Underinvestment and Overinvestment: The Market's View

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First Version November 24, 2006
This Version April 24, 2007

ABSTRACT

Mandatory contributions to defined benefit pension plans provide an original strategy to identify price reactions to a shift in internal resources controlling for investment opportunities. The price impact of an exogenous reduction in cash flows is larger for firms that rank higher according to a number of indices of financial constraints, consistent with an economically significant effect of costly external finance on investment. On the other hand, price reactions to mandatory contributions are smaller in absolute value for firms protected by more anti-takeover provisions, suggesting that the market expects poor governance to significantly reduce the value of cash flows. The evidence supports theories of both under- and overinvestment. The comparison of the two effects suggests that the market considers underinvestment to be the prevalent distortion for a large majority of firms.

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In interpreting the empirical literature on corporate investment, Jeremy Stein (2003) argues that while it is hardly questionable that financial slack matters for investment, it is less clear to what extent this relationship is due to financing constraints or to empire building. On the one hand, by raising the costs of external funds, financial frictions may cause a sub-optimal level of investment.[1] This makes investment sensitive to the availability of cheap internal funds. On the other hand, if managers have empire-building preferences (Jensen (1986)), they will use free cash flows to fund investment projects beyond the level that maximizes shareholders' value. Like costly external finance, also this argument leads to the prediction that investment is increasing in internal resources. While the two theories are observationally equivalent in this respect, their policy implications are obviously different. Furthermore, the two hypothesis may very well coexist in a unified model that admits both under- and overinvestment. Then, the more interesting question is which distortion prevails empirically.

The goal of this paper is to address this question. More broadly, the paper aims at quantifying the absolute and relative importance of financial constraints and empire-building in an empirical frameworks that allows a unified treatment of the two distortions.

One obvious dimension along which the empirical predictions of the two set of theories differ is the value that the market attaches to internal funds. In an efficient market, an unexpected cash flow drain should trigger a larger price drop in a firm that faces costly external finance than in a firm that is unconstrained. For the constrained company, the shadow price of cash flows includes the higher cost of external finance and the net present value (NPV) of the investment projects that are forsaken because of the disappearance of these funds. On the other hand, the same drop in internal resources is less costly for the investors of firms where managers are prone to overinvest. In these companies, one needs to adjust the face value of internal resources by the negative NPV of the wasteful projects that empire-building managers would undertake.

These considerations suggest the empirical strategy of the paper. After identifying a source of unexpected cash flow decrease, one can study the price reaction to this shock across *a priori* measures of financial constraints and measures of potential conflicts between managers and shareholders. The evidence that price reactions are magnified for firms that are supposedly more constrained would suggest that some companies are indeed underinvesting. Moreover, finding that post-event returns are attenuated for firms that are more afflicted by agency conflicts would suggest

[1] The theoretical literature has motivated financial constraints on the basis of both asymmetric information (Myers and Majluf (1984), Greenwald, Stiglitz, and Weiss (1984)) and moral hazard (Jensen and Meckling (1976), Grossman and Hart (1982), Stulz (1990), Hart and Moore (1995)).

that the market anticipates overinvestment. Given that incentives for under- and overinvestment may coexist even within the same firm, one can compute the aggregate price reaction to the shock to find which distortion prevails. Then, the study of the cross-sectional distribution of the aggregate price reactions allows to determine which distortion is globally more important for the firms in the sample.

This simple empirical approach is complicated by an identification issue. Following the cash flow drop, one should expect heterogeneity in price reactions across firms as a function of the different investment opportunities that they face. Specifically, in firms with poor investment prospects the value of cash flows is naturally lower. This effect could muddle identification if investment opportunities correlate with the chosen measures of financial constraints and agency conflicts. For example, a firm may be classified as financially constrained if it has a large need of funds (Hennessy and Whited (2007)). This need in turn can result from the availability of profitable investment opportunities. For this firm, the shadow price of cash flows is magnified by both investment opportunities and financial constraints. Identification requires the separation of the two effects.[2]

The paper draws the solution to this identification problem from the literature on investment-cash flow sensitivities. Some recent work by Rauh (2006) focuses on mandatory contributions to defined benefit pension plans as a source of variation in internal resources. These contributions are determined as a non-trivial function of the funding status of the company pension plans. The funding status, in turn, is a choice variable of the firm, which depends on investment opportunities. Thanks to the regression discontinuity approach, by controlling for functions of the plan funding status, one absorbs the effect of investment opportunities. Consequently, one can use the orthogonal component of mandatory contributions to capture independent variation in cash flows.

