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ARTICLE

Financial health and the valuation of corporate pension plans

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Abstract
We return to the long-standing question 'Who owns the assets in a defined benefit pension plan?' Unlike earlier studies, we condition the market's assessment of implicit property rights on the sponsoring firm's financial health. Valuations of financially strong firms, and those that are strengthening, are more responsive to pension plan funding. For these firms, each extra dollar of net plan assets is valued at between \$0.50 and \$1.00. In contrast, for weak and weakening firms, valuation effects are statistically indistinguishable from zero. This result is consistent with the higher likelihood that they will renege on their pension obligations.

Key words: Bankruptcy scores; defined benefit pension plans; financial distress; property rights; stock market valuation effects

JEL codes: G23; G32

U.S. corporate defined benefit pension plans have received considerable recent attention from both practitioners and academic researchers, and for good reason. Pension assets are substantial, averaging roughly 1/6 of total firm value. Among about 2,000 firms that sponsor defined benefit plans, average total plan assets exceed \$1,500 billion and average total projected benefit obligations exceed \$1,600 billion.

In addition to the importance conveyed by their size, corporate pension plans involve some interesting and complex financial contracts. One long-standing issue is the question 'Who owns the assets in a defined benefit pension plan?' (Bulow and Scholes, 1983). Prior to the passage of the Employee Retirement Income Security Act (ERISA) in 1974, pension liabilities were not liabilities of the firm. At plan termination, beneficiaries or employees had claims on the assets of the pension fund, but no claims on other corporate assets. However, under ERISA, pension assets and liabilities are treated as corporate assets and liabilities; this is explicitly recognized by the Financial Accounting Standards Board (FASB) in its statement SFAS 87. On the other hand, ERISA shifted some of the firm's contingent obligations onto the government, specifically the Pension Benefit Guaranty Corporation (PBGC). If the company is in distress and pension funding is inadequate to support promised retirement benefits, the PBGC's insurance program guarantees those benefits up to specified limits. This implicit put option implies that the government has a stake in – and implicit property rights to – the firm's pension assets and liabilities.

A number of empirical studies (Oldfield, 1977; Feldstein and Seligman, 1981; Feldstein and Morck, 1983; Daley, 1984; Dhaliwal, 1986; Landsman, 1986; Bodie *et al.*, 1987; Bulow *et al.*, 1987; Barth, 1991) have attempted to infer the implicit property rights to pension plan assets and liabilities. These studies commonly entail regressions of firm value on pension plan assets and liabilities. The typical empirical question is whether the estimated coefficients on pension assets or liabilities in firm valuation regressions are close to 1.0 in absolute value. Feldstein and Seligman (1981) find that while each dollar increase in pension liabilities reduces equity value by about a dollar, a corresponding increase in pension assets increases equity value by less than a dollar. These results support the hypothesis that stock market investors treat the liabilities, but not the assets, of pension plans as fully internalized by the firm. This is consistent with the view that firms are responsible for underfunded liabilities but may have difficulty in accessing assets of overfunded plans for general corporate purposes. However, a more recent study by Liébana and Vincent (2004) finds no consistent evidence that investors value firms' pension assets and liabilities as though they belonged to the sponsoring firms. The valuation implications of pension plans remain unanswered.

Unlike earlier studies, we explicitly consider in this paper the implications of the firm's financial health on the valuation effects of pension plan funding. To the extent that firms may eventually walk away from their obligations, other parties (e.g., the PBGC as well as incompletely insured employees) may be partial claimants on the assets and liabilities of the plan. This will attenuate the response of firm value to pension funding. We therefore distinguish among firms based on their financial condition.

Firms may behave strategically in allocating pension assets and choose to 'go for broke' (Harris and Raviv, 1992) when they approach financial distress to maximize the value of the PBGC put option. Besides increasing the total value of the pension put, such risk-shifting behavior can reallocate the values of the contingent claims (and liabilities) across the various stakeholders.[1]

Moreover, flexible actuarial choices make pension valuation soft and discretionary (Scholes and Wolfson, 1992). Several papers show that firms' actuarial choices, such as pension discount rate, cost method, and the forecast salary growth rate vary with the financial condition of the firm (Morris *et al.*, 1983; Bodie *et al.*, 1987; Ghicas, 1990; Gopalakrishnan and Sugrue, 1995; Godwin *et al.*, 1996; Petersen, 1996; Asthana, 1999). As a general rule, firms appear to choose conservative pension assumptions in good times and aggressive ones in bad times. All these factors, i.e., the PBGC put option, strategic risk-taking, and discretionary actuarial choices, call for pension valuation models that are conditioned on financial status.

We capture the level and change in firms' financial conditions over consecutive fiscal years using bankruptcy-risk scores developed in three studies: Ohlson (1980), Altman (1968), and Campbell *et al.* (2008). We use *Ohlson_BS*, *Altman_Z*, and *CHS_BS* to denote the corresponding bankruptcy scores and Δ*Ohlson_BS*, Δ*Altman_Z*, and Δ*CHS_BS* to denote changes in firms' financial status between consecutive fiscal years. (In all cases, we add minus signs to the changes in bankruptcy-risk scores, so that positive numbers indicate improvements in financial condition.)

Taking Ohlson's (1980) bankruptcy score as an example, we use the estimated parameters from his model to calculate *Ohlson_BS* for each firm in our sample. Then, we compute the difference between the *Ohlson_BS* during consecutive fiscal years. The resulting Δ*Ohlson_BS* index measures the change in firms' financial condition. Depending on whether Δ*Ohlson_BS* is positive or negative, we partition firms into healthier versus more distressed groups and estimate a pension valuation model for the two groups. We ask how changes in various items on the pension fund balance sheet affect changes in the market valuation of the firm, and whether those valuation effects depend on the firm's financial status.

Aside from conditioning valuation effects on financial status, the other major novelty in our approach is a focus on the *changes* in market value that result from changes in pension funding. Earlier studies have focused almost exclusively on the level of plan assets and obligations. The valuation measures employed by these studies typically include the ratio of market value to book value of assets (Tobin's *Q*), the ratio of market value to book value of equity (*M*/*B*), or the market value

[1]The extent to which this happens in reality is an open question. Rauh (2009) concludes that firms, in fact, actually assume less risky pension asset allocations when their financial condition is weaker. The broader point is that pension plan risk may respond endogenously to the firm's financial condition, with implications for ownership claims.

of equity normalized by other deflators. However, these multiples, which are used to assess the impact of pension funding on the market valuation of the firm, are relatively crude benchmarks. Moreover, because levels of balance sheet items are highly persistent from year to year, time series variation in pension funding and financial condition does not provide much power to tease out corresponding variation in valuation multiples, especially given the imprecision in the benchmarks.

Our focus on changes rather than levels offers several advantages. First, changes in stock market value are observed as contemporaneous rates of return, and we have a rich set of tools to net out the impact of broad market movements to obtain cleaner estimates of the firm-specific return in any period. By controlling more precisely for benchmark returns, we enhance the signal-to-noise ratio and improve the power of our empirical tests. Second, an excess-return valuation method allows for an adjustment for the difference in risk factors and, therefore, the difference in discount rates on firm valuation (Faulkender and Wang, 2006).

Third, whereas variation in pension funding may have only a second-order impact on valuation multiples, they are far more telling for changes in market value. Changes in pension net funding will have a far greater proportional impact on equity value (which will show up as an excess return) than on variation in any valuation multiple.[2] Therefore, we adopt a long-run benchmark-adjusted stock return approach, as in Grinblatt and Moskowitz (2004) and Faulkender and Wang (2006). The dependent variable is a stock's excess return relative to the matched size and book-to-market portfolio of Fama and French (1992, 1993).

The focus on changes rather than levels poses one potential problem in interpretation. For consistency with our other variables, we partition firms based on *change* in financial status, but are ultimately interested in the difference between healthy and distressed firms. However, we demonstrate that changes in bankruptcy scores display considerable serial correlation: on average, strengthening firms in 1 year continue to strengthen in the following year, and weakening firms continue to weaken. Therefore, it would be reasonable for the market to conclude that firms with currently worsening financial metrics are more likely to encounter serious challenges in the future, and therefore present a higher likelihood of offloading pension obligations onto other parties. Moreover, when we partition on the level of distress score, we obtain results broadly consistent with, albeit less powerful than, results based on partitions using changes.

We summarize our major results as follows. First, using all firm-year observations from June 1988 to June 2017, we find the following point estimates, given in the 'All firms' column, for the impact on firm value changes in various pension assets or liabilities:

		Healthier firms, partitioned by:		
Valuation impact of an extra dollar of:	All firms	Δ*Ohlson_BS*	Δ*Altman_Z*	Δ*CHS_BS*
Plan assets	\$0.43	\$0.65	\$0.66	\$0.61
Projected benefit obligations	−\$0.28	−\$0.48	−\$0.55	−\$0.34
Accumulated benefit obligations	−\$0.58	−\$0.95	−\$0.98	−\$0.87
Funding status (plan assets – projected benefit obligations)[3]	\$0.37	\$0.52	\$0.50	\$0.51
Off-balance sheet pension items	\$0.34	\$0.51	\$0.48	\$0.49

Source: Excerpted from Tables 2 and 7.

These results thus offer new evidence suggesting that pension fund assets and liabilities are valued by the market much less than one-for-one. Moreover, the market seems to value an extra dollar of accumulated benefit obligations more highly than an extra dollar of projected benefit obligations. Investors seem to recognize off-balance sheet pension liabilities as well. Not only is the estimate large in magnitude, but the significance level is high.

[2]For example, the change in value of Hewlett-Packard's 2013 pension assets was about 3% of the firm's total market capitalization. This change induces a first-order impact on excess return, but not on the ratio of pension assets to total capitalization. Pension assets were about 58% of HP's market capitalization in that year.

[3]Following SFAS 158, funding status is determined by comparing PBO, not ABO, to plan assets.

Second, the valuation impact of pension funding differs considerably by financial health. The three columns to the right show that for all three indices of financial strength, the valuation effects for healthier firms are greater than those for the entire sample of firms. This result suggests that pension funding has stronger valuation effects on financially secure firms.

In contrast, for firms that have become financially weaker, valuation effects (not shown in this excerpt) are much smaller; in fact in the majority of cases, they are statistically indistinguishable from zero. Below, we present formal tests of the hypothesis that the valuation effects are the same for healthy and distressed stocks. We firmly reject the null hypothesis for each of the five pension asset and liability variables above, regardless of which distress-risk score is used to partition the sample. These patterns are even stronger following the adoption of pension accounting rules that increased disclosure requirements.

While we employ several measures of financial strength, our results are robust to the particular measure used to partition the sample. This is in part because the alternative indices of the evolution of financial status (Δ*Ohlson_BS*, Δ*Altman_Z*, and Δ*CHS_BS*) are themselves generally consistent. The pairwise correlations between the three are all positive, ranging from 0.29 to 0.47, and are all statistically significant at better than a 5% level.

Third, both improvements and deterioration in financial health tend to persist. On average, firms in the improving or *STRONGER* group in 1 year exhibit further reductions in financial risk in the next year. Conversely, firms in the *WEAKER* group exhibit higher risk in the next year. This pattern implies that forecasts of future financial distress can be sensitive to recent changes in bankruptcy-risk scores. This is consistent with our finding that valuation effects are highly associated with such changes.

Fourth, we estimate the valuation impacts of mandatory contributions to pension plans. While overfunded plans do not have to make contributions, firms operating underfunded plans (for which projected benefit obligations exceed pension assets) are required by law to make catch-up contributions. We construct two measures of mandatory contributions following Moody (2006) and Campbell *et al.* (2012). The first, *Mand*1, equals service cost (retirement benefits accrued by plan participants during the year) plus one-thirtieth of current underfunding, measured as the difference between accrued benefits and plan assets. The second, *Mand*2, is simply that year's service cost. (If either statistic is negative, e.g., due to plan overfunding, the variable is set equal to zero.) The stock market reacts to either measure of mandatory contributions with a sensitivity far exceeding one-for-one. Every incremental dollar of mandatory contributions reduces stock market value by more than \$5 for *STRONGER* firms versus roughly zero to \$2 for *WEAKER* firms. We interpret these results as evidence that a change in mandatory contributions is generally long lasting, and the market discounts the present value of the entire stream of required future contributions. For distressed firms, for which there is a greater likelihood that the stream of contributions will be interrupted, the impact is significantly lower.

Fifth, we find that small firms, which generally are less likely to hedge uncertainty (Nance *et al.*, 1993), are more sensitive to variation in pension valuation. We carry out a two-way partition based on both market capitalization and each of the three distress-risk indices. The partition divides all firms into the following four interactive groups: large and healthier, large and more distressed, small and healthier, and small and more distressed. We find the differential valuation effects between healthier and more distressed firms are stronger for small firms than large firms.

Our work is also related to the accounting literature on pension valuation focusing on the relevance of a fair or market value model versus that of a smoothed model under SFAS 87, the pension accounting recognition and measurement rule instituted by the Financial Accounting and Standards Board. The accounting literature is typically based on empirical variants of the Ohlson (1995) model that provide a direct link between accounting measures and firm value. These studies include Daley (1984), Landsman (1986), Barth (1991), Barth *et al.* (1992), Coronado and Sharpe (2003), Coronado *et al.* (2008), and Hann *et al.* (2007). Glaum (2009) provides a detailed review of the value relevance of pension accounting information. There is also a growing related literature on the management of pension assumptions and actuarial choices (Thomas, 1988; Thomas and Tung, 1992; Blankley and Swanson, 1995; Amir and Gordon, 1996; Amir and Benartzi, 1998; Bergstresser *et al.*, 2006).[4]

The rest of the paper proceeds as follows. Section 1 develops our main hypothesis. Section 2 describes data sources, sample screening, and variable definitions. Section 3 presents summary statistics. Section 4 provides the baseline valuation for corporate pension plans. Section 5 constructs three bankruptcy scores used to develop partitioning indices. Section 6 examines the persistence of firms' changes in financial condition. Section 7 estimates the valuation model for firms' pension assets and liabilities, conditional on recent changes in financial condition. Section 8 examines the conditional valuation of mandatory contributions. Section 9 explores the role of firm size in pension valuation. Finally, Section 10 concludes.

1. Hypothesis development

ERISA has important implications regarding the property rights of corporate pension plans. Since its implementation, researchers have tended to view pension assets as corporate assets and pension liabilities as corporate liabilities (Sharpe, 1976; Treynor, 1977; Barnow and Ehrenberg, 1979; Tepper, 1981; Bulow, 1982). Nevertheless, both are subject to a few important special considerations (Bulow *et al.*, 1983).

First, ERISA created the PBGC, which insures obligations to plan beneficiaries. If distressed firms terminate their plans, the PBGC pays the difference between (i) the guaranteed benefit and (ii) the assets in the pension funds plus 30% of the firms' net worth. Employers pay insurance premiums against plan termination. Second, ERISA restricts the transfer of pension plan assets to the corporation, making it costly for firms to overfund pension plans and later pull back the funds. In addition, the Tax Reform Act of 1986 introduced an excise tax on reversion of pension plan assets. These considerations imply that the ownership of net pension assets is only partially retained by the firm.

We will focus primarily on the financial status of the firm. Regardless of the current level of pension funding, financially healthy firms cannot simply walk away from pension obligations; the PBGC put comes into play only in the event of bankruptcy. Pension obligations, like other prospective liabilities, are more likely to be internalized by shareholders when the firm is healthier and more likely to honor them, so this is the variable on which we focus. For example, an increase in the value of pension assets due to strong market returns should increase the value of a financially strong firm that is likely to pay its pension obligations by more than the value of a distressed firm that is viewed as likely to renege on those obligations in any event. Therefore, the response of firm value to the value of pension funding should depend on the firm's financial status. This suggests these two hypotheses:

Hypothesis 1: Only a fraction of the change in net pension assets should be internalized in the market's assessment of a firm's intrinsic value.

