Low	High	Low	High	Low
DEF/Assets	-0.929***	-0.404	0.075***	0.085	0.176**	0.330	-1.056***	0.339	2.631**	5.030
	(-3.7)	(-0.8)	(3.3)	(1.5)	(2.2)	(1.6)	(-3.1)	(0.6)	(2.1)	(1.3)
N	5,381	2,901	5,303	2,844	931	782	878	750	1,016	830

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1 The banking literature suggests that as inside debtholders, banks have a competitive advantage over other capital market participants in collecting information about the borrowing firms (Fama (1985), Rajan (1992)) and thus can intervene quickly and informally.

2 A DB pension plan commits the sponsoring firm to pay a pre-specified benefit to its employees at a given future date. Under such a plan, if the value of pension assets is insufficient to pay the promised benefits, the sponsoring firm is responsible for the shortfall (pension deficit). This commitment is financially equivalent to a firm's legal promise to pay off conventional debt on its balance sheet.

3 The aggregate pension deficits are computed based on all U.S. firms with a DB pension plan covered in Compustat.

4 Francis and Reiter (1987) maintain that underfunded DB pension plans make employees as sponsoring firms' debtholders. Similarly, Edward Burrows, the former president of the American Society of Pension Professionals & Actuaries, argues that unpaid accrued pension benefit obligations in DB plans are loans from plan participants to sponsors. He suggests that "…These loans (pension deficits) differ from most commercial loans in an important respect: the lenders (the employees) are not establishing diversified loan portfolios. They must deal with just one borrower: the plan sponsor…"

5 The maximum pension benefit guaranteed by the PBGC is adjusted annually. For single-employer plans that end in 2011, workers who retire at age 65, 60, and 50 can receive up to $4,500, $2,925, and $1,575 per month, respectively. The PBGC does not guarantee benefits for which employees do not have a vested right or have not met all age, service, or other requirements at the time when the plan terminates. Benefit increases and new benefits that have been in place for less than five years are also only partly guaranteed.

6 Anecdotal evidence also indicates that pension deficits are far from fully guaranteed by the PBGC. For example, Reuters reports that "…the underfunded liability was estimated at $41 billion for GM, Chrysler and Ford Motor Co. at the end of 2008, the latest PBGC figures show. GM accounted for half the shortfall and only $4 billion of that gap would be insured if plans were terminated now, according to the PBGC. Chrysler had a $9 billion shortfall of which $2 billion would be covered." (Reuters, April 23, 2009).

7 For example, in 2003, the annual insurance premium was $19 per employee, plus $9 per $1,000 of shortfall. In 2009, it was increased to $34 per employee.

8 Unlike conventional debt whose periodic interest payments are regular and predictable, the pension contributions made by sponsoring firms are unstable and difficult to predict since they are influenced by various factors such as pension funding status, the availability of firms' cash flows, and managerial discretion. However, a significant part of pension contributions such as mandatory pension contributions and insurance premiums paid to the PBGC are regular and nondiscretionary and thus can serve as an important control mechanism in limiting managers'

discretionary behavior.

[9] The sample period is constrained by the availability of the Compustat data items related to DB pension plans, which are available starting from 1980. Since our empirical analysis uses one-year lagged pension variables in regressions, our sample period starts from 1981.

[10] U.S. Internal Revenue Service (IRS) Form 5500 is an alternative data source for a firm's pension-related variables, as it offers detailed information concerning plans' finances, participants, and administrators. However, using Form 5500 would significantly reduce our sample size, since complete filings are available only for fiscal years 1990 through 1998.

[11] To measure Assets, we follow the netting approach described in Table 1 of Shivdasani and Stefanescu (2010). Our results remain qualitatively the same if we use the unadjusted book value of assets as in Rauh (2006a).

[12] Untabulated results show that compared with acquiring firms sponsoring non-DB pension plans, DB plan sponsors are larger, older, and more leveraged. DB plan sponsors also have more employees, more tangible assets, lower growth opportunities, and higher profitability,

[13] Because many targets are privately held and their stock prices are not available, the sample sizes in calculating WCAR and Premium are reduced to 3,515 and 3,301, respectively.

[14] Previous studies show that stock-financed M&As, diversifying M&As, hostile M&As, and acquisitions of publicly listed firms generally result in lower bidder returns. See, for example, Morck, Shleifer, and Vishny (1990), Schwert (2000), and Fuller, Netter, and Stegemoller (2002) for detailed results.

[15] Data on the unionization rate, which is the percentage of unionized workers in an industry, come from the Union Membership and Coverage Database compiled by Hirsch and Macpherson (2003).