This paper adopts Rauh's approach and uses mandatory contributions as a source of cash flow variation, while controlling for investment opportunities through the pension plan funding status. The dependent variable is the return in the year that follows the contributions. To absorb the effect of known determinants of returns, I control for beta, size, book-to-market, momentum, and other variables. This identification strategy distinguishes the paper from previous studies that look at price reactions but do not deal with the problem of omitted investment opportunities.

To capture financial constraints, I rank firms according to three indices that are standard in the literature: the Whited and Wu (2006) index, the Kaplan and Zingales (1997) index, and the

[2]In the context of overinvestment, a comparable example is the following. Managers of firms that lack investment opportunities may have more incentives to divert cash. Here, the severity of agency conflicts and the lack of investment prospects both contribute to reduce the value of cash flows.

Cleary (1999) index. Instead, to capture potential exposure to empire-building I refer to the studies that identify an effect of corporate governance on performance. In this context, Gompers, Ishii, and Metrick (2003), Bebchuck, Cohen, and Ferrell (2004), and Bebchuck and Cohen (2005) develop some indices that are meant to measure the extent of agency conflicts between managers and shareholders. In particular, they focus on the anti takeover provisions that shield managers from the disciplinary role of the market for corporate control.

The results of the analysis are easily summarized. The market recognizes an economically important role for financial constraints. Specifically, a drop in cash is more costly in constrained firms, irrespectively of the chosen measure of financial constraints. According to a rough estimate, a cash decrease equal to 1% of market value reduces the stock price by about 3.6% for the average constrained firm, while the effect on the value of an unconstrained company merely reflects the initial loss of cash (that is, 1%). These estimates seem consistent with realistic investment behavior, as a simple back-of-the-envelope calculation suggests that the internal rate of return of the forsaken projects by the average constrained firm ranges between 10.6% and 20.8%. The results are robust to alternative definitions of the indices, to controls for the information surrounding the firm, and to different sample selection criteria. Further supporting evidence comes from a trading strategy, which extends the results in Franzoni and Marin (2006) using financial constraints as an additional sorting variable.

On the other hand, investors also seem to anticipate that empire-building is an important determinant of investment for some firms. Specifically, a loss of cash in a firm with worse corporate governance is valued less by the market. For the average firm in the sample, a drop in free cash flows of 1% of market capitalization prevents a waste of resources that ranges between 4.6% and 14% of company value, depending on the governance index one uses. Given the sample of firms for which the governance indices are available is tilted towards large companies, one should expect this number to overestimate the impact of empire building on the average listed firm. Still, the findings suggest that overinvestment is an economically significant distortion in investors' perception.

When comparing the incidence of under- and overinvestment for the firms in the sample, the conclusion is that the first distortion is prevalent. Depending on the chosen financial constraints index, for a fraction of firm-year observations between 58% and 73% the aggregate price reaction to a drop in cash flow is magnified by the presence of financial constraints rather than attenuated by potential empire-building. Again, given the bias of this sample towards large firms, the prevalence of underinvestment should be even larger in more representative panels.

The approach of the paper is to measure the market's assessment of these distortions without taking a stance on the issue of market efficiency. In reality, price formation may be deeply affected by investors' irrationality. If that is the case, the measured price reactions do not capture the true long run impact on firm value of credit constraints and empire building. In this scenario, the results of the paper should be interpreted as a snapshot of investors' subjective view on these distortions.

Finally, the paper's interpretation of the evidence hinges on the assumption that the chosen indices of financial constraints and governance actually capture heterogeneity in the exposure to under- and overinvestment. To corroborate the validity of this assumption, the paper looks at some descriptive evidence on the firms that, based on the aggregate price reaction, rank differently in the spectrum from under- to overinvestment. The firms that are predicted to overinvest are larger, older, have more coverage by analysts and rating agencies, and generate more cash flow than the companies for which underivestment is the prediction. The fact that these statistics correspond to the theoretical presumption on the identity of firms that under/overinvest is reassuring about the empirical soundness of the proposed interpretation.