Hypothesis 2: Pension valuation effects should be stronger for financially healthy firms than for those vulnerable to financial distress.

Nevertheless, *conditional* on financial distress, pension funding will matter, as shown in Bulow and Scholes (1983). In practice, of course, pension funding tends to be correlated with financial strength, so these effects are intertwined. We will present evidence on the interaction of these two variables.

[4]Our work complements recent studies that focus on other aspects of defined benefit corporate pension plans. These include studies on determinants of corporate pension funding strategies (Francis and Reiter, 1987; Ippolito, 2001); reversions of excess pension assets (Pontiff *et al.*, 1990; Petersen, 1992); equity risk and pension plans (Jin *et al.*, 2006); information efficiency and pension accounting information (Franzoni and Marín, 2006; Picconi, 2006); the impact of mandatory contributions on corporate investment (Rauh, 2006); stock market reaction to pension contributions (Franzoni, 2009); the relation between pension plans and capital structure (Shivdasani and Stefanescu, 2010); the effect of pension plan funding on cost of capital (Campbell *et al.*, 2012); and asset allocation and managerial assumptions in pension plans (Addoum *et al.*, 2010).

2. Data sources, sample screening, and variable definitions

2.1 Data sources and sample stocks

The data for U.S. equity markets are from the CRSP and COMPUSTAT merged files. We obtain monthly returns, monthly stock prices, and market capitalization from CRSP. The annual accounting items, such as fiscal year-end shares outstanding and book value of equity, and pension related variables, such as plan assets and projected benefit obligations, are taken from COMPUSTAT. We use NYSE, AMEX, and NASDAQ firms, excluding financial firms with 4-digit SIC codes between 6000 and 6999.

2.2 Sample period and FASB statements

Our sample period begins in June 1988 when SFAS 87 imposed new standards on pension reporting. SFAS 87 (FASB, 1985) requires that accumulated benefit obligations determine recognition of minimum liability and dictates a smoothed rather than a fair or market-value model for pension accounting. Under SFAS 132 (FASB, 1988), effective in December 1997, firms are no longer required to report separate pension items for over- and under-funded plans. Under SFAS 158, effective as of December 2006 (FASB, 2006), firms must incorporate fair value funding status, or the difference between plan assets and projected benefit obligations, in their consolidated statements. Minimum pension liability adjustments associated with accumulated benefit obligations under SFAS 87 are no longer required. The sample ends in June 2017.

2.3 Variable definitions

Our variables fall into four groups: pension variables, raw and benchmark-adjusted excess stock returns, accounting variables, and financial strength measures. We discuss the first three categories in this section. Pension-plan related variables include plan assets (*PA*), projected benefit obligations (*PBO*), accumulated benefit obligations (*ABO*), funding status (*Funding_Status*), off-balance sheet items (*OFF_BAL*), and two measures of mandatory contributions (*Mand*1 and *Mand*2). These are all scaled by beginning-of-fiscal-year market capitalization. The details of the construction of these variables are provided in Appendix A. The measures of financial strength are treated in Section 5.

Plan assets (*PA*) refer to funds set aside to meet a firm's obligations. They increase due to capital gains on existing assets as well as from the difference between firm contributions and benefit payouts. *PBO* is the present value of employees' projected future benefits, which requires firms to make several actuarial assumptions, for example, concerning number of years until retirement, post-retirement life expectancy, final salary, and the appropriate discount rate. Whereas *PBO* is based on expected future salaries, *ABO* calculates employees' future benefits using their current salaries; it is the current value of obligations already earned, and equal to the present value of benefits if the plans were terminated immediately.

Funding_Status is the difference between *PA* and *PBO*. According to SFAS 158, this difference must be immediately incorporated into the balance sheet. However, until 2006, U.S. GAAP kept some pension gains and losses off the firm's financial statements to smooth the periodic pension cost. Off-balance-sheet pension assets or liabilities (*OFF_BAL*) included unrecognized gains and losses (*Unreg_GL*), unrecognized prior service costs (*Unreg_SC*), and unrecognized transitional assets and liabilities (*Unreg_TAL*). *Unreg_SC* measures retroactive benefits awarded to employees due to plan amendments. COMPUSTAT integrates *Unreg_GL* and *Unreg_TAL* together into one number. It estimates changes in *PBO* due to changes in actuarial assumptions and the deferred gains and losses that result from the difference between expected and actual returns on plan assets.[5]

Following the literature, we control for risk and macroeconomic factor exposure using the 25 size and book-to-market portfolios of Fama and French (1992, 1993). Liew and Vassalou (2000) show that these factors predict GDP growth, and may therefore control for some aspects of business cycle risk. To the extent that market returns depend on interest rate innovations, the FF factors at least partially control for that

[5]Appendix A of Picconi (2006) offers a clear overview of pension accounting.

source of risk as well. Stock *i*'s excess return in year *t* is its return minus that of the benchmark portfolio to which it belongs at the beginning of the fiscal year. We calculate annual returns (including dividends) for the fiscal year by cumulating monthly returns for both individual stocks and the 25 FF portfolios.

Our regression model for valuation of pension variables incorporates the following ten control variables; for each, we calculate changes between consecutive fiscal years of: cash holdings (*CASH*); interest expenses (*INT*); earnings before interest and taxes (*EBIT*); and non-cash total assets (*Non-Cash_Assets*). We adjust *EBIT* for pension cost (*PC*) by adding back the pension cost that is typically deducted. We also control for beginning-of-period cash holding ($CASH_{t-1}$); accruals (*ACCRUALS*); external financing (*XFIN*), which includes both debt and equity financing; asset growth (*ASST_GWTH*); market leverage (*MLEV*); and beginning-of-fiscal year firm size (Mkt_Eq_{t-1}). We deflate the firm-specific factors (except for asset growth, market leverage, and firm size) by the 1-year lagged market value of equity (Mkt_Eq_{t-1}). This standardization enables us to interpret the estimated coefficients as the dollar change in value for a \$1 change in the corresponding independent variable. The details of the construction of these variables using COMPUSTAT items are provided in Appendix A.

3. Summary statistics

The sample for NYSE/AMEX/NASDAQ covers 348 months, from June 1988 to June 2017. There are 2,737 stocks for which the necessary pension variables are available in this period. While we have 23,446 firm-year observations for ΔPA, ΔPBO, $\Delta Funding\_Status$, and ΔOFF_BAL, there are fewer firm-year observations for ΔABO because disclosure of *ABO* was not required between 1999 and 2003. The number of firms in our sample in each year ranges from 606 in fiscal year 2016 to 997 in 1994. However, these are not the same firms in each year, so 2,117 distinct firms appear at least once in the sample. Panel A of Table 1 provides summary statistics, including mean, median, 25th and 75th percentile values, and standard deviations. Panel B presents pairwise correlations of key variables.

Panel A shows that the average increase in pension assets is 1.6% and the average increase in projected benefit obligations is 1.8%, resulting in an average annual change in funding status of −0.3% of the beginning of the fiscal year market value. The mean annual stock return in excess of the Fama and French (1992, 1993) size and book-to-market adjusted benchmark is 1.4%. This is roughly the same magnitude as the average change in pension-cost adjusted earnings before interest and taxes, 1.7%.

Panel B reveals that ΔPA has significant correlations of 0.43, 0.70, and 0.74, respectively, with ΔPBO, ΔABO, and $\Delta Funding\_Status$. ΔABO and ΔPBO are highly correlated (0.81), which is not surprising. However, the positive correlation (0.30) between ΔABO and $\Delta Funding\_Status$ combined with the negative correlation between ΔPBO and $\Delta Funding\_Status$ (−0.21) suggests that pension assets track accumulated benefit obligations more closely than projected benefit obligations.

4. Baseline valuation

4.1 Pension assets and liabilities

Table 2 presents estimates of the following baseline valuation model for the entire sample of firms, in which the dependent variable $R_{i,t} - R_{B,i,t}$ is the return of stock *i* during fiscal year *t*, net of the return on the benchmark portfolio (matched by size and book-to-market ratio):

$$\begin{aligned} R_{i,t} - R_{B,i,t} = {} & \alpha_0 + \alpha_1 \Delta Y_{i,t} + \gamma_1 \frac{CASH_{i,t-1}}{Mkt\_Eq_{i,t-1}} + \gamma_2 \frac{\Delta CASH_{i,t}}{Mkt\_Eq_{i,t-1}} + \gamma_3 \frac{\Delta INT_{i,t}}{Mkt\_Eq_{i,t-1}} \\ & + \gamma_4 \frac{\Delta(EBIT + PC)_{i,t}}{Mkt\_Eq_{i,t-1}} + \gamma_5 \frac{Non{-}Cash\ Assets_{i,t}}{Mkt\_Eq_{i,t-1}} + \gamma_6 \frac{ACCRUALS_{i,t}}{Mkt\_Eq_{i,t-1}} \\ & + \gamma_7 \frac{XFIN_{i,t}}{Mkt\_Eq_{i,t-1}} + \gamma_8 ASST\_GWTH_{i,t} + \gamma_9 MLEV_{i,t} + \gamma_{10} \log(Mkt\_Eq_{i,t-1}) + \varepsilon_{i,t} \end{aligned} \tag{1}$$

Table 1. Summary statistics

Variables	Firm-year obs.	Mean	25%	Median	75%	Std. Dev.
Panel A: Summary statistics						
Pension variables						
ΔPA	23,446	0.016	0.000	0.006	0.023	0.117
ΔPBO	23,446	0.018	−0.001	0.007	0.023	0.081
ΔABO	17,028	0.016	−0.001	0.006	0.021	0.086
$\Delta Funding\_Status$	23,446	−0.003	−0.010	−0.001	0.007	0.106
ΔOFF_BAL	23,446	−0.004	−0.010	−0.001	0.005	0.102
$Mand1$	18,198	0.005	0.000	0.001	0.005	0.015
$Mand2$	18,198	0.004	0.000	0.001	0.004	0.012
Return variables						
$R_{i,t} - R_{B,t}$	23,446	0.014	−0.216	−0.031	0.166	0.460
Accounting variables						
$CASH_{t-1}$	23,446	0.110	0.019	0.054	0.129	0.196
$\Delta CASH$	23,446	0.009	−0.015	0.002	0.024	0.159
ΔINT	23,446	0.001	−0.002	0.000	0.003	0.040
$\Delta(EBIT + PC)$	23,446	0.017	−0.017	0.007	0.032	0.275
$\Delta Non\text{-}Cash\_Assets$	23,446	0.177	−0.010	0.058	0.193	0.792
$ACCRUALS$	23,446	−0.078	−0.106	−0.047	−0.011	0.214
$XFIN$	23,446	0.009	−0.043	−0.005	0.038	0.245
$ASST\_GWTH$	23,446	0.088	−0.013	0.048	0.123	0.264
$MLEV$	23,446	0.266	0.103	0.234	0.400	0.200
Mkt_Eq_{t-1}	23,446	5,687	200	898	3,359	19,497

Panel B: Pairwise correlations

	ΔPBO	ΔABO	$\Delta Funding\_Status$	ΔOFF_BAL
ΔPA	0.43**	0.70**	0.74**	0.71**
ΔPBO		0.81**	−0.21**	−0.20**
ΔABO			0.30**	0.29**
$\Delta Funding\_Status$				0.96**

	$Mand2$
$Mand1$	0.96**

	$\Delta CASH$	ΔINT	$\Delta(EBIT + PC)$	$\Delta Non\text{-}Cash\_Assets$	$ACCRUALS$	$XFIN$	$ASST\_GWTH$	$MLEV$	Mkt_Eq_{t-1}
$CASH_{-1}$	−0.01*	−0.21**	0.22**	−0.07**	−0.15**	−0.16**	−0.08**	0.06**	−0.21**
$\Delta CASH$		−0.41**	0.22**	−0.08**	−0.24**	0.05**	0.08**	−0.01	−0.03**
ΔINT			−0.26**	0.09**	0.36**	0.30**	0.14**	0.04**	0.03**
$\Delta(EBIT + PC)$				−0.03**	−0.15**	−0.15**	0.05**	0.02**	−0.06**
$\Delta Non\text{-}Cash\_Assets$					0.11**	0.40**	0.37**	0.07*	0.07**
$ACCRUALS$						0.33**	0.20**	−0.26**	0.13**
$XFIN$							0.44**	0.08**	0.02**
$ASST\_GWTH$								−0.02**	0.02**
$MLEV$									−0.15**

The sample consists of U.S. firms on the NYSE/AMEX/NASDAQ and the CRSP and COMPUSTAT merge files during the June 1988 to June 2017 period. Panel A reports summary statistics, constructed using pooled firm-year observations. Pension asset and liability related items include changes in the following variables: pension assets (ΔPA), projected benefit obligations (ΔPBO), accumulated benefit obligations (ΔABO), funding status ($\Delta Funding\_Status$), and off-balance-sheet pension assets or liabilities (ΔOFF_BAL). Mandatory contributions include two measures, *Mand*1 and *Mand*2, following Moody's (2006) and Campbell *et al*. (2012), respectively. The return variables include the raw returns (R_{it}) and the Fama–French 25 size and book-to-market ratio adjusted returns ($R_{it} - R_{Bt}$) measured over the corresponding fiscal year. The accounting variables include lagged cash holdings ($CASH_{-1}$), changes in cash holdings ($\Delta CASH$), changes in interest payments (ΔINT), changes in pension-cost adjusted earnings before interest and taxes ($\Delta EBIT + PC$), changes in non-cash assets ($\Delta Non\text{-}Cash\_Assets$), the level of accrual ($ACCRUALS$), the level of external financing ($XFIN$), asset growth ($ASST\_GWTH$), market leverage ($MLEV$), and beginning-of-fiscal-year market capitalization (Mkt_Eq_{-1}). Pension variables and accounting variables are all scaled by beginning of the fiscal year market capitalization except for $ASST\_GWTH$ and $MLEV$. The details of the construction of the variables are given in Appendix A. Panel B reports pair-wise correlation coefficients. * and ** indicate significance at the 10% and 5% levels, respectively.

The independent variables in equation (1) include the control variables discussed earlier.[6] The variable of interest is denoted by ΔY, which in each column of Table 2 equals the change in one of the following pension variables respectively: plan assets (ΔPA), projected benefit obligations (ΔPBO), accumulated

[6] We also tried specifications including changes in dividend payments, R&D expenditures, employee growth, and other interactive terms as in Faulkender and Wang (2006). These variables were not significant.