[16] To further account for the impact of merger waves on announcement returns, we control for the combined industry/year fixed effects in the regressions instead of controlling for industry and year fixed effects, separately. Untabulated results show that our results still hold.

[17] Since a firm may make more than one acquisition in a given year, the number of acquirers (23,072) is smaller than the number of acquisitions (26,325) in our sample.

[18] We find that DB is positively related to the likelihood of making diversifying mergers. The marginal effect of changing from a non-DB firm to a DB plan sponsor increases the likelihood of making diversifying mergers by 7%. Column (2), which is estimated using the bootstrapped approach, also shows that this effect is significant at less than the 5% level.

[19] In untabulated tests, using negative binomial regressions, we examine the impact of pension deficits on the number of diversifying mergers and find similar results.

[20] Using either the abnormal return on the announcement day (CAR(0)) or the cumulative abnormal return from two days before to two days after the announcement date (CAR(-2,2)) as the dependent variable does not change our main results.

[21] Masulis, Wang, and Xie (2007) find that for their sample of acquiring firms during the 1990 to 2003 period, a one-standard deviation increase in leverage is associated with an approximately 0.09% increase in acquirer CAR(-2,2).

[22] Masulis, Wang, and Xie (2007) show that the quality of corporate governance has a significant effect on acquirer returns. To check whether our results are robust to the inclusion of the governance index, we reestimate the regressions in Table 5 by adding G-index constructed by Gompers, Ishii, and Metrick (2003) as an additional control variable. Untabulated results show that the coefficient estimate on DB×DEF/Assets remains positive and significant in both OLS regressions and regressions using the bootstrapping procedure.

[23] Because we include target characteristics in regressions reported in Tables 6 and 7, our tests in these tables are conducted over a smaller sample for which target characteristics are available in Compustat. In untabulated tests, we also estimate the regressions in Tables 6 and 7 using the full sample of acquisitions and not including target characteristics, and find that the coefficient estimates on DB×DEF/Assets are positive and significant at the 1% level in all regressions.

[24] In untabulated tests, we estimate probit regressions in which the dependent variable (PureCash) is a dummy variable that takes a value of one if the acquirer pays for acquisitions using cash only and zero otherwise. The results are similar to those reported in Table 7.

[25] Inconsistent with this view, however, Franzoni and Marin (2006) and Franzoni (2009) show that firms with underfunded pension plans tend to be overvalued and have lower stock returns than those with overfunded pension

[26] However, some of the results in our paper, such as the findings that acquiring firms with larger pension deficits realize higher announcement returns and are more likely to use cash to purchase targets, are consistent with the disciplinary effects of pension deficits rather than financial constraint effects.

[27] Although both the Retirement Protection Act of 1994 and the Pension Protection Act of 2006 also increase the levels of deficit reduction contributions of severely underfunded pension plans, compared with the Pension Protection Act of 1987, the incremental increase in contribution requirements imposed by these acts for firms with large pension deficits is substantially smaller.

[28] Since the fraction of active employees is highly correlated with firm size, firm age, and the size of pension plan, to purge out the size and age effects, we first regress the fraction of active employees on firm size, firm age, and PBO/Assets. We then use the residual from this regression as a measure for the fraction of active employees. Following Rauh (2009), we obtain information on the numbers of active and retired employees from the Form 5500 data.

[29] In Panel A1 of Table 9, the sample size for Active is larger than that for Retired because we include firms whose plan age is equal to the sample median in the Active subsample. Including these firms in the Retired subsample does not change our results.

[30] Pension asset reversion is an acquiring firm's act to redistribute wealth from target employees to its shareholders by terminating the target's overfunded pension plan or replacing the overfunded plan with a plan that offers lower payments to target employees. .

[31] This result is somewhat different from the evidence documented by Pontiff, Shleifer, and Weisbach (1990), who show that the reversions can explain approximately 11% of the takeover premium in cases where they actually occur. Although our result is not directly comparable to theirs since our result is based on the multivariate regression analysis while their analysis is derived from the univariate analysis, we propose two possible explanations for this difference in results. First, while their sample period covers only from 1981 to 1988, the most of our sample period covers the period after 1986. After the passage of the Tax Reform Act of 1986, there was a sequence of escalating reversion tax rates from 10% in 1986 to 50% in 1990 on reversion amounts, significantly weakening the importance of pension asset reversions as a source of takeover gains. Second, Pontiff, Shleifer, and Weisbach (1990) only look at tender offers that are reported in the Wall Street Journal, while we include all M&As in which acquiring firm owns less than 50% of the target's shares before the announcement date and controls 100% of the target's shares after acquisition.