The literature on financial constraints is abundant. Following Fazzari, Hubbard, and Petersen (1988), a strand of literature has adopted reduced-form regressions of investment on cash flow, using average Tobin's Q to control for investment opportunities. An alternative approach, which circumvents the use of Q, relies on natural experiments to identify exogenous variation in internal resources (Blanchard, Lopez-de-Silanes, and Shleifer (1994), Lamont (1997)).[3] In this context, Rauh's (2006) study represents the first natural experiment in a large sample setting. This paper is closely related to Rauh, as it adopts the same identification device. Moreover, the finding that price reactions triggered by pension contributions are larger for constrained firms can be interpreted as the projection in the space of asset prices of the real effect identified in Rauh's paper.

The marginal contribution of this work to the financing constraints literature is the focus on price reactions. The empirical studies on financial frictions have focused on a wide range of investments that, besides expenditure on plant and equipment, include R&D, inventories, pricing for market share, and labor hoarding during recessions.[4] The advantage of the methodology in this paper is that is provides a comprehensive measure of the costs of external finance and the value of forsaken investment that includes quantities that are not otherwise measurable (i.e human capital and other intangibles). In this sense, while it is a reduced form approach, it achieves a comparable result to

[3]Erickson and Whited (2000) and Gomes (2001) point out the potential failures of average Q to control for investment opportunities due to measurement error.

[4]See Stein (2003) for references.

the structural models of investment (Hennessy and Whited (2007)).

Like this paper, a number of previous studies look at price reactions after corporate events across levels agency conflicts and finds evidence in favor of Jensen's (1986) free cash flow hypothesis. These papers focus on returns after the announcement of capital expenditures (McConnell and Muscarella (1985)), acquisitions (Lewellen, Loderer, and Rosenfeld (1985), Morck, Shleifer, and Vishny (1990), Lang, Stulz, and Walkling (1991), and Masulis, Wang, and Xie (2007)), and dividend initiations (Pinkowitz, Stulz, and Williamson (2006), and Officer (2007)). The fundamental difference in the approach in this paper relative to this previous work is the fact that the research design allows to control for the omitted investment opportunity set, which is an obvious determinant of post-event price reactions, and is potentially correlated with the chosen measure of agency conflicts.[5]

Closely related to this paper, a number of studies focus on the marginal value of cash and find evidence of financial constraints (Pinkowitz and Williamson (2004), Faulkner and Wang (2005)) and corporate governance effects (Dittmar and Mahrt-Smith (2007)). Relative to this literature, the incremental contribution of the paper is to identify a cash flow component (mandatory contributions) that is clearly orthogonal to the investment opportunity set, once one controls for the pension plan funding status.[6]

Overall, the main contribution of the paper is to provide a unified empirical framework to separate the implications of under- and overinvestment theories. With respect to this goal, the paper to which the present work is closest in spirit is the one by Blanchard, Lopez-de-Silanes, and Shleifer (1994). These authors look at the investment behavior of a small sample of firms that receive cash windfalls. Instead, my large sample approach focuses on the market's valuation of the cash flows of companies that are mandated to pay pension contributions. In the same vein, the large sample study by Chaney, Sraer, and Thesmar (2007) finds results that are consistent with the evidence in this paper. Using variation in real estate prices as an exogenous measure of changes in collateral value, they show that investment is affected by net worth. However, the relaxation of financial constraints is good news only for firms with strong shareholder protection, as measured by the Gompers, Ishii, and Metrick's (2003) index.

The paper is organized as follows. Section I explains in detail the empirical strategy. Section II briefly describes the data leaving variable definitions to the Appendix. Section III applies the

[5]A notable exception is the paper by Officer (2007), who uses a change in the tax rate for dividends as a natural experiment to study the market response to dividend initiations.

[6]It is then no longer necessary to resort to the estimation of excess cash (cash in excess to the normal cash holdings), which may fail to effectively control for investment opportunities.

empirical strategy to identify price reactions to mandatory contributions for firms with different financial constraints status. The same apparatus is used in Section IV to estimate price reactions along the corporate governance dimension. Section V determines the aggregate reaction to financial constraints and agency conflicts and ranks firms on the spectrum from under- to overinvestment. The same section contains descriptive statistics on the firms in this ranking. Finally, Section VI draws the conclusions of this work.