Table 2. Stock market valuation of corporate pension plans

Model	1	2	3	4	5	6	7
ΔPA	0.43						
	(2.27)**						
ΔPBO		−0.28					
		(−2.41)**					
ΔABO			−0.58				
			(−2.06)**				
$\Delta Funding\_Status$				0.37			
				(2.21)**			
ΔOFF_BAL					0.34		
					(1.83)*		
$Mand1$						−5.03	
						(−3.37)**	
$Mand2$							−6.37
							(−3.54)**
$CASH_{-1}$	0.18	0.24	0.31	0.21	0.21	0.21	0.23
	(2.72)**	(1.95)*	(3.17)**	(2.61)**	(2.57)**	(3.05)**	(3.56)**
$\Delta CASH$	0.13	0.08	0.50	0.14	0.13	0.47	0.47
	(0.85)	(0.46)	(4.90)**	(0.88)	(0.80)	(4.92)**	(5.10)**
ΔINT	0.48	0.42	−0.58	0.50	0.50	−0.47	−0.39
	(1.57)	(1.11)	(−0.98)	(1.62)	(1.57)	(−1.01)	(−0.76)
$\Delta(EBIT+PC)$	0.33	0.38	0.35	0.34	0.34	0.32	0.33
	(5.20)**	(8.14)**	(5.16)**	(5.28)**	(5.57)**	(4.27)**	(4.82)**
$\Delta Non\text{-}Cash\_Assets$	0.04	0.03	0.04	0.04	0.04	0.03	0.04
	(3.75)**	(3.32)**	(3.31)**	(3.84)**	(3.83)**	(2.08)**	(2.65)**
$ACCRUALS$	−0.24	−0.23	−0.06	−0.25	−0.25	−0.08	−0.10
	(−2.63)**	(−2.77)**	(−0.67)	(−2.63)**	(−2.63)**	(−0.88)	(−0.86)
$XFIN$	−0.19	−0.19	−0.20	−0.19	−0.19	−0.16	−0.17
	(−3.03)**	(−2.94)**	(−2.46)**	(−2.91)**	(−2.92)**	(−2.33)**	(−2.41)**
$ASST\_GWTH$	0.22	0.24	0.22	0.25	0.25	0.24	0.24
	(5.35)**	(5.04)**	(4.58)**	(5.43)**	(5.39)**	(5.11)**	(4.90)**
$MLEV$	−0.60	−0.60	−0.52	−0.58	−0.58	−0.54	−0.53
	(−11.21)**	(−11.50)**	(−12.57)**	(−10.72)**	(−10.78)**	(−12.05)**	(−11.59)**
$Log(Mkt\_Eq_{t-1})$	−0.02	−0.02	−0.01	−0.02	−0.02	−0.02	−0.02
	(−5.07)**	(−4.28)**	(−4.60)**	(−4.98)**	(−4.92)**	(−5.36)**	(−5.21)**
Intercept	0.17	0.16	0.14	0.17	0.17	0.15	0.15
	(4.31)**	(3.33)**	(3.25)**	(4.17)**	(4.17)**	(4.36)**	(4.09)**
Industry dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes
R^2	0.17	0.17	0.21	0.17	0.17	0.22	0.22
Obs.	23,446	23,446	17,028	23,446	23,446	18,198	18,198

Regression estimates of equation (1):

$$
\begin{aligned}
R_{i,t} - R_{B,i,t} = {} & \alpha_0 + \alpha_1 \Delta Y_{i,t} + \gamma_1 \frac{CASH_{i,t-1}}{Mkt\_Eq_{i,t-1}} + \gamma_2 \frac{\Delta CASH_{i,t}}{Mkt\_Eq_{i,t-1}} + \gamma_3 \frac{\Delta INT_{i,t}}{Mkt\_Eq_{i,t-1}} \\
& + \gamma_4 \frac{\Delta(EBIT+PC)_{i,t}}{Mkt\_Eq_{i,t-1}} + \gamma_5 \frac{Non{-}Cash\,Assets_{i,t}}{Mkt\_Eq_{i,t-1}} + \gamma_6 \frac{ACCRUALS_{i,t}}{Mkt\_Eq_{i,t-1}} \\
& + \gamma_7 \frac{XFIN_{i,t}}{Mkt\_Eq_{i,t-1}} + \gamma_8 ASST\_GWTH_{i,t} + \gamma_9 MLEV_{i,t} + \gamma_{10} log(Mkt\_Eq_{i,t}{-}1) + \varepsilon_{i,t}
\end{aligned}
$$

The sample consists of U.S. firms on the NYSE/AMEX/NASDAQ and the CRSP and COMPUSTAT merge files during the June 1988 to June 2017 period. The table reports the estimation results for the model in Grinblatt and Moskowitz (2004) and Faulkender and Wang (2006). The dependent variable is the Fama–French 25 size and book-to-market ratio adjusted return ($R_{it} - R_{Bt}$). The independent variables include lagged cash holdings ($CASH_{-1}$), changes in cash holdings ($\Delta CASH$), changes in interest payments (ΔINT), changes in pension-cost adjusted earnings before interest and taxes ($\Delta(EBIT+PC)$), changes in non-cash assets ($\Delta Non\text{-}Cash\_Assets$), the level of accrual ($ACCRUALS$), the level of external financing ($XFIN$), asset growth ($ASST\_GWTH$), market leverage ($MLEV$), and beginning-of-fiscal-year market value (Mkt_Eq_{-1}). Pension variables and accounting variables are all scaled by beginning-of-fiscal-year capitalization except for $ASST\_GWTH$ and $MLEV$. * and ** indicate significance at the 10% and 5% levels, respectively. Petersen (2009) and Thompson (2011) two-dimension firm and year clustered *t*-statistics are reported.

benefit obligations (ΔABO), funding status ($\Delta Funding\_Status$), off-balance sheet pension items (ΔOFF_BAL), and mandatory contributions. The control variables in Table 2 generally exhibit high levels of significance with the expected signs.

Because ΔPA is scaled by market capitalization, its coefficient in column (1) implies that each extra dollar of plan assets is valued by shareholders at \$0.43, with a *t*-statistic of 2.27. Similarly, each dollar in other pension items are valued on the margin as follows: projected benefit obligations (ΔPBO), −\$0.28, with a *t*-statistic of −2.41; accumulated benefit obligations (ΔABO), −\$0.58, with a *t*-statistic of −2.06; funding status ($\Delta Funding\_Status$), \$0.37, with a *t*-statistic of 2.21; off balance sheet pension items (ΔOFF_BAL), \$0.34, with a *t*-statistic of 1.83. We include industry fixed effects in the ordinary least squares (OLS) regressions; standard errors have been adjusted for clustering in both firm and year (Petersen, 2009; Thompson, 2011).

If the market views the firm as completely responsible for the additional assets and liabilities generated during the fiscal year, then the estimated coefficients should be close to \$1.00 for asset-related items and −\$1.00 for liability-related items. In fact, the regression coefficients are generally closer to one-half. Investors seem to be most sensitive to changes in accumulated benefit obligations (ΔABO), for which the −\$0.58 valuation coefficient is closest to \$1.00 in absolute value. The fact that these coefficients are all below 1.0 may reflect in part the fact that the net-of-tax value-impact of pension contributions and reversions would reflect the tax deductibility of such cash flows. However, even allowing for such effects, these coefficients are still considerably below the values one would expect if the firm had full and certain ownership claims and obligations for the pension plan.

The glaring exception is the response to mandatory contributions. The estimates (*t*-statistics) on *Mand*1 and *Mand*2 are −5.03 (−3.37) and −6.37 (−3.54), respectively. This high response superficially suggests market overreaction, but far more likely, it reflects a recognition that any change in mandatory contribution this year signifies a repeated obligation in following years for catch-up funding. Thus, the market appears to be imputing the present value of an *annuity* of mandatory contributions. In contrast, changes in the other pension fund measures, for example, accrued benefits, would not be expected to be persistent.

In Table 3, we check the robustness of these results by considering five alternative estimation methods. Method 1 employs OLS regressions with year and industry fixed effects. Method 2 employs OLS regressions with year and industry fixed effects, but with *t*-statistics adjusted for the clustering by firm. Method 3 includes only industry fixed effects. The *t*-statistics are adjusted for the clustering-in-year effects. Whereas the first three OLS regression methods do not employ firm dummies, method 4 uses a fixed-effect panel regression controlling for both year and firm effects with *t*-statistics adjusted for clustering by firm. Finally, method 5 employs a random-effect panel regression controlling for year, industry, and firm effects. The *t*-statistics are adjusted for clustering by firm. Because these OLS and panel regressions all generate similar results, in subsequent analysis, we rely on the simpler OLS approach with standard errors adjusted for clustering by both firm and year, as in Table 2.

5. Indices measuring changes in financial conditions

To test Hypothesis 2 (that pension funding effects depend on financial status), we require measures of financial strength. Therefore, we construct three measures to capture the evolution of firms' financial conditions and link that evolution to the valuation of corporate pension plans. Our measures are based on three popular models of bankruptcy risk: Ohlson (1980), Altman (1968), and Campbell *et al.* (2008). We use original estimates from these papers, except for the Altman (1968) model, for which we use the updated estimates from Shumway (2001). Appendix A summarizes the construction of these bankruptcy scores.[7]

[7]As a robustness check, we also classify change in financial strength using the two most important accounting ratios in almost all bankruptcy prediction models, net income and leverage. Our results using these variables are effectively identical to those we find using these three measures.

Table 3. Alternative regression methods

	Method 1	Method 2	Method 3	Method 4	Method 5
Pension variables					
ΔPA	0.52	0.52	0.43	0.41	0.49
	(4.22)**	(4.16)**	(2.29)**	(3.25)**	(4.27)**
ΔPBO	−0.30	−0.30	−0.28	−0.15	−0.29
	(−3.19)**	(−3.21)**	(−2.39)**	(−1.59)	(−3.09)**
ΔABO	−0.63	−0.63	−0.58	−0.44	−0.57
	(−3.69)**	(−3.69)**	(−2.06)**	(−3.46)**	(−4.08)**
ΔFunding_Status	0.47	0.47	0.37	0.40	0.43
	(2.72)**	(2.69)**	(2.24)**	(2.49)**	(2.61)**
ΔOFF_BAL	0.44	0.44	0.34	0.38	0.40
	(2.18)**	(2.17)**	(1.85)*	(2.16)**	(2.10)**
*Mand*1	−5.12	−5.12	−5.03	−4.62	−5.29
	(−5.03)**	(−4.89)**	(−3.41)**	(−5.76)**	(−5.59)**
*Mand*2	−6.64	−6.64	−6.37	−5.64	−6.76
	(−6.13)**	(−6.16)**	(−3.54)**	(−5.89)**	(−7.16)**
Control variables	Yes	Yes	Yes	Yes	Yes
Year dummy	Yes	Yes	No	Yes	Yes
Industry dummy	Yes	Yes	Yes	No	Yes
Firm effect (dummy)	No	No	No	Yes	Yes
t-statistic	Robust	Cluster on firm	Cluster on year	Cluster on firm	Cluster on firm
Method	OLS	OLS	OLS	Panel fixed effects	Panel random effects

Alternative regression estimates of equation (1). The sample consists of U.S. firms on the NYSE/AMEX/NASDAQ and the CRSP and COMPUSTAT merge files during the June 1988 to June 2017 period. The dependent variable is the Fama–French 25 size and book-to-market ratio adjusted return ($R_{it} - R_{Bt}$). The independent variables include lagged cash holdings ($CASH_{-1}$), changes in cash holdings (*ΔCASH*), changes in interest payments (*ΔINT*), changes in pension-cost adjusted earnings before interest and taxes (*Δ(EBIT + PC)*), changes in non-cash assets (*ΔNon-Cash_Assets*), the level of accrual (*ACCRUALS*), the level of external financing (*XFIN*), asset growth (*ASST_GWTH*), market leverage (*MLEV*), and beginning-of-fiscal-year market value (Mkt_Eq_{-1}). Pension variables, including changes in pension assets and liabilities (*ΔPA*, *ΔPBO*, *ΔABO*, *ΔFunding_Status*, and *ΔOFF_BAL*), are included one-at-a-time in each equation. Pension variables and accounting variables are all scaled by beginning of the fiscal year market capitalization except for *ASST_GWTH* and *MLEV*. * and ** indicate significance at the 10% and 5% levels, respectively.

5.1 Ohlson (1980) bankruptcy score

Ohlson (1980) estimates a static binary logit bankruptcy model. The probability of bankruptcy is modeled as $P(X_i,\beta) = [1 + \exp(-X_i\beta)]^{-1}$, where X_i is a vector of variables constructed from the firm's financial statements observed 1 year before bankruptcy. He finds four factors to be statistically significant for assessing the probability of bankruptcy: size, leverage, performance, and liquidity. The original estimated parameters from model 1 in Table 4 of Ohlson (1980) are used to compute Ohlson's bankruptcy score for firm i in fiscal year t: $Ohlson\_BS_{i,t} = -X_{i,t}\hat{\beta}$.[8]

5.2 Altman bankruptcy score

Altman (1968) applies discriminant analysis to a sample of bankrupt and non-bankrupt firms during the 1946–65 period. The following variables best distinguish between these two types of firms: working capital, retained earnings, earnings before interest and taxes, and sales, all scaled by total assets, and the ratio of market value of equity to book value of total debt. Shumway (2001) updates the estimates using data from 1962 to 1992. We use the estimates in Table 2 of Shumway (2001) to construct Altman's (1968) Z-score. We also add a negative sign to Z-scores, so that, like our other measures, higher algebraic values indicate riskier firms, $Altman\_Z_{i,t} = -X_{i,t}\hat{\beta}$.

5.3 Campbell, Hilscher, and Szilagyi bankruptcy score

Campbell *et al.* (2008) estimate bankruptcy risk using a multi-period dynamic logit model over the 1963–2003 period. Their model is similar to those of Shumway (2001) and Chava and

[8]Begley *et al.* (1996) and Hillegeist *et al.* (2004) re-estimate the Ohlson (1980) and Altman (1968) models. Shumway (2001) also re-estimates the Altman (1968) model. We repeat all empirical analyses using these alternative estimates and find similar results to our hypothesis testing.

Table 4. Bankruptcy scores and change in financial condition indices

	Obs.	Mean	Median	Std. Dev.	Mean	Median	Std. Dev.
Panel A: Summary statistics for bankruptcy scores and partitioning indices							
Ohlson (1980)			*Ohlson_BS*			*ΔOhlson_BS*	
All	23,446	−1.57	−1.45	1.75	−0.01	0.01	1.25
STRONGER	11,802	−1.97	−1.83	1.69	0.73	0.46	0.95
WEAKER	11,644	−1.15	−1.10	1.72	−0.76	−0.48	1.06
Altman (1968)			*Altman_Z*			*ΔAltman_Z*	
All	23,446	−1.09	−0.87	1.44	−0.02	0.01	0.81
STRONGER	12,012	−1.27	−1.00	1.51	0.34	0.19	0.73
WEAKER	11,434	−0.91	−0.74	1.34	−0.39	−0.21	0.72
Campbell *et al.* (2008)			*CHS_BS*			*ΔCHS_BS*	
All	23,446	−7.82	−7.97	0.74	0.01	0.01	0.70
STRONGER	11,823	−8.05	−8.14	0.56	0.47	0.31	0.51
WEAKER	11,623	−7.58	−7.78	0.82	−0.47	−0.31	0.51

Level of bankruptcy score			Change in bankruptcy score		
Panel B: Pairwise correlations for bankruptcy scores and partitioning indices					
	Altman_Z	*CHS_BS*		*ΔAltman_Z*	*ΔCHS_BS*
Ohlson_BS	0.48**	0.63**	*ΔOhlson_BS*	0.36**	0.47**
Altman_Z		0.41**	*ΔAltman_Z*		0.29**

		Second classification. Each entry equals percent agreement between two partitions		
First classification	Firm-year observations	*ΔOhlson_BS*	*ΔAltman_Z*	*ΔCHS_BS*
Panel C: Comparing classification results under alternative partitioning indices				
ΔOhlson_BS				
STRONGER firms	11,802		76	67
WEAKER firms	11,644		73	66
ΔAltman_Z				
STRONGER firms	12,012	74		64
WEAKER firms	11,434	75		64
ΔCHS_BS				
STRONGER firms	11,823	67	65	
WEAKER firms	11,623	66	63	

The sample consists of U.S. firms on the NYSE/AMEX/NASDAQ and the CRSP and COMPUSTAT merge files during the June 1988 to June 2017 period. Firms are partitioned into *STRONGER* and *WEAKER* groups based on whether annual changes in the following three bankruptcy scores are positive or negative: the Ohlson (1980) score (*ΔOhlson_BS*), the Altman (1968) index (*ΔAltman_Z*), and the Campbell *et al.* (2008) score (*ΔCHS_BS*). Panel A tabulates simple summary statistics for bankruptcy scores and partitioning indices for the entire sample and for the *STRONGER* and *WEAKER* subgroups of firms. Panel B presents pairwise correlations between bankruptcy scores and partitioning indices, using all firm-year observations. Panel C compares the classification results for *STRONGER* and *WEAKER* firms based on the partitioning indices. * and ** indicate significance at the 10% and 5% levels, respectively.

Jarrow (2004). The probability of bankruptcy in the next year is computed as $P(X_{i,t-1},\beta) = [1 + \exp(-X_{i,t-1}\beta)]^{-1}$. They identify several variables that are significant in predicting the probability of failure: relative size, lagged stock excess return, volatility, profitability, leverage, and current liquidity. The model is employed to predict 1, 6, 12, 24, and 36-month-ahead probabilities of failure. We use the original estimated coefficients for 1-year-ahead predictions in Table 4 of Campbell *et al.* (2008) to compute the bankruptcy score for firm i during fiscal year t, $CHS\_BS_{i,t} = -X_{i,t}\hat{\beta}$.

5.4 Indices measuring changes in financial conditions

Bankruptcy scores provide measures of each firm's financial condition. A simple *INDEX* of the change in financial condition over consecutive fiscal years is the difference in the bankruptcy score from t to $t-1$. For example, using the Ohlson bankruptcy score measure as the partitioning index:

$$INDEX_{i,t} = \Delta Ohlson\_BS_{i,t} = -(Ohlson\_BS_{i,t} - Ohlson\_BS_{i,t-1})$$

The higher the value of $\Delta Ohlson\_BS_{i,t}$, the more the firm's financial condition has improved. We construct analogous indexes for the two other bankruptcy scores, thus generating $INDEX_{i,t} = \Delta Altman\_Z_{i,t}$ and $\Delta CHS\_BS_{i,t}$ respectively.

We use these indexes to partition firms in each fiscal year into healthier and more distressed groups according to whether the *INDEX* is positive or negative and create two dummy variables (*STRONGER* and *WEAKER*)[9]:

$$STRONGER_{i,t} = \begin{cases} 1 & if \;\; INDEX_{i,t} > 0 \\ 0 & if \;\; INDEX_{i,t} \leq 0 \end{cases}$$

and

$$WEAKER_{i,t} = \begin{cases} 1 & if \;\; INDEX_{i,t} \leq 0 \\ 0 & if \;\; INDEX_{i,t} > 0 \end{cases}$$

Our partition generates a roughly equal number of firm-year observations for firms in both the healthier and more distressed groups for each of the three partitioning indices.

Figure 1 shows the mean values of each of these financial health indexes in each year. The figure shows that small funds exhibit noticeably greater variation in the evolution of their financial health: for each measure of financial strength, improving small firms enjoy greater average increases in financial health than improving large firms, and weakening small firms suffer greater average declines in financial health than weakening large firms.

5.5 Consistency of classification

The three partitioning indices, $\Delta Ohlson\_BS$, $\Delta Altman\_Z$, and ΔCHS_BS, each measuring changes in financial conditions, are constructed from historical estimates based on differing statistical models and sample periods. This raises the question of whether the indices generally identify the same set of firm-years as exhibiting improved or worsening financial conditions.

Panel A of Table 4 reports the mean, median, and standard deviation of the three bankruptcy-risk scores (*Ohlson_BS*, *Altman_Z*, and *CHS_BS*) as well as their changes. We report summary statistics for all firms, healthier firms, and more distressed firms, respectively. Panel B reports correlations between bankruptcy scores and between partitioning indices. The correlations between bankruptcy scores range from 0.41 to 0.63. Correlations between the partitioning indices that measure changes in financial conditions range from 0.29 to 0.47. All correlations are significant at better than the 1% level.

Panel C examines the consistency of classification under these alternative partitioning indices. There are a total of 23,446 firm-year observations in our entire sample. Of these, 11,802 and 11,644 are classified as *STRONGER* and *WEAKER*, respectively, under the first partitioning index, $\Delta Ohlson\_BS$. The question is how many of these 11,802 firm-year observations will be similarly classified as *STRONGER* under the two other partitioning indices. Panel C shows that of the 11,802 *STRONGER* firm-year observations, 76% and 67% are also classified as *STRONGER* under $\Delta Altman\_Z$ and ΔCHS_BS, respectively. Similarly, of the 11,644 firm-year observations classified as *WEAKER* under $\Delta Ohlson\_BS$, 73% and 66% are also classified as *WEAKER* under $\Delta Altman\_Z$ and ΔCHS_BS, respectively.

The other parts of panel C summarize the classification results when we alternate the first partitioning index and the second partitioning index. Between 64% and 76% of firms identified as *STRONGER* under the first index remain so under the second index. Between 63% and 75% of firms identified as *WEAKER* under the first index remain so under the second. We conclude that these measures are largely in agreement.

[9]Our results are essentially the same if we partition firms every year based on the median value of the change in the bankruptcy score rather than on a positive versus negative criterion.

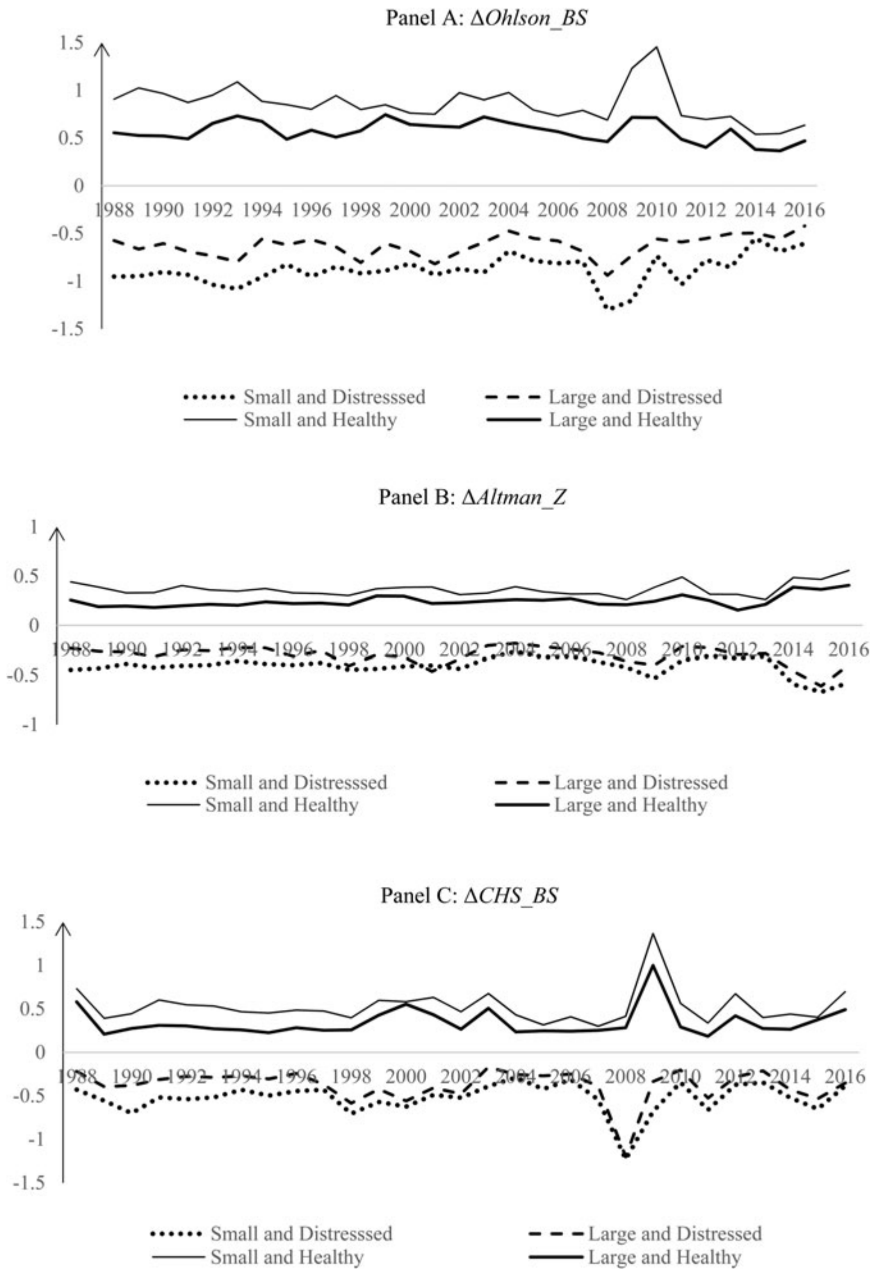

Figure 1. The evolution of indices measuring changes in financial conditions over time. Panel A: Δ*Ohlson_BS*; panel B: Δ*Altman_Z*; panel C: Δ*CHS_BS*.

6. Persistence of firms' financial conditions

The stability of financial ratios may change over time, especially when firms approach bankruptcy (Dambolena and Khoury, 1980). Therefore, in this section, we examine two issues related to the reliability of the bankruptcy scores. First, we compare the information content of their current levels as well as their change over consecutive years. Second, we study the persistence of changes in condition of firms in the *STRONGER* and *WEAKER* groups.

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6.1 The information content of bankruptcy scores and changes in bankruptcy score

We examine the information content of the level and change in the three alternative bankruptcy scores following the approach in Hillegeist *et al.* (2004). However, rather than applying these models to predict actual failure during the sample period, we apply them to predict which firms are at risk of entering financial distress. A firm is defined to be at risk of distress in year $t+1$ if net income as a percent of beginning-of-fiscal year market value, NI_t/Mkt_Eq_{t-1}, is less than −20%. We use 'at risk of distress' rather than actual failure because financial distress is sufficient to encourage risk-shifting behavior. This criterion for distress is still demanding: the fraction of firms in our sample that exhibit financial distress according this criterion is small, only 3.45% of the firm-year observations. Our results are robust to alternative cut-off levels of NI_t/Mkt_Eq_{t-1} such as −30%.

We test the 1-year-ahead predictive power for financial distress constructed from alternative bankruptcy scores (*BS*) using the following logit regression:

$$P(At\_Risk_{i,t+1} = 1) = \frac{\exp(a_0 + a_1 BS_{i,t})}{1 + \exp(a_0 + a_1 BS_{i,t})}$$

where $At\_Risk_{i,t+1}$ is a dummy variable that takes the value 1 if firm i in fiscal year $t+1$ is at risk of distress and 0 otherwise. Panel A of Table 5 reports logit regressions when current financial conditions measured by the three bankruptcy scores (*Ohlson_BS*, *Altman_Z*, and *CHS_BS*) are used as explanatory variables. Panel B reports estimates when both lagged levels and changes in financial conditions are used as explanatory variables.

By and large, these credit risk scores do predict financial distress over short horizons, with highly significant coefficient estimates. Table 5 shows that the *CHS_BS* score from the Campbell *et al.* (2008) model has the highest predictive power for financial distress measured by both the R^2 and log-likelihood function, followed by the *Altman_Z* score and the *Ohlson_BS* score. The models that use both lagged level and change in scores slightly outperform those that use only the current level of bankruptcy.

6.2 Persistence of changes in firms' financial conditions

We now examine the serial correlation of changes in bankruptcy scores following the method in Ali and Zarowin (1992). In each year, we estimate the following cross-sectional regression for both groups of firms (*STRONGER* and *WEAKER*), respectively:

$$INDEX_{i,t} = c_0 + c_1 INDEX_{i,t-1} + \varepsilon_{i,t} \tag{2}$$

where $INDEX_{i,t}$ is one of the three measures capturing the change in firms' financial conditions.[10] Table 6 partitions firms each year into *STRONGER* and *WEAKER* groups based on their change in financial condition (Δ*Ohlson_BS*, Δ*Altman_Z*, Δ*CHS_BS*). The table reports the annual average of the intercept and slope coefficients of equation (2) as well as the corresponding Fama and MacBeth (1973) t-statistics for the two partitioned groups. The table also reports the Wilcoxon (1945) signed rank test statistic and the corresponding p-values for the null hypothesis that the intercept and slope coefficients for the *STRONGER* and *WEAKER* groups are equal.

Table 6 reveals two broad patterns. On the one hand, average intercepts are positive for *STRONGER* firms and negative for *WEAKER* ones. Therefore, improving (*STRONGER*) firms in 1 year tend to strengthen again in the following year, while *WEAKER* firms continue to weaken. On the other hand, average slope coefficients for both groups are negative. The more negative the slope coefficient,

[10]We also estimate equation (2) treating the $INDEX_{i,t}$ as the change in net income ($\Delta NI/Mkt\_Eq_{t\text{-}1}$) or market leverage ($\Delta MLEV$), respectively. These variables resulted in essentially identical results to those obtained from the three bankruptcy scores.

Table 5. Predictive power of levels and changes in financial conditions

Panel A: Predictive power of current level of financial conditions			
Constant	−2.84	−2.78	6.77
	(−26.59)**	(−22.59)**	(10.59)**
$Ohlson\_BS_{i,t}$	−0.52		
	(−11.33)**		
$Altman\_Z_{i,t}$		−0.95	
		(−10.76)**	
$CHS\_BS_{i,t}$			−1.37
			(−15.80)**
Log likelihood	− 3,118.6	− 2,978.9	− 2,684.4
Pseudo-R^2	0.079	0.120	0.207
Observations	22,796	22,796	22,796
Panel B: Predictive power of lagged level and change in financial conditions			
Constant	−2.90	−2.81	8.14
	(−28.52)**	(−23.58)**	(15.65)**
$Ohlson\_BS_{i,t-1}$	−0.49		
	(−11.26)**		
$(Ohlson\_BS_{i,t} - Ohlson\_BS_{i,t-1})$	−0.59		
	(−10.79)**		
$Altman\_Z_{i,t-1}$		−0.92	
		(−9.82)**	
$(Altman\_Z_{i,t} - Altman\_Z_{i,t-1})$		−1.05	
		(−9.87)**	
$CHS\_BS_{i,t-1}$			−1.55
			(−21.32)**
$(CHS\_BS_{i,t} - CHS\_BS_{i,t-1})$			−1.25
			(−13.76)**
Log likelihood	−3,110.1	−2,976.6	−2,660.8
Pseudo-R^2	0.081	0.121	0.214
Observations	22,796	22,796	22,796

The sample consists of U.S. firms on the NYSE/AMEX/NASDAQ and the CRSP and COMPUSTAT merge files during the June 1988 to June 2017 period. The table reports the 1-year ahead predictive power of bankruptcy scores (*BS*) for future firm 'distress' in the following logit regression:

$$P(At\_Risk_{i,t+1} = 1) = \frac{e^{\alpha_0 + \alpha_1 BS_{i,t}}}{1 + e^{\alpha_0 + \alpha_1 BS_{i,t}}},$$

where $At\_Risk_{,t+1}$ is a dummy variable that takes the value one if firm *i* in fiscal year $t+1$ is deemed to be in distress and zero otherwise. A firm is defined to be in distress if its net income scaled by beginning-of-fiscal-year market equity is less than −20%. A total of 3.37% of firm-years in our sample exhibit distress. Panel A reports the logit regressions when current levels of financial conditions constructed from three bankruptcy scores (*Ohlson_BS*, *Altman_Z*, and *CHS_BS*) are used as explanatory variables. Panel B reports the logit regressions when both lagged levels and changes in financial conditions are used as explanatory variables. * and ** indicate significance at the 10% and 5% levels, respectively.

the greater the mean reversion in financial condition. Nevertheless, the values of c_1 and the changes in bankruptcy scores are sufficiently small that these mean-reversion terms do not come close to offsetting the impact of the intercepts.