I. Empirical Strategy

A. Theoretical Framework

The first goal of this paper is to measure the impact of financial constraints on investment by looking at price reactions following a cash flow shock. The theoretical prediction that motivates the empirical analysis is the simple statement that internal resources are more valuable in financially constrained firms than in unconstrained firms. The premium in the shadow value of internal finance in constrained firms reflects the additional cost of external finance and the NPV of the investment projects that could not otherwise be undertaken.

While this statement is self-evident in the comparison between a firm that is underinvesting and a frictionless firm, one can argue that it is true also when comparing firms at different levels of financial constraints. The appendix provides a formal proof in the context of Kaplan and Zingales's (1997) one-period model of investment. The basic intuition for the results in the appendix is that, given decreasing marginal productivity of capital, a positive shift in internal resources produces a larger increase in value for more constrained firms that are farther below the optimal investment level. Although not proved here, it seems reasonable to believe that the intuition extends to dynamic models in which firms engage in precautionary savings to avoid future financing constraints (Almeida, Campello, and Weisbach (2004), Whited and Wu (2006), Hennessy and Whited (2007)).[7]

This discussion highlights a major benefit of analyzing price reactions to assess the importance of financial constraints. This approach circumvents the controversy that surrounds the study of investment-cash flow sensitivities. Specifically, Kaplan and Zingales (1997 and 2000) argue that the practice initiated by Fazzari, Hubbard, and Petersen (1988) of sorting firms according to *a*

[7]Indeed, in a typical dynamic model, more constrained firms have lower investment-cash flow sensitivities, because they face higher costs of raising finance. This fact reveals a higher shadow price of internal funds in firms that are more constrained. I am grateful to Toni Whited for suggesting this intuition.

priori measures of financing constraints and comparing investment-cash flow sensitivities is not theoretically founded (see also Fazzari, Hubbard, and Petersen (2000) for a reply).[8]

There is an even more important advantage of studying price reactions to cash flow shocks. Price reactions provide the market's comprehensive assessment of the value of investment opportunities that are forsaken because of the cash flow shock. In this sense, all types of investment are accounted for, including the ones that are not otherwise measurable (e.g., investment in human resources and other intangible assets). Moreover, the drop in value accounts for the costs of external finance, which have been the subject of extensive studies (see, for example, Hennessy and Whited (2007)). In this sense, the study of price reactions can point out whether these distortions are economically significant from the market's point of view.

The second and related goal of this paper is to find evidence of overinvestment. In the context of the overinvestment studies, it is typical to focus on the market response to capital expenditures, acquisitions, or dividend initiations. The main challenge in this literature is that one cannot directly measure the NPV of the marginal investment project. Hence, one has to rely on the market's reaction to infer the value of managers' actions for shareholders. These studies interpret a lower post event return in firms that are *a priori* more exposed to a conflict of interest between managers and shareholders as evidence of overinvestment.

This paper focuses on price reactions following a shock to cash flows. In a firm that overinvests, the theoretical prediction on price reactions is specular to the case of underinvestment. Specifically, one can compare a firm where managers maximize shareholders' value and no other frictions are present with a firm where managers pursue their own interest. In the first case, the market price of internal resources is their face value. In the second case, the face value is reduced by the negative NPV of the wasteful projects that empire-building managers undertake. Hence, the same cash flow shock should induce an attenuated price reaction in the firm that overinvests.[9]

[8]Kaplan and Zingales's argument is corroborated by the results of the dynamic models in Moyen (2004) and Hennessy and Whited (2007).

[9]This statement is trivially true when comparing a firm that overinvests the funds that originate from a positive cash flow shock with a firm that does not overinvests them. If one wanted to extend the comparison to two firms that are differently prone to overinvestment (for example in the context of the one-period model in the appendix), the statement would be unambiguously true only in some circumstances. In particular, one would have to compare two firms that overinvest only their free cash flows, as suggested by Jensen (1986). In fact, assuming that managers overinvest costly external resources generates an effect that goes in the other direction, as a positive shock to internal resources partly increases shareholders' value by reducing the cost of external finance. A further sufficient condition is that the managers in one of the two firms have already achieved their desired level of overinvestment, while the

This paper aims at complementing the results in the previous studies by developing a research design that allows to control for investment opportunities. The potential coexistence of high agency conflicts and poor investment opportunities can be responsible for the low marginal value of funds in firms that are *a priori* more likely to overinvest. Then, to identify the effect of interest one wants to compare firms with the same marginal profitability. This issue is discussed next.