For example, panel A of Table 4 shows that the mean values of $\Delta Ohlson\_BS$ are 0.73 and −0.76, respectively for the *STRONGER* and *WEAKER* groups. The mean values of c_0 are 0.64 and −0.70, respectively, and the mean values of c_1 are −0.24 and −0.16, respectively (Table 6). Therefore, the predicted change in the $\Delta Ohlson\_BS$ for each group of firms is:

$$STRONGER: \quad 0.64 \; - 0.24 \times 0.73 = 0.46$$
$$WEAKER: \quad -0.70 \; - 0.16 \times (-0.76) = -0.58$$

Therefore, changes in financial condition continue into the following year, with improving firms continuing to improve and deteriorating firms continuing to worsen. The product of the average value of c_1 and the mean value of the *INDEX* is generally much less than the average value of the intercept; the product is typically less than 30% of the intercept.

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Table 6. Persistence of changes in financial condition indices

	STRONGER	*WEAKER*	*Z*-statistic (p-value) for equality of coefficients
INDEX = Δ*Ohlson_BS*			
Mean of c_0	0.64	−0.70	6.54 (0.00)**
t-statistic	(27.04)**	(−25.69)**	
Mean of c_1	−0.24	−0.16	−2.81 (0.01)**
t-statistic	(−9.99)**	(−6.13)**	
INDEX = Δ*Altman_Z*			
Mean of c_0	0.33	−0.37	6.54 (0.00)**
t-statistic	(17.75)**	(−17.65)**	
Mean of c_1	−0.20	−0.03	−3.24 (0.00)**
t-statistic	(−4.44)**	(−0.69)	
INDEX = Δ*CHS_BS*			
Mean of c_0	0.35	−0.43	6.54 (0.00)**
t-statistic	(21.87)**	(−12.20)**	
Mean of c_1	−0.28	−0.07	−5.16 (0.00)**
t-statistic	(−9.42)**	(−3.67)**	

Regression estimates of equation (2):

$$INDEX_{i,t} = c_0 + c_1 INDEX_{i,t}-1 + \varepsilon_{i,t}$$

where $INDEX_{i,t}$ is one of the three measures capturing the change in firms' financial condition. The sample consists of U.S. firms on the NYSE/AMEX/NASDAQ and the CRSP and COMPUSTAT merge files during the June 1988 to June 2017 period. Every year, the table partitions firms into *STRONGER* and *WEAKER* groups based on the change in financial conditions (Δ*Ohlson_BS*, Δ*Altman_Z*, Δ*CHS_BS*). The table reports the annual average of the regression coefficients and the corresponding Fama and MacBeth (1973) *t*-statistics for the two partitioned groups. The table also reports the Wilcoxon (1945) signed rank test statistics and the corresponding p-values for the null hypothesis that the slope coefficients from the *STRONGER* and *WEAKER* groups are the same. * and ** indicate significance at the 10% and 5% levels, respectively.

Given this pattern, recent *changes* in financial condition appear to be informative for assessing longer-horizon likelihood of financial distress. The strength of this persistence alleviates concerns that random fluctuations in probability of distress would lead to misclassification of improving or deteriorating firms.[11] It seems plausible that the market would extrapolate recent changes into the future as it forecasts a firm's financial condition.

7. Conditional valuation of pension assets and liabilities

We now proceed to examine Hypothesis 2 regarding the relation between financial strength and pension valuation effects. Equation (3) allows the response of excess return to changes in various components of pension funding to depend on financial condition:

$$\begin{aligned} R_{i,t} - R_{B,i,t} = &\alpha_0 + (\alpha_1 \times STRONGER + \alpha_2 \times WEAKER)\Delta Y_{i,t} \\ &+ \alpha_3 \times INDEX_{i,t} + \boldsymbol{\gamma} \cdot \boldsymbol{Z}_{i,t} + \varepsilon_{i,t} \end{aligned} \tag{3}$$

where $INDEX_{i,t}$ denotes $\Delta Ohlson\_BS_{i,t}$, $\Delta Altman\_Z_{i,t}$, and $\Delta CHS\_BS_{i,t}$, respectively and $\Delta Y_{i,t}$ is the change in one of the pension variables. The response of STRONGER firms to each item is α_1, while that of WEAKER firms is α_2. We hypothesize that α_1 is greater than α_2, indicating that market values of STRONGER firms are more sensitive to changes in pension variables. The vector $\boldsymbol{Z}_{i,t}$ contains the same control variables as those specified in equation (1); $\boldsymbol{\gamma} = [\gamma_1,\ldots,\gamma_{10}]$ represents the vector of regression coefficients associated with the control variables.

Table 7 presents OLS estimates for equation (3). We first discuss the Δ*Ohlson_BS* index. Looking across the top row of panel A (which report the estimates of α_1), we see that market valuations of *STRONGER* firms respond between \$0.48 and \$0.95 to each dollar change in various measures of pension funding; all coefficients are significant at better than the 1% level.

In sharp contrast, the second row shows that the estimated coefficients, α_2, on *WEAKER* × Δ*Y* are much smaller and generally not statistically significant. These results thus indicate strong differential

[11]To the extent that firms are misclassified, that would actually weaken our results rather than lead to false positives.

effects between firms experiencing improvement in financial health and those encountering financial difficulty. The *F*-statistics in panel A confirm that one can easily reject the null hypothesis that regression coefficients from the *STRONGER* and *WEAKER* firms are equal, i.e., that $\alpha_1 = \alpha_2$ for all five pension variables.

We also include the health *INDEX* directly in the regressions (see the estimates of α_3 in the third row). For the first set of regressions, the *INDEX* is taken to be $\Delta Ohlson\_BS$. The estimated coefficient on $\Delta Ohlson\_BS$, α_3, is positive and highly significant in all cases, indicating that stock returns respond to improvements in financial condition. The R^2s of the model fit range from 0.18 to 0.23.

Panels B and C of Table 7 summarize the results when we use $\Delta Altman\_Z$ and ΔCHS_BS as the partitioning indices. Overall, the results are remarkably stable across each measure of financial health. The null hypothesis that regression coefficients from the *STRONGER* and *WEAKER* subsamples are the same, i.e.,$\alpha_1 = \alpha_2$, is firmly rejected for all five pension variables under both the $\Delta Altman\_Z$ and ΔCHS_BS indices. The R^2s range from 0.21 to 0.26 under the $\Delta Altman\_Z$ index and from 0.28 to 0.30 under the ΔCHS_BS index.

The valuation impact of an extra dollar of *ABO* is approximately $1.00 under all three indices, whereas the coefficients on *PBO* are generally only about half of that value. This suggests that shareholders are far more sensitive to changes in accumulated benefit obligations (ΔABO) than to changes in projected benefit obligations (ΔPBO). This is consistent with the fact that *ABO* measures the firm's legal liability.

As a robustness check, we present Table 8, in which we distinguish firms using dummy variables for the *level* rather than the *change* in financial condition. *WEAK_FIRM*s are the riskiest, with the highest 10% of credit-risk scores, while *STRONG_FIRM*s are the remaining 90% of firm-year observations. As noted above, total firm-year observations is 23,446, corresponding to 2,117 unique firms. The 10% highest credit-risk firms therefore produce 2,344 firm-year observations. The numbers of unique firms are 842, 754, and 999 when measuring credit-risk by *OBS*, *ABS*, and *CBS* respectively. This number of unique firms is quite large, so our results are not driven by a small number of extremely weak firms.

We estimate the following equation, and report in Table 8 the estimates of α_1 and α_2 as well as *F*-statistics for the difference between them:

$$\begin{aligned} R_{i,t} - R_{B,i,t} = \alpha_0 &+ (\alpha_1 \times STRONG\_FIRM + \alpha_2 \times WEAK\_FIRM)\Delta Y_{i,t} \\ &+ \alpha_3 \times INDEX_{i,t} + \boldsymbol{\gamma} \cdot \boldsymbol{Z_{i,t}} + \varepsilon_{i,t} \end{aligned} \tag{4}$$

The results for levels are generally consistent with those for which the sample was split by change in credit risk. The coefficients on strong firms all have the expected sign and are (with only one exception) statistically significant. The coefficients on weak firms are less consistent, which is not surprising, since these firms would have the weakest claim to net pension assets. In every case that the difference in coefficients is statistically significant, the coefficient on strong firms has higher absolute value than the one on weak firms. Other splits (e.g., 20% weakest/80% other firms) yielded similar results. On balance, however, the stronger results obtained in Table 7 using changes in financial condition support the hypothesis that the market recognizes the high persistence of these changes and extrapolates them into the future.

We also examine interactions between financial condition and pension funding. We define *UNDERFUNDED* plans as those for which the pension plan asset shortfall relative to *PBO* exceeds 10% of the market value of equity. The remainder of funds is deemed *ADEQUATELY_FUNDED*. The following regression specification allows the value impact of changes in pension funding to vary according to a two-way classification by both funding and financial strength. Results are presented in Table 9.

$$\begin{aligned} R_{i,t} - R_{B,i,t} = \alpha_0 &+ (\alpha_1 \times UNDERFUNDED + \alpha_2 \times ADEQUATELY\_FUNDED) \\ &\times STRONG\_FIRM + (\alpha_3 \times UNDERFUNDED + \alpha_4 \times ADEQUATELY\_FUNDED) \\ &\times WEAK\_FIRM + \alpha_5 \times INDEX_{i,t} + \boldsymbol{\gamma} \cdot \boldsymbol{Z_{i,t}} + \varepsilon_{i,t}. \end{aligned} \tag{5}$$

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Table 7. Stock market valuation conditional on change in financial condition

		ΔPA	ΔPBO	ΔABO	$\Delta Funding\_Status$	ΔOFF_BAL
Panel A: Firms partitioned by change in the Ohlson (1980) score						
	α_1 : $STRONGER \times \Delta Y$	0.65	−0.48	−0.95	0.52	0.51
		(4.54)**	(−2.66)**	(−2.73)**	(4.97)**	(4.30)**
	α_2 : $WEAKER \times \Delta Y$	−0.07	−0.06	−0.07	−0.25	−0.30
		(−0.93)	(−1.27)	(−1.55)	(−2.11)**	(−2.65)**
	α_3 : $\Delta Ohlson\_BS$	0.05	0.05	0.04	0.05	0.05
		(5.25)**	(5.03)**	(3.70)**	(5.30)**	(5.30)**
F-statistic for $\alpha_1 = \alpha_2$		33.72**	5.99**	6.42**	87.38**	74.95**
(p-value)		(0.00)	(0.01)	(0.01)	(0.00)	(0.00)
Control variables		Yes	Yes	Yes	Yes	Yes
Industry dummies		Yes	Yes	Yes	Yes	Yes
R^2		0.20	0.18	0.23	0.19	0.19
Obs.		23,446	23,446	17,028	23,446	23,446
		ΔPA	ΔPBO	ΔABO	$\Delta Funded\_Status$	ΔOFF_BAL
Panel B: Firms partitioned by change in the Altman (1968) Z-score						
	α_1: $STRONGER \times \Delta Y$	0.66	−0.55	−0.98	0.50	0.48
		(4.59)**	(−2.73)**	(−2.90)**	(4.30)**	(3.63)**
	α_2 : $WEAKER \times \Delta Y$	−0.12	0.03	0.05	−0.11	−0.13
		(−1.45)	(0.47)	(0.56)	(−0.82)	(−0.96)
	α_3 : $\Delta Altman\_Z$	0.12	0.12	0.10	0.13	0.13
		(4.15)**	(4.12)**	(3.25)**	(4.15)**	(4.14)**
F-statistic for $\alpha_1 = \alpha_2$		44.14**	7.44**	11.45**	200.57**	157.07**
(p-value)		(0.00)	(0.01)	(0.00)	(0.00)	(0.00)
Control variables		Yes	Yes	Yes	Yes	Yes
Industry dummies		Yes	Yes	Yes	Yes	Yes
R^2		0.23	0.21	0.26	0.22	0.22
Obs.		23,446	23,446	17,028	23,446	23,446
		ΔPA	ΔPBO	ΔABO	$\Delta Funding\_Status$	ΔOFF_BAL
Panel C: Firms partitioned by change in the Campbell *et al.* (2008) score						
	α_1: $STRONGER \times \Delta Y$	0.61	−0.34	−0.87	0.51	0.49
		(3.89)**	(−2.21)**	(−2.81)**	(4.19)**	(3.65)**
	α_2 : $WEAKER \times \Delta Y$	−0.30	0.01	0.07	−0.58	−0.64
		(−2.06)**	(0.15)	(0.92)	(−2.92)**	(−3.16)**
	α_3 : ΔCHS_BS	0.23	0.23	0.20	0.24	0.24
		(7.651)**	(7.18)**	(5.75)**	(8.24)**	(8.24)**
F-statistic for $\alpha_1 = \alpha_2$		19.12**	4.38**	8.96**	24.04**	23.34**
(p-value)		(0.00)	(0.04)	(0.00)	(0.00)	(0.00)
Control variables		Yes	Yes	Yes	Yes	Yes
Industry dummies		Yes	Yes	Yes	Yes	Yes
R^2		0.29	0.27	0.30	0.28	0.28
Obs.		23,446	23,446	17,028	23,446	23,446

Regression estimates of equation (3):

$$R_{i,t} - R_{B,i,t} = \alpha_0 + (\alpha_1 \times STRONGER + \alpha_2 \times WEAKER)\Delta Y_{i,t} + \alpha_3 \times INDEX_{i,t} + \boldsymbol{\gamma} \cdot \boldsymbol{Z}_{i,t} + \varepsilon_{i,t}$$

The sample consists of U.S. firms on the NYSE/AMEX/NASDAQ and the CRSP and COMPUSTAT merge files during the June 1988 to June 2017 period. Panel A reports the results of regressing the Fama–French 25 size and book-to-market ratio adjusted excess returns ($R_{it} - R_{Bt}$) on pension funding variables. Pension variables, including changes in pension assets and liabilities (ΔPA, ΔPBO, ΔABO, $\Delta Funding\_Status$, and ΔOFF_BAL), are included one-at-a-time in each equation. In panel A, dummy variables are constructed based on the $\Delta Ohlson\_BS$ index to indicate firms that experience improvement in financial conditions (*STRONGER*) and the firms that experience deterioration in financial conditions (*WEAKER*). The table reports estimated coefficients on the interactive terms $STRONGER \times \Delta Y$ and $WEAKER \times \Delta Y$, where ΔY denotes one of the pension variables. Panel B repeats the analysis using the $\Delta Altman\_Z$ index. Panel C repeats the analysis using the ΔCHS_BS index. The reported F-statistic is for the test of the null hypothesis that the regression coefficients for *STRONGER* and *WEAKER* firms are the same, i.e., $\alpha_1 = \alpha_2$. Other control variables and industry dummies are included in the regressions. * and ** indicate significance at the 10% and 5% levels, respectively. Petersen (2009) and Thompson (2011) two-dimension firm and year clustered t-statistics are reported.

Table 8. Stock market valuation conditional on level of financial condition

	Ohlson_BS			*Altman_Z*			*CHS_BS*		
	Strong firms	Weak firms	*F*-statistic (p-value)	Strong firms	Weak firms	*F*-statistic (p-value)	Strong firms	Weak firms	*F*-statistic (p-value)
Δ*PA*	0.51	0.04	4.07**	0.48	0.21	12.93**	0.48	−0.21	6.39**
	(2.56)**	(0.32)	(0.04)	(2.67)**	(1.14)	(0.00)	(2.36)**	(−1.35)	(0.01)
Δ*PBO*	−0.25	−0.35	0.63	−0.28	−0.20	0.19	−0.18	−0.17	0.01
	(−2.10)**	(−2.46)**	(0.43)	(−2.77)**	(−0.89)	(0.66)	(−1.55)	(−1.40)	(0.92)
Δ*ABO*	−0.68	−0.34	2.80*	−0.62	−0.46	2.57	−0.65	0.14	3.36*
	(−2.13)**	(−2.40)**	(0.09)	(−2.08)**	(−1.60)	(0.11)	(−2.02)**	(0.86)	(0.07)
Δ*Funding_Status*	0.50	−0.280	10.13**	0.45	0.05	1.22	0.46	−0.45	12.07**
	(3.24)**	(−2.02)**	(0.00)	(2.33)**	(0.26)	(0.27)	(2.98)**	(−2.37)**	(0.00)
Δ*OFF_BAL*	0.46	−0.30	8.10**	0.44	−0.03	1.05	0.43	−0.39	9.72**
	(2.64)**	(−2.19)**	(0.00)	(1.97)**	(−0.10)	(0.31)	(2.52)**	(−1.95)*	(0.00)
*Mand*1	−6.40	−1.43	10.70**	−6.32	−0.96	14.55**	−6.31	0.58	13.43**
	(−5.73)**	(−0.84)	(0.00)	(−7.59)**	(−0.55)	(0.00)	(−5.51)**	(0.49)	(0.00)
*Mand*2	−7.53	−1.49	8.63**	−7.65	−1.21	19.22**	−7.32	0.35	7.57**
	(−4.92)**	(−0.75)	(0.00)	(−5.67)**	(−0.66)	(0.00)	(−4.01)**	(0.20)	(0.01)

Regression estimates of equation (4):

$$R_{i,t} - R_{B,i,t} = \alpha_0 + (\alpha_1 \times STRONG\_FIRM + \alpha_2 \times WEAK\_FIRM)\Delta Y_{i,t} + \alpha_3 \times INDEX_{i,t} + \boldsymbol{\gamma} \cdot \mathbf{Z}_{i,t} + \varepsilon_{i,t}$$

The sample consists of U.S. firms on the NYSE/AMEX/NASDAQ and the CRSP and COMPUSTAT merge files during the June 1988 to June 2017 period. Panel A reports the results of regressing the Fama–French 25 size and book-to-market ratio adjusted excess returns ($R_{it} - R_{Bt}$) on pension-funding variables. Pension variables, including changes in pension assets and liabilities (Δ*PA*, Δ*PBO*, Δ*ABO*, Δ*Funding_Status*, and Δ*OFF_BAL*), are included one-at-a-time in each equation. Dummy variables are constructed based on the *Ohlson_BS*, *Altman_Z*, and *CHS_BS*, respectively, to indicate firms that have low or high probability of distress. The *WEAK_FIRM* dummy equals 1 if the firm is the 10% of firms with the highest probability of financial distress. The *STRONG_FIRM* dummy equals 1 for the remaining 90% of firms. The reported *F*-statistic is for the test of the null hypothesis that the regression coefficients for *STRONG_FIRM*s and *WEAK_FIRM*s firms are the same, i.e., $\alpha_1 = \alpha_2$. Other control variables and industry dummies are included in the regressions. * and ** indicate significance at the 10% and 5% levels, respectively. Petersen (2009) and Thompson (2011) two dimension firm and year clustered *t*-statistics are reported.

Results of Table 9 add nuance to those in Table 8, but do not reverse any of its key conclusions. For adequately funded weak firms, increases in net pension funding actually tend to be value reducing: the estimates of α_4 are negative and significant for increases in assets or funding status. For example, using the Ohlson (1980) score, the estimated coefficient (*t*-statistic) on Δ*PA* is $\alpha_4 = -0.27$ (−2.56). This may signify that firms already at high risk of financial distress reduce value by transferring funds into pension accounts, at least when those plans are already adequately funded and the contributions are not immediately necessary. α_3 has the opposite sign as α_4. The corresponding estimate of α_3 is 0.33 (1.96), implying that contributions by weak firms to underfunded plans are value increasing. This may reflect an information effect, with those contributions signaling confidence that the firm will overcome its current financial weakness, making the shoring up of underfunded pension plans a reasonable strategy.

For strong firms, funding underfunded pension plans is consistently value enhancing, with increases in assets receiving consistently positive coefficients and increases in pension liabilities receiving negative coefficients (see the estimates for α_1). Again, we interpret this as an information effect, in which the decision to shore up an underfunded plan is a signal of confidence in the firm's prospects as a going concern. In contrast, when pension plans of strong firms are already adequately funded, further funding does not increase firm value: the estimates of α_2 are generally insignificant, and when significant, are only about one-third the size of α_1 (see the columns for mandatory contributions).

7.1 Value relevance and accounting standards

Pension accounting standards have undergone major changes over the years, which allows us to test for the effect of such rules on value relevance. Dhaliwal (1986), Davis-Friday *et al.* (1999), Amir *et al.* (2010), Chuk (2013), and Yu (2013) examine the impact of two major changes in pension accounting

Table 9. Interaction effects between financial strength and pension funding

	ΔPA	ΔPBO	ΔABO	ΔFunding_Status	ΔOFF_BAL	Mand1	Mand2
Panel A: Firms partitioned using *WEAK_FIRM/STRONG_FIRM* dummies defined using Ohlson (1980) score							
α_1	0.80	−0.48	−1.01	0.62	0.60	−6.82	−8.60
	(7.51)**	(−2.63)**	(−2.93)**	(7.21)**	(6.07)**	(−6.91)**	(−6.91)**
α_2	−0.04	−0.03	−0.04	−0.24	−0.29	−2.94	−2.58
	(−0.81)	(−0.48)	(−0.43)	(−1.73)*	(−2.12)**	(−2.78)**	(−2.40)**
α_3	0.33	−0.53	−0.53	−0.27	−0.20	−1.55	−1.97
	(1.96)**	(−2.78)**	(−3.92)**	(−1.27)	(−1.04)	(−0.84)	(−0.82)
α_4	−0.27	0.07	0.17	−0.46	−0.57	0.69	1.14
	(−2.56)**	(0.65)	(1.25)	(−3.97)**	(−4.32)**	(0.51)	(0.85)
F-stat for $\alpha_1 = \alpha_2$	52.99**	5.80**	7.11**	40.01**	47.62**	8.69**	14.19**
p-value	(0.00)	(0.02)	(0.01)	(0.00)	(0.00)	(0.00)	(0.00)
F-stat for $\alpha_3 = \alpha_4$	8.48**	8.25**	10.55**	0.61	2.57	1.44	1.99
p-value	(0.00)	(0.00)	(0.00)	(0.43)	(0.11)	(0.23)	(0.16)
R^2	0.20	0.18	0.23	0.19	0.19	0.24	0.24
Obs.	23,446	23,446	17,028	23,446	23,446	18,198	18,198
Panel B: Firms partitioned using *WEAK_FIRM/STRONG_FIRM* dummies defined using Altman (1968) *Z*-score							
α_1	0.75	−0.60	−0.94	0.57	0.58	−6.65	−8.65
	(9.22)**	(−3.45)**	(−2.63)**	(4.22)**	(3.64)**	(−9.35)**	(−8.34)**
α_2	−0.07	0.01	−0.03	−0.23	−0.29	−3.27	−2.82
	(−1.11)	(0.17)	(−0.27)	(−1.60)	(−1.96)**	(−3.06)**	(−2.64)**
α_3	0.74	−0.29	−0.77	0.28	0.13	−1.25	−1.95
	(1.84)*	(−0.96)	(−2.70)**	(0.85)	(0.35)	(−0.62)	(−0.81)
α_4	−0.22	−0.01	0.26	−0.43	−0.45	1.76	1.80
	(−2.60)**	(−0.09)	(1.70)*	(−4.23)**	(−3.33)**	(0.88)	(0.90)
F-stat for $\alpha_1 = \alpha_2$	73.98**	12.23**	6.27**	23.53**	25.23**	9.31**	19.77**
p-value	(0.00)	(0.00)	(0.01)	(0.00)	(0.00)	(0.00)	(0.00)
F-stat for $\alpha_3 = \alpha_4$	4.97**	0.86	9.32**	4.09**	2.07	1.07	1.30
p-value	(0.03)	(0.35)	(0.00)	(0.04)	(0.15)	(0.30)	(0.25)
R^2	0.23	0.21	0.26	0.22	0.22	0.27	0.26
Obs.	23,446	23,446	17,028	23,446	23,446	18,198	18,198
Panel C: Firms partitioned using *WEAK_FIRM/STRONG_FIRM* dummies defined using Campbell *et al.* (2008) score							
α_1	0.74	−0.39	−0.96	0.58	0.57	−6.75	−8.45
	(6.47)**	(−2.34)**	(−2.87)**	(6.58)**	(5.81)**	(−6.83)**	(−5.49)**
α_2	−0.15	0.06	0.06	−0.34	−0.38	−2.32	−1.99
	(−1.60)	(0.76)	(0.53)	(−1.82)*	(−2.17)**	(−2.05)**	(−1.74)*
α_3	−0.08	−0.36	−0.00	−0.33	−0.23	0.50	−0.23
	(−0.26)	(−2.09)**	(−0.02)	(−1.51)	(−0.98)	(0.40)	(−0.11)
α_4	−0.26	0.03	0.29	−0.58	−0.65	2.16	3.11
	(−1.95)*	(0.31)	(1.86)*	(−2.59)**	(−2.61)**	(0.95)	(1.52)
F-stat for $\alpha_1 = \alpha_2$	38.30**	6.47**	8.32**	26.07**	31.30**	11.57**	14.44**
p-value	(0.00)	(0.01)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
F-stat for $\alpha_3 = \alpha_4$	0.37	5.57**	0.93	1.21	2.25	0.58	1.80
p-value	(0.55)	(0.02)	(0.33)	(0.27)	(0.13)	(0.44)	(0.18)
R^2	0.29	0.27	0.31	0.28	0.28	0.32	0.31
Obs.	23,466	23,446	17,028	23,446	23,446	18,198	18,198

Regression estimates of equation (5):

$$\begin{aligned} R_{i,t} - R_{B,i,t} = \alpha_0 &+ (\alpha_1 \times UNDERFUNDED + \alpha_2 \times ADEQUATELY\_FUNDED) \times STRONG\_FIRM \\ &+ (\alpha_3 \times UNDERFUNDED + \alpha_4 \times ADEQUATELY\_FUNDED) \times WEAK\_FIRM \\ &+ \alpha_5 \times INDEX_{i,t} + \boldsymbol{\gamma} \cdot \boldsymbol{Z}_{i,t} + \varepsilon_{i,t} \end{aligned}$$

The sample consists of U.S. firms on the NYSE/AMEX/NASDAQ and the CRSP and COMPUSTAT merge files during the June 1988 to June 2017 period. The dependent variable is the Fama–French 25 size and book-to-market ratio adjusted return ($R_{it} - R_{Bt}$). *ADEQUATELY_FUNDED* is a dummy variable equal to 1 if $(PA - PBO)/Mkt\_Eq_{-1}$ is greater than −10%. *UNDERFUNDED* is a dummy variable equal to 1 if $(PA - PBO)/Mkt\_Eq_{-1}$ is less than −10%. Control variables are the same as in equation (1). * and ** indicate significance at the 10% and 5% levels, respectively. Petersen (2009) and Thompson (2011) two-dimension firm and year clustered *t*-statistics are reported.

standards. The first is SFAS 132R, which became effective at the end of 2003. This rule requires annual disclosure of the asset allocation of the pension fund across equities, bonds, real estate, and other assets. The second is SFAS 158, which went into effect at the end of 2006. This rule requires firms to incorporate fair value funding status, or the difference between plan assets and projected benefit obligations in their consolidated statements. Under the original SFAS 87, firms had been allowed to use a smoothing approach to gradually recognize funding status.

To examine whether these two changes in pension accounting standards have any effect on our valuation results, we construct two dummy variables, DUM132R and DUM158. DUM132R takes the value of 1 for fiscal years after December 2003, and 0 otherwise. DUM158 takes the value of 1 for fiscal years after December 2006, and 0 otherwise.

We interact these dummy variables with the right-hand side variables in equation (3), and find that the differential pension valuation effects (for *STRONGER* versus *WEAKER* firms) after passage of SFAS 132R and 158 are greater, consistent with the hypothesis that the disclosure rules increased the flow of value-relevant information to the market. To conserve space, Table 10 reports only the *F*-statistics and p-values for the hypothesis that the pension valuation effects across strong and weak firms are the same in the periods before and after passage of these rules. They indicate that our main hypothesis is unaffected by the change in disclosure rules: both before and after passage, the coefficients on pension variables for *STRONGER* firms are significantly larger than for *WEAKER* firms.

8. Stock market valuation of mandatory contributions

Underfunded plans are required to make contributions to the pension fund. Under ERISA rules, the required contribution is a nonlinear function of funding status ($PA - PBO$). An unfunded liability may be amortized over a period between 5 and 30 years. Rauh (2009) uses IRS 5500 filings to the U.S. Labor Department between 1990 and 1998 to compute funding requirements for the individual pension plans of each firm.

As an alternative to these formulas, which rely in part on data that are publicly available only with a significant lag, Moody (2006) develops a measure for mandatory pension contributions using data from 10-K reports. Campbell *et al.* (2012) use a proxy for mandatory pension contributions based on Moody's definition of required pension contributions. This estimate is nearly identical to Rauh's (2009).

We employ two measures of mandatory contributions. First, according to Moody's (2006), mandatory pension contributions equal the sum of (i) the portion of pension expense earned by employees during the current period (*Serv_Cost*) and (ii) the amortization of any funding shortfall. Therefore, our primary measure of mandatory contribution, *Mand*1, is:

$$Mand1_{i,t} = \begin{cases} Serv\_Cost_{i,t} + (ABO_{i,t} - PA_{i,t})/30 & \text{if } PBO_{i,t} > PA_{i,t} \\ 0 & \text{if } PBO_{i,t} \leq PA_{i,t} \end{cases}$$

where the funding shortfall, $ABO - PA$, is amortized over a 30-year period before 2006. Under PPA 2006, firms must fully fund their pension plans within seven years.[12] An alternative measure, *Mand*2, follows Campbell *et al.* (2012):

$$Mand2_{i,t} = \begin{cases} Serv\_Cost_{i,t} & \text{if } PBO_{i,t} > PA_{i,t} \\ 0 & \text{if } PBO_{i,t} \leq PA_{i,t} \end{cases}$$

During the sample period, from June 1988 to June 2017, there are a total of 23,446 firm-year observations with pension asset and liability data. Among them, 8,597 firm-year observations register a funding shortfall relative to *ABO*, i.e., *ABO* exceeds *PA*. Among these observations, the mean level of funding shortfall amortization $(ABO_{i,t} - PA_{i,t})/30$ is $8.45 million or 0.19% of the beginning-of-fiscal-year market value.

[12]Our valuation results are essentially the same if the funding shortfall is amortized over a 7-year period. Notice that FASB determines funding status using projected benefit obligations, *PBO*, but the mandatory contribution is based on accumulated benefit obligations, *ABO*.