B. Identification and Empirical Specification

The empirical studies of the effect of capital market imperfections on investment are confronted with the challenge of identifying exogenous variations in net worth. If cash flow is correlated with future profitability, a link between cash flow and investment also results from a frictionless model in which investment is a function of expected profitability. In other words, changes in cash flows are an endogenous variable in the investment equation. In principle, controlling for investment opportunities with Tobin's Q should solve this problem. In practice, however, measurement error in Q is likely to invalidate this strategy (Erickson and Whited (2000)). Hence, the researcher needs to find variation in cash flows that is independent of investment opportunities.

This problem is transferred to a strategy that, like this paper, tries to measure the projection onto market value of an expected change in investment due to a shock to net worth. As argued, the price reaction following a drop in cash flow is the result of two effects. First, there is the direct effect of the change in cash onto company value. In this case, there should be a dollar decrease in value for each dollar drop in cash. Second, there is the effect related to investment. For firms facing costly external finance, this component includes the extra-cost of funding the projects that are still going to be undertaken, but which cannot be financed using internal resources. Also part of the investment effect is the present value of the projects that have to be given up as a result of the cash flow drop. The sign of this component depends on whether the firm is underinvesting (that is, the marginal project has positive NPV) or overinvesting (that is, the marginal project has negative NPV). In a frictionless world, only the first effect is present. What makes identification of the second effect more complicated is that it is magnified by the firm's investment opportunities. Specifically, the same drop in net worth is more costly for firms with good investment prospects. If the goal is identifying the second effect, one has to compare firms with the same marginal profitability. In other words, one needs to control for investment opportunities.

managers in the other firm still have incentives to undertake negative NPV projects. The details are available upon request.

This paper adopts the original identification strategy in Rauh (2006). The author uses mandatory contributions to a firm's defined benefit pension plans as a source of cash flow variation. As these contributions depend on the funding status, which is an endogenous variable that can be correlated with investment opportunities, he identifies the effect of interest by controlling for linear and nonlinear functions of the plans' funding status. For the details on the regulation that governs defined benefit pension plans, I refer the reader to Rauh's article. For the present purposes, suffices it to say that mandatory contributions for underfunded plans, i.e. plans for which the liabilities exceed the assets set aside, are determined according to a nonlinear function of the current funding status. Hence, along the lines of the regression discontinuity approach in labor economics (see, e.g., Angrist and Lavy (1999)) the identifying assumption in Rauh's approach is that the function that relates the pension funding status to investment opportunities does not have the same exact kinks, jumps, and asymmetries as the function that relates the pension funding status to mandatory contributions. This assumption is easily satisfied, as the rules determining mandatory contributions follow an arbitrary structure. For example, contributions are zero for plans that are fully funded or overfunded, while there is no reason of discontinuity at the level of full funding in the relationship between funding status and investment opportunities. This example is not exclusive, because the relationship between contributions and funding status displays kinks at different levels of funding (see Figure 2 in Rauh (2006)). Consequently, the identification comes from different points in the range of funding status, and not just from the comparison between zero and nonzero mandatory contributions at the full funding level, as one might suspect.

I use Rauh's (2006) approach to identify the component of the price reaction to a cash flow shock that is orthogonal to information about investment opportunities. In particular, the typical specification regresses the stock return ($R_{i,t+1}$) on the amount of mandatory contributions in the previous period (MC_{it}), an index of either financial constraints or agency conflicts (I_{it}), the interaction between mandatory contributions and the index I_{it}, linear and nonlinear functions of the funding status (FS_{it}), and a set of control variables (X_{it}):

$$R_{i,t+1} = \beta_0 + \beta_1 MC_{it} + \beta_2 MC_{it} I_{it} + \beta_3 I_{it} + \beta_4 FS_{it} + \beta_5 X_{it} + \varepsilon_{it}. \quad (1)$$

Under the assumptions of the regression discontinuity approach, controlling for the funding status identifies mandatory contributions as a cash flow decrease that is exogenous with respect to investment opportunities. Therefore, the coefficient β_2 captures the incremental effect of either financial constraints or agency conflicts on the price reaction to a cash drain, keeping investment opportunities constant. If I_{it} is a financial constraints index, and investment is sensitive to internal finance, then β_2 is predicted to be significantly negative. On the other hand, if I_{it} measures the extent of agency conflicts, a positive β_2 supports an empire-building story where part of the lost cash flows are wasted in negative NPV projects.

Some papers identify a positive cross-sectional link between expected returns and measures of financial constraints. For example, Whited and Wu (2006) find convincing evidence of a financial constraints factor. Moreover, following Gompers, Ishii, and Metrick's (2003) seminal paper, a number of studies point out a negative cross-sectional relationship between returns and indices of corporate governance (Bebchuck, Cohen, and Ferrell (2004), Bebchuck and Cohen (2005), Cremers and Nair (2005)). To control for this type of asset pricing effects, which could interfere with the identification of the effect of interest, the index I_{it} also appears directly in equation 1.

The vector X_{it} includes Tobin's Q at the beginning of the period as a further control for investment opportunities. In most specifications, X_{it} includes time effects to control for aggregate phenomena that could affect both returns and mandatory contributions, such as a market downturn, and to partly absorb the cross-sectional correlation in the errors term. To a similar purpose, the preferred specification also includes industry effects. Instead, while firm effects are included in some specifications, they are not part of the main specification, as the desired identification comes from cross-sectional variation in I_{it}. Also, the inclusion of firm fixed effects entails a look-ahead bias that prevents the interpretation of the regression results as returns from a feasible trading strategy.

To account for known determinants of returns, X_{it} also includes the stock beta, market capitalization, book-to-market (Fama and French (1992 and 1993)), past-twelve-months returns (Jegadeesh and Titman (1993)), and accruals (Sloan (1996)). Assuming that these variables control for all systematic sources of variation in asset prices, the estimated return reaction to mandatory contributions can be interpreted as the component of returns that is attributable to the shock to net worth.[10]

As the regressions are run on longitudinal data, one needs to be careful about the structure of the error terms in equation 1. Heteroskedestacity is certainly a concern given that returns on different stocks have different volatilities. Moreover, in spite of the inclusion of the cross-sectional determinants of asset prices, time effects, and industry effects, the residuals could still be cross-sectionally correlated. Finally, although the regression controls for price momentum, one may still be wary of time-series correlation in the residuals. These considerations suggest that the

[10]This conclusion also holds under the milder assumption that any omitted source of variation in asset prices is uncorrelated with mandatory contributions.

OLS assumptions are unlikely to be satisfied by ε_{it} and the standard errors need to account for that. Historically, the solution to cross-sectional correlation has been provided by the Fama and MacBeth (1973) procedure. More recently, clustering has been used as a remedy to either time-series or cross-sectional correlation, as it allows the researcher to fully exploit the time-series and cross-sectional dimensions of the data in a pooled regression. However, as pointed out by Thompson (2006), if one clusters by firm one must assume independence across time and if one clusters by time one must assume independence across firms. Neither of these assumptions is granted in the current context. Fortunately, this author proposes a simple solution that corrects standard errors for both cross-sectional and time-series correlation. The resulting standard errors are also robust to heteroskedasticity. The Monte Carlo experiments in that paper suggest that, even at sample sizes that are comparable to the one in the present article, the corrected standard errors perform better than the OLS ones, or the ones that cluster only by firm or time. Hence, I compute the standard errors throughout this work using Thompson's adjustment.

C. Mandatory Contributions as Cash Flow Shocks

In the discussion so far, I have interpreted mandatory contributions as a cash flow shock, i.e. an unanticipated drop in internal resources. This interpretation is inspired by the empirical results in Franzoni and Marin (2006). These authors show that using the year-end level of underfunding of the pension plan one can predict future negative abnormal returns. As an explanation, they argue that investors fail to fully anticipate the contributions to the pension plan that are predictable on the basis of the firm's funding history. Consequently, investors are negatively surprised when these cash transfers occur. These results, and additional evidence that I present below, suggest that mandatory contributions represent a shock to cash flows that triggers a price reaction.

Another piece of evidence that supports the interpretation of required contributions as a cash flow shock is Rauh's (2006) decomposition of contributions into the expected and unexpected part using assumptions on the evolution of pension liabilities and assets. The author shows that the two components have impacts on investment that are statistically indistinguishable from each other, suggesting that even company managers fail to forecast future contributions.

In spite of the evidence in these two papers, there could still be a fraction of mandatory contributions that is anticipated by the market. This fact can be problematic for the strategy in the current paper to the extent that the degree of anticipation is correlated with the measures of financial frictions. Specifically, I_{it} in equation 1 could act as a proxy for the component of mandatory contributions that is unanticipated and that causes the price reaction. For example, one may argue that it is harder for the market to predict future mandatory contributions for small firms on which there is less information. At the same time, small firms are arguably more financially constrained. The paper addresses this concern by including interactions of mandatory contributions with variables that proxy for information about the company, such as the size and the number of analysts following the firm.

It is useful to be more explicit about the timing of some of the variables in equation 1. Total contributions to the pension plan, that is the sum of mandatory and voluntary contributions that the firms makes to all its pension plans in a given year, are reported in the forms 5500 that the company files with IRS. While these forms may be filed with some delay and are slow to become public information, the firm also reports total pension contributions in the statement of cash flows and provides a discussion of its pension obligations in the footnotes to the financial statements.[11] This fact allows one to consider total contributions as public information at the end of the fiscal year. Consequently, the empirical analysis is structured as a replicable trading strategy. To allow for the diffusion of information in the market, and following a standard practice (see, e.g., Fama and French (1993)), mandatory contributions for the fiscal period ending in year $t-1$ are associated with returns between July of year t and June of year $t+1$. In the main specification, the return frequency is annual, as one would like to measure the total price reaction over the year following the release of the public information about pension contributions. The choice of a one-year window is motivated by the evidence in Franzoni and Marin (2006) that the return reaction to pension information is long lasting. The authors argue that abnormal return reactions to pension plan underfunding can be predicted for at least five years after portfolio formation, suggesting that the market is slow in impounding this information into prices. In the current paper, the choice of a one-year window after the contributions seems a reasonable compromise between the needs of capturing investors' reaction to the information and having enough independent observations. As a robustness check to the choice of the frequency, the paper also reports results where monthly returns between July of year t and June of year $t+1$ are regressed on contributions from year $t-1$.

It is worth stressing that the analysis could not be set up as a short-run event study where one measures the returns around the announcement dates of pension contributions. First, the exact dates in which pension contributions become public information are not known. Second, as said

[11] The recent reform of pension accounting contained in SFAS 158 causes the pension information to appear more explicitly in the balance sheet. However, this reform does not concern the sample years covered in this paper.

above, the evidence in Franzoni and Marin (2006) indicates that the market takes a long time to react to pension information. Hence, the choice of a long post-event window is the natural one.

II. The Data

The data for this paper come from the intersection of different sources.

Mandatory contributions to defined benefit pension plans are constructed by Rauh (2006) from Form 5500 filings.[12] These data are only available for fiscal years from 1990 through 1998. A necessary but not sufficient condition for a firm to incur mandatory contribution is that one of its pension plans be underfunded. The formula to compute mandatory contributions is fairly complex and changes over the period of interest. Broadly speaking, contributions depend on the funding history and the pension cost for the current year on a plan-by-plan basis (details in Rauh (2006)).[13] Rauh's data contain mandatory contributions at the firm level. As reported by the author, approximately one-quarter of Compustat firms had defined benefit pension plans in the nineties, but these firms account for more than half of Compustat firm book value. Out of these observations, one selects only those that have valid Form 5500 information. The final data set provided by the author contains 8,030 firm-year observations on 1,522 firms. Mandatory contributions are divided by total assets in the previous year to obtain the variable MC, which is finally winsorized at the first and ninety-ninth percentiles. Also available in Rauh's data is the aggregate funding status of all defined benefit plans sponsored by a firm (FS). It is computed by subtracting the aggregate pension liabilities from the aggregate pension assets at the firm level and normalizing by total assets in the previous year.

Contributions are matched by firm and year to the accounting information from Compustat.

[12] I am grateful to Joshua Rauh for providing his data.

[13] Specifically, contributions depend on the funding status of the domestic Defined Benefit plans, on a plan by plan basis. The plan funding status results from comparing the plan liabilities (present value of future benefits to which the workers are entitled) with the plan assets (the market value of the assets that have been set aside to provide for the pensions). The plans that are more than fully funded do not incur mandatory