Table 10. Impact of accounting standards on stock market valuation of pension funding

	PRE132R *F*-statistic ($\alpha_1 = \alpha_2$)	POST132R *F*-statistic ($\alpha_3 = \alpha_4$)	PRE158 *F*-statistic ($\alpha_1 = \alpha_2$)	POST158R *F*-statistic ($\alpha_3 = \alpha_4$)
Panel A: Firms partitioned by change in the Ohlson (1980) score				
ΔPA	6.80**	21.71**	7.53**	21.83**
	(0.01)	(0.00)	(0.01)	(0.00)
ΔPBO	3.38*	4.91**	4.52**	6.33**
	(0.07)	(0.03)	(0.03)	(0.01)
ΔABO	1.40	16.10**	2.48	19.06**
	(0.24)	(0.00)	(0.12)	(0.00)
$\Delta Funding\_Status$	0.01	6.36**	0.04	6.45**
	(0.95)	(0.01)	(0.85)	(0.01)
ΔOFF_BAL	0.05	13.90**	0.04	25.68**
	(0.83)	(0.00)	(0.84)	(0.00)
*Mand*1	3.79**	5.62**	9.23**	1.87
	(0.05)	(0.02)	(0.00)	(0.17)
*Mand*2	5.16**	14.74**	11.28**	4.03**
	(0.02)	(0.00)	(0.00)	(0.04)
Panel B: Firms partitioned by change in the Altman (1968) *Z*-score				
ΔPA	7.29**	30.38**	5.85**	38.34**
	(0.01)	(0.00)	(0.02)	(0.00)
ΔPBO	8.68**	3.27*	4.76**	6.26**
	(0.00)	(0.07)	(0.03)	(0.01)
ΔABO	11.35**	24.86**	14.98**	22.33**
	(0.00)	(0.00)	(0.00)	(0.00)
$\Delta Funding\_Status$	0.05	12.78**	0.01	13.59**
	(0.82)	(0.00)	(0.97)	(0.00)
ΔOFF_BAL	0.54	18.38**	0.55	19.94**
	(0.46)	(0.00)	(0.46)	(0.00)
*Mand*1	23.99**	15.81**	28.21**	9.99**
	(0.00)	(0.00)	(0.00)	(0.00)
*Mand*2	17.56**	31.13**	21.04**	14.62**
	(0.00)	(0.00)	(0.00)	(0.00)
Panel C: Firms partitioned by change in the Campbell *et al.* (2008) score				
ΔPA	3.07*	23.94**	4.13**	26.06**
	(0.08)	(0.00)	(0.04)	(0.00)
ΔPBO	6.57**	2.42	5.69**	1.81
	(0.01)	(0.12)	(0.02)	(0.18)
ΔABO	5.32**	12.37**	8.40**	12.20**
	(0.02)	(0.00)	(0.00)	(0.00)
$\Delta Funding\_Status$	0.18	10.53**	0.24	13.22**
	(0.67)	(0.00)	(0.63)	(0.00)
ΔOFF_BAL	0.04	16.24**	0.20	19.50**
	(0.85)	(0.00)	(0.66)	(0.00)
*Mand*1	10.43**	2.04	8.59**	0.15
	(0.00)	(0.15)	(0.00)	(0.69)
*Mand*2	10.38**	5.92**	9.05**	0.85
	(0.00)	(0.02)	(0.00)	(0.36)

The table reports *F*-statistics for the equality of coefficients on pension funding variables for *WEAKER* versus *STRONGER* firms before and after passage of SFAS 132R and 158. The sample consists of U.S. firms on the NYSE/AMEX/NASDAQ and the CRSP and COMPUSTAT merge files during the June 1988 to June 2017 period. Panel A reports the results of regressing the Fama–French 25 size and book-to-market ratio adjusted excess returns ($R_{it} - R_{Bt}$) on pension-funding variables. The pension funding variables (ΔPA, ΔPBO, ΔABO, $\Delta Funding\_Status$, and ΔOFF_BAL) are included one-at-a-time in each equation. The regression specifications for the two SFAS rules are:

$$\begin{aligned}R_{i,t} - R_{B,i,t} &= \alpha_0 + (\alpha_1 \times STRONGER \times PRE132R + \alpha_2 \times WEAKER \times PRE132R)\Delta Y_{i,t} \\ &\quad + (\alpha_3 \times STRONGER \times POST132R + \alpha_4 \times WEAKER \times POST132R)\Delta Y_{i,t} \\ &\quad + \alpha_5 \times INDEX_{i,t} + \boldsymbol{\gamma} \cdot \mathbf{Z}_{i,t} + \varepsilon_{i,t}, \\ R_{i,t} - R_{B,i,t} &= \alpha_0 + (\alpha_1 \times STRONGER \times PRE158 + \alpha_2 \times WEAKER \times PRE158)\Delta Y_{i,t} \\ &\quad + (\alpha_3 \times STRONGER \times POST158 + \alpha_4 \times WEAKER \times POST158)\Delta Y_{i,t} \\ &\quad + \alpha_5 \times INDEX_{i,t} + \boldsymbol{\gamma} \cdot \mathbf{Z}_{i,t} + \varepsilon_{i,t}.\end{aligned}$$

The *STRONGER* and *WEAKER* dummy variables indicate firms that experience improvement or deterioration in financial strength using three alternative measures of financial condition. *PRE132R* takes the value of 1 for fiscal years before December 2003, and 0 otherwise. *POST132R* equals one minus *PRE132R*. *PRE158* takes the value of 1 for fiscal years before December 2006, and 0 otherwise. *POST158* equals one minus *PRE158*. The *F*-statistic is computed for the null hypothesis that the regression coefficients for *STRONGER* × *PRE132R* and *WEAKER* × *PRE132R* are the same ($\alpha_1 = \alpha_2$). The *F*-statistic for the null hypothesis that the regression coefficients for *STRONGER* × *POST132R* and *WEAKER* × *POST132R* are the same ($\alpha_3 = \alpha_4$) is also reported. Similar tests are implemented for *PRE158* and *POST158* dummies. Other control variables and industry dummies are included in the regressions. * and ** indicate significance at the 10% and 5% levels, respectively. Petersen (2009) and Thompson (2011) two-dimension firm and year clustered *t*-statistics are reported.

The summary statistics from Table 1 show that mandatory contributions are estimated to be 0.5% and 0.4% of the beginning of the fiscal year market value, respectively, under the two alternative measurements. The two measures of mandatory contributions are highly correlated (correlation coefficient = 0.96). We estimate and report in Table 11 the following regression equation for the stock market valuation of mandatory contributions:

$$R_{i,t} - R_{B,i,t} = \alpha_0 + (\alpha_1 \times STRONGER + \alpha_2 \times WEAKER)Mand_{i,t} + \alpha_3 \times INDEX_{i,t} + \boldsymbol{\gamma} \cdot \boldsymbol{Z_{i,t}} + \varepsilon_{i,t}. \tag{6}$$

Since the results for the two measures of mandatory contributions are essentially the same, we will focus our discussion on the first measure, *Mand*1.

When the Δ*Ohlson_BS* index is used to partition firms, the estimated coefficients (t-statistic) from equation (6) for *STRONGER* × *Mand*1 and *WEAKER* × *Mand*1 are −5.85 (−3.84) and −2.01 (−3.06), respectively. Thus, the value impact of mandatory contributions is substantially greater than one-for-one. This reaction to mandatory contributions is similar to estimates reported in Moody (2006) and Campbell *et al.* (2012). As noted above, these high regression coefficients presumably reflect the fact that, unlike other pension items, changes in mandatory contributions generally signify an entire annuity of additional required payments.

The high sensitivity to mandatory contributions particularly characterizes *STRONGER* firms. Distressed firms exhibit far lower coefficients. The differential impact using the Δ*Ohlson_BS* index to partition firms is highly significant with an F-statistic (p-value) of 5.80 (0.02) for *Mand*1 and 10.81 (0.00) for *Mand*2. The results are quite similar with even higher F-statistics using the other indexes of financial health. Therefore, the empirical evidence provides strong support for the differential valuation of mandatory contributions between *STRONGER* and *WEAKER* firms. This is consistent with the hypothesis that the market values the entire stream of required payments for healthy firms, but is less confident that distressed firms will actually make all the stipulated payments.

9. Size, financial distress, and pension valuation

Small firms tend to fall on hard times and face more difficulties in distressed times (Ofek, 1993).[13] Griffin and Lemmon (2002) employ Ohlson's (1980) bankruptcy score to conclude that firms in the highest default probability quintile tend to be small. While firms can hedge to reduce the variance of firm value, and thereby the expected costs of financial distress, Nance *et al.* (1993) argue that small firms have lower incentives to hedge, and are in fact less likely to do so.

In this section, we first examine the persistence properties of large and small firms. Every year, all firms are sorted independently based on size as well as change in financial conditions. Firms are classified into *LARGE* and *SMALL* groups based on market capitalization and then into *STRONGER* and *WEAKER* subgroups. For each subsample of *LARGE* and *SMALL* firms, panel A of Table 12 reports the average slope coefficients from equation (2), Fama and MacBeth (1973) t-statistics, and Wilcoxon signed rank test statistics and corresponding p-values for the two partitioned subgroups *STRONGER* and *WEAKER*.

Overall, the earlier conclusion that changes in financial condition are highly persistent is preserved in panel A of Table 12. In all cases, c_0 is positive and highly significant for *STRONGER* firms and negative and significant for *WEAKER* firms. As in Table 10, c_1 is generally negative, but far too small to reverse the persistence that results from the intercepts. Moreover, for both *STRONGER* and *WEAKER* firms, *SMALL* firms on average exhibit stronger persistence than *LARGE* firms. The absolute value of c_0 is larger for *SMALL* than for *LARGE* firms, while the absolute value of c_1 is comparable.

[13] Ofek (1993) finds that larger firms often operate within several lines of business and different geographic regions and may be better able to restructure at the onset of distress, whereas smaller firms are more limited.

Table 11. Stock market valuation of mandatory contributions

	INDEX = Δ*Ohlson_BS*		*INDEX* = Δ*Altman_Z*		*INDEX* = Δ*CHS_BS*	
	*Mand*1	*Mand*2	*Mand*1	*Mand*2	*Mand*1	*Mand*2
α_1 : *STRONGER* × Δ*Y*	−5.85	−7.73	−6.29	−8.03	−5.61	−7.30
	(−3.84)**	(−4.53)**	(−4.29)**	(−5.48)**	(−3.80)**	(−4.12)**
α_2 : *WEAKER* × Δ*Y*	−2.01	−1.79	−0.55	0.58	−1.68	−1.07
	(−3.06)**	(−2.12)**	(−0.52)	(0.56)	(−2.14)**	(−1.03)
α_3 : *INDEX*	0.04	0.04	0.10	0.10	0.20	0.20
	(3.96)**	(3.87)**	(3.50)**	(3.50)**	(5.90)**	(5.78)**
F-statistic for $\alpha_1 = \alpha_2$	5.80**	10.81**	25.93**	36.76**	6.01**	11.82**
(p-value)	(0.02)	(0.00)	(0.00)	(0.00)	(0.01)	(0.00)
Control variables	Yes	Yes	Yes	Yes	Yes	Yes
Industry dummy	Yes	Yes	Yes	Yes	Yes	Yes
R^2	0.24	0.24	0.26	0.26	0.31	0.31
Obs.	18,198	18,198	18,198	18,198	18,198	18,198

Regression estimates of equation (6):

$$R_{i,t} - R_{B,i,t} = \alpha_0 + (\alpha_1 \times STRONGER + \alpha_2 \times WEAKER)Mand_{i,t} + \alpha_3 \times INDEX_{i,t} + \boldsymbol{\gamma} \cdot \mathbf{Z}_{i,t} + \varepsilon_{i,t}.$$

The sample consists of U.S. firms on the NYSE/AMEX/NASDAQ and the CRSP and COMPUSTAT merge files during the June 1988 to June 2017 period. The table reports the results of regressing the Fama–French 25 size and book-to-market ratio adjusted excess returns ($R_{it} - R_{Bt}$) on mandatory pension contributions. *Mand*1 and *Mand*2 are included one-at-a-time in each equation. The following indices are used to partition firms into *STRONGER* and *WEAKER* groups: Δ*Ohlson_BS*, Δ*Altman_Z*, and Δ*CHS_BS*. The table reports estimated coefficients on the interactive terms *STRONGER* × Δ*Y* and *WEAKER* × Δ*Y*. Δ*Y* denotes *Mand*1 or *Mand*2. *F*-statistics are for the test of the null hypothesis that the regression coefficients for *STRONGER* and *WEAKER* firms are the same, i.e., $\alpha_1 = \alpha_2$. Other control variables and industry dummies are included in the regressions. * and ** indicate significance at the 10% and 5% levels, respectively. Petersen (2009) and Thompson (2011) two dimension firm and year clustered *t*-statistics are reported.

Second, we examine differential pension valuation effects when firms are partitioned by both size and changes in financial conditions. We estimate the following regression:

$$\begin{aligned} R_{i,t} - R_{B,i,t} &= \beta_0 + (\beta_1 \times LARGE \times STRONGER + \beta_2 \times LARGE \times WEAKER \\ &\quad + \beta_3 \times SMALL \times STRONGER + \beta_4 \times SMALL \times WEAKER)\Delta Y_{i,t} \\ &\quad + \beta_5 \times INDEX_{i,t} + \boldsymbol{\gamma} \cdot \mathbf{Z}_{i,t} + \varepsilon_{i,t} \end{aligned} \tag{7}$$

where the dummy variable *LARGE* × *STRONGER* equals 1 for large firms that are financially healthier than in the previous year while *SMALL* × *WEAKER* equals 1 for small firms that weaken. As before, *INDEX* refers to the partitioning variables, Δ*Ohlson_BS*, Δ*Altman_Z*, Δ*CHS_BS*, respectively. Δ*Y* refers to changes in the following seven pension variables, Δ*PA*, Δ*PBO*, Δ*ABO*, Δ*Funding_Status*, Δ*OFF_BAL*, *Mand1*, and *Mand2*, respectively.

We formally implement empirical tests for the following four null hypotheses. The first is that regression coefficients for the *LARGE* × *STRONGER* and *LARGE* × *WEAKER* firms are the same, i.e., $\beta_1 = \beta_2$. The second null is that regression coefficients for the *SMALL* × *STRONGER* and *SMALL* × *WEAKER* firms are the same, i.e., $\beta_3 = \beta_4$. The third is that regression coefficients for the *LARGE* × *STRONGER* and *SMALL* × *STRONGER* firms are the same, i.e., $\beta_1 = \beta_3$. The fourth is that regression coefficients for the *LARGE* × *WEAKER* and *SMALL* × *WEAKER* firms are the same, i.e., $\beta_2 = \beta_4$.

We observe the following patterns from panel B of Table 12. First, for the majority of partitions and the majority of the seven pension variables (the five pension asset and liability variables and two mandatory contribution measures), the differential valuations between *SMALL* × *STRONGER* and *SMALL* × *WEAKER* firms are highly significant. The second hypothesis ($\beta_3 = \beta_4$) can be rejected at the 5% level in 20 out of 21 cases (three partitioning criteria × seven pension variables). In other words, we find strong support for differential valuation effects among small firms. In contrast, for large firms, the null hypothesis of no differential valuation effects ($\beta_1 = \beta_2$) is rejected in only 7 out of 21 cases. The result for the third hypothesis ($\beta_1 = \beta_3$, i.e., healthy large firms versus healthy

Table 12. Size effects in pension valuation

	LARGE			*SMALL*		
	STRONGER	*WEAKER*	*Z*-statistic (p-values)	*STRONGER*	*WEAKER*	*Z*-statistic (p-values)
Panel A: Persistence properties for firms partitioned by size and financial conditions						
INDEX = Δ*Ohlson_BS*						
Mean of c_0	0.51	−0.60	6.54 (0.00)**	0.77	−0.82	6.54 (0.00)**
t-statistic	(25.23)**	(−25.64)**		(24.68)**	(−25.20)**	
Mean of c_1	−0.26	−0.12	−4.53 (0.00)**	−0.23	−0.17	−2.08 (0.04)**
t-statistic	(−12.15)**	(−6.04)**		(−8.68)**	(−5.27)**	
INDEX = Δ*Altman_Z*						
Mean of c_0	0.25	−0.30	6.54 (0.00)**	0.40	−0.43	6.54 (0.00)**
t-statistic	(14.57)**	(−12.39)**		(19.06)**	(−19.81)**	
Mean of c_1	−0.07	−0.07	0.43 (0.67)**	−0.23	0.01	−3.46 (0.00)**
t-statistic	(−1.69)*	(−1.44)		(−4.41)**	(0.13)	
INDEX = Δ*CHS_BS*						
Mean of c_0	0.26	−0.36	6.54 (0.00)**	0.43	−0.51	6.54 (0.00)**
t-statistic	(16.60)**	(−9.57)**		(23.26)**	(−15.05)**	
Mean of c_1	−0.29	−0.05	−4.30 (0.00)**	−0.28	−0.07	−5.36 (0.00)**
t-statistic	(−7.83)**	(−1.18)		(−10.02)**	(−5.07)**	

	Partitioning criterion		
	Δ*Ohlson_BS*	Δ*Altman_Z*	Δ*CHS_BS*
Panel B: Number of pension variables for which the null hypothesis is rejected			
Test for $\beta_1 = \beta_2$	0	5	2
Test for $\beta_3 = \beta_4$	7	7	6
Test for $\beta_1 = \beta_3$	6	6	6
Test for $\beta_2 = \beta_4$	1	3	4

Panel A: Regression estimates of equation (2), which tests for persistence of changes in financial strength, with firms stratified according to size. The sample consists of U.S. firms on the NYSE/AMEX/NASDAQ and the CRSP and COMPUSTAT merge files during the June 1988 to June 2017 period. Every year, all firms are sorted independently based on size and change in financial condition. Firms are classified into *LARGE* and *SMALL* groups based on market capitalization. Firms are further sorted into the *STRONGER* and *WEAKER* groups based on the change in financial condition indices (Δ*Ohlson_BS*, Δ*Altman_Z*, Δ*CHS_BS*). For *LARGE* and *SMALL* groups respectively, panel A estimates the persistence of changes in financial condition indices by running the following cross-sectional regression:

$$INDEX_{i,t} = c_0 + c_1 INDEX_{i,t}-1 + \varepsilon_{i,t}$$

where $INDEX_{i,t}$ is one of the three measures capturing the change in firms' financial conditions. Panel A reports the annual average of the slope coefficients, c_0 and c_1, and their Fama and MacBeth (1973) *t*-statistics. The panel also reports the Wilcoxon (1945) signed rank test statistic *Z* and the corresponding p-values for the null hypothesis that the slope coefficients from the *STRONGER* and *WEAKER* groups are the same. * and ** indicate significance at the 10% and 5% levels, respectively.
Panel B: Estimates of equation (7), which tests for size effects in pension valuation. The dependent variable is the Fama–French-25 size and book-to-market ratio adjusted annual return ($R_{it} - R_{Bt}$). The independent variables include pension-funding variables. Pension variables, including changes in pension assets and liabilities (Δ*PA*, Δ*PBO*, Δ*ABO*, Δ*Funding_Status*, and Δ*OFF_BAL*, *Mand*1, and *Mand*2), are included one-at-a-time in each equation. The table presents tests for the following four null hypotheses: (i) regression coefficients for the *LARGE* × *STRONGER* and *LARGE* × *WEAKER* firms are the same, i.e., $\beta_1 = \beta_2$; (ii) regression coefficients for the *SMALL* × *STRONGER* and *SMALL* × *WEAKER* firms are the same, i.e., $\beta_3 = \beta_4$; (iii) regression coefficients for the *LARGE* × *STRONGER* and *SMALL* × *STRONGER* firms are the same, i.e., $\beta_1 = \beta_3$; and (iv) regression coefficients for the *LARGE* × *WEAKER* and *SMALL* × *WEAKER* firms are the same, i.e., $\beta_2 = \beta_4$. For each of the three partitioning indices and two fundamental variables, the table reports the number of pension variables for which the null hypothesis is rejected at the 5% level.

small firms) is also strong, with the null hypothesis rejected in 18 of 21 cases. However, the fourth null hypothesis that there are no differential valuation effects between distressed large firms and distressed small firms ($\beta_2 = \beta_4$) is rejected less consistently, in only 8 of 21 cases.

10. Conclusion

Despite significant progress in the literature on valuation issues surrounding defined benefit pension plans, some questions concerning defined benefit corporate pension plan property rights remain unanswered. Pension economics and legislative efforts both suggest that plan sponsors, plan beneficiaries, and the PBGC share claims to pension plan assets. But the empirical evidence so far has not allowed for definitive conclusions on how those claims are perceived to be allocated. A potentially determinative issue typically not addressed in these studies is that perceived property rights may depend on the firm's financial strength. Default risk has important implications for pension-plan property rights in light of PBGC and bankruptcy put options. Moreover, the stock market may attach differential importance to changes in plan assets or plan liabilities between consecutive fiscal years because managers have different funding contribution and financial reporting incentives in good and bad times. Managers can strategically allocate assets or manipulate actuarial assumptions depending on their firm's financial outlook.

This paper differs from earlier studies on pension valuation in two important ways. First, we focus on changes in market value and pension funding rather than on levels. This procedure improves the signal-to-noise ratio and the power of our tests. Second, we implement a conditional valuation study, where valuation effects depend on a firm's financial status. We construct three indices based on the bankruptcy scores of Ohlson (1980), Altman (1968), and Campbell *et al.* (2008) to measure the evolution of firms' financial conditions between consecutive fiscal years. Our simple indices, which measure changes in financial condition, are significantly positively correlated and consistently partition firms into either financially stronger or financially weaker groups.

We find that on average, firms that have recently become financially healthier continue to improve. Similarly, firms that have become financially weaker tend to weaken further. The fact that recent changes in financial condition seem to continue into the next year has important implications for the valuation of pension assets and liabilities. As firms that have become more distressed further deteriorate, they are more prone to renege on pension obligations.

We estimate pension valuation models for stronger versus weaker firms and for strengthening and weakening firms and test for differential valuation effects. We show that excess stock returns of healthy firms strongly track changes in pension assets and liabilities, but that such tracking is not evident in firms facing higher likelihoods of financial distress. For healthy firms, the evidence suggests that the stock market views the ownership claim of the sponsoring firm at somewhere between 50% and 100%, depending on the particular pension asset or liability under consideration. But implied ownership claims for distressed firms are generally not statistically distinguishable from zero. This evidence therefore implies that the stock market valuation of pension assets and liabilities is dynamic and depends critically on the evolution of the individual firms' financial status. In this regard, our empirical results provide important context and new evidence on the long-standing question of the effective ownership of corporate defined benefit pension plans.

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Appendix A

Construction of pension variables, stock returns, accounting ratios, and financial distress measures

This appendix provides the definitions, references, and details of the COMPUSTAT accounting items used to construct the pension variables, stock return variables, and accounting variables, and the financial distress measure.

Variable name and references	COMPUSTAT items
Pension variables	
Plan assets (*PA*)	*PA* = pension plan assets + underfunded pension plan assets = *PPLAO* + *PPLAU*
Plan benefit obligations (*PBO*)	*PBO* = projected benefit obligations + underfunded projected benefit obligations = *PBPRO* + *PBPRU*
Funding status (*Funding_Status*)	*Funding_Status* = plan assets – projected benefit obligations = *PA* – *PBO*
Accumulated benefit obligations (*ABO*)	*ABO* = accumulated benefit obligations + underfunded accumulated benefit obligations = *PBACO* + *PBACU*
Mandatory contributions (*Mand*1)	*Mand*1 = service cost + minimum pension liabilities/30 = *Serv_Cost* + *MPL*/30 if *PBO* > *PA*; *Mand1* = 0 if otherwise. *MPL* = minimum pension liabilities = *ABO* – *PA* if *PBO* > *PA*; *MPL* = 0 if otherwise.
Mandatory contributions (*Mand*2)	*Mand*2 = service cost = *Serv_Cost* if *PBO* > *PA*; *Mand2* = 0 if otherwise.
Off balance sheet asset/liability (*OFF_BAL*)	*OFF_BAL* = unrecognized gain and loss + unrecognized prior service cost = *Unreg_GL* + *Unreg_SC*
Unrecognized gain and loss (*Unreg_GL*)	*Unreg_GL* = pension other adjustments + underfunded pension other adjustments = *POAJO* + *POAJU*
Unrecognized prior service cost (*Unreg_SC*)	*Unreg_SC* = pension unrecognized prior service cost + underfunded pension unrecognized prior service cost = *PCUPSO* + *PCUPSU*
Service cost (*Serv_Cost*)	*Serv_Cost* = pension plans service cost = *PPSC*
Periodic pension cost (*PC*)	*PC* = periodic pension cost = *PPC*
Stock return variables	
R_{it}	Cumulative monthly return over fiscal year *t* for firm *i*
$R_{it} - R_{Bt}$	Cumulative monthly return on individual stocks minus cumulative returns on size and book to market matched portfolio returns during the same fiscal year

Accounting variables	
Book value (*BE*)	*BE* = Total assets – liabilities – book value of preferred stocks + balance sheet deferred taxes and investment tax credit = *AT* – *LT* – *PSTKL* + *TXDITC*
Market size (*Mkt_Eq*)	*Mkt_Eq* = market equity in June at end of year *t* = June-end stock price × common shares outstanding = June-end stock price × *CASHO*
Cash holdings (*CASH*)	*CASH* = cash and short-term investments = cash + short-term investments = *CHE*
Interest expenses (*INT*)	*INT* = interest and related expenses = *XINT*
Earnings before interest expenses and taxes (*EBIT*)	*EBIT* = income before extraordinary items + interest + deferred income taxes and investment tax credits = *IB* + *XINT* + *TXDITC*

Total assets minus cash and cash equivalents (*Non-Cash_Assets*)	$Non\text{-}Cash\_Assets$ = total assets − cash holding = $AT - CASH$ $CASH$ = cash and short-term investments = CHE
Accrual (*ACCRUALS*)	$ACCRUALS$ = (Δcurrent assets − Δcash and cash equivalents) − (Δcurrent liabilities − Δdebt included in current liabilities − Δincome taxes payable) − depreciation and amortization expenses $= (\Delta ACT - \Delta CHE) - (\Delta LCT - \Delta DLC - \Delta TXP) - DP$
Asset growth (*ASST_GWTH*)	$ASST\_GWTH = AT/AT(-1) - 1$ AT = total assets
Net external financing (*XFIN*)	$XFIN$ = long-term debt − long-term debt reduction + sale of common and preferred stock − purchase of common and preferred stock $= DLTIS - DLTR + SSTK - PRSTKC$
Market leverage (*MLEV*)	$MLEV$ = (long-term debt + short-term debt)/(long-term debt + short-term debt + market equity at fiscal year end) $= (DLTT + DLC)/(DLTT + DLC + ME\_F)$ Market equity at fiscal year end = $ME\_F = PRCC\_F \times CHSO$
Financial distress measures	
Ohlson's (1980) bankruptcy score (*Ohlson_BS*)	$Ohlson\_BS = -1.320 - 0.407\log(AT) + 6.030\left(\frac{LT}{AT}\right) - 1.430\left(\frac{ACT-LCT}{AT}\right) + 0.076\left(\frac{LCT}{ACT}\right) - 2.370\left(\frac{NI}{AT}\right) - 1.830\left(\frac{FO}{LT}\right) - 0.521\left(\frac{NI-NI_{-1}}{\lvert NI\rvert + \lvert NI_{-1}\rvert}\right) - 1.720 \times D1 + 0.285 \times D2$ $D1 = 1$ if $LT > AT$; $D1 = 0$ if otherwise. $D2 = 1$ if $NI < 0$ for the last two years; $D2 = 0$ if otherwise. AT = total assets LT = total liabilities ACT = current assets LCT = current liabilities FO = funds from operation = $IBC + DPC + TXDC + FOPO$ IBC = net income before extraordinary items DPC = depreciation and amortization $TXDC$ = deferred income taxes $FOPO$ = total other cash flow
Change in financial health (distress) index based on *Ohlson_BS* (Δ*Ohlson_BS*)	$\Delta Ohlson\_BS_{i,t} = -(Ohlson\_BS_{i,t} - Ohlson\_BS_{i,t-1})$
Altman's (1968) *Z*-score. We use the estimates from Table 2 of Shumway (2001) and add a negative sign. (*Altman_Z*)	$Altman\_Z = -\left[1.200\left(\frac{ACT-LCT}{AT}\right) + 0.600\left(\frac{RE}{AT}\right) + 10.000\left(\frac{EBIT}{AT}\right) + 0.050\left(\frac{ME}{LT}\right) - 0.470\left(\frac{SALE}{AT}\right)\right]$ ACT = current assets LCT = current liabilities RE = retained earnings AT = total assets LT = total liabilities $SALE$ = sales revenue $EBIT$ = income before extraordinary items + interest + deferred income taxes and investment tax credits $= IB + XINT + TXDITC$ ME = market equity at the end of fiscal year = end of fiscal year stock price × common shares outstanding $= PRCC\_F \times CASHO$
Change in financial health (distress) index based on *Altman_Z* (Δ*Altman_Z*)	$\Delta Altman\_Z_{i,t} = -(Altman\_Z_{i,t} - Altman\_Z_{i,t-1})$
Campbell *et al.* (2008) bankruptcy score (*CHS_BS*)	$CHS\_BS = -9.164 - 20.264 \times NIMTAAVG + 1.416 \times TLMTA - 7.129 \times EXRETAVG + 1.411 \times SIGMA - 0.045 RSIZE - 2.132 \times CASHMTA + 0.075/BM - 0.058 \times PRICE$

	NIMTAAVG = moving average of *NIMTA* *NIMTA* = quarterly profitability = net income/(market equity for the fiscal quarter + total liabilities) = *NIQ*/(*Price_F* × *CASHO* + *LTQ*) *TLTMA* = quarterly market leverage = total liabilities/(market equity for the fiscal quarter + total liabilities) = LTQ/(Price_F × CASHO + LTQ) *EXRETAVG* = moving average of *EXRET* Excess return = $EXRET_{i,t} = \log(1 + R_{i,t}) - \log(1 + R_{SP500,t})$ *SIGMA* = annualized 3-month rolling sample standard deviation *RSIZE* = relative size = $\log(ME/ME_{SP500})$ *CASHMTA* = quarterly liquidity = cash and short-term investments/(market equity for the fiscal quarter + total liabilities) = *CHEQ*/(*Price_F* × *CASHO* + *LTQ*) *BM* = book equity/market equity = (*CEQ* + *TXDITCQ*)/(*Price_F* × *CASHO*) *CEQ* = common equity *TXDITCQ* = deferred taxes and investment tax credit *PRICE* = log(*CRSP* stock price per share)
Change in financial health (distress) index based on *CHS_BS* (Δ*CHS_BS*)	$\Delta CHS\_BS_{i,t} = -(CHS\_BS_{i,t} - CHS\_BS_{i,t-1})$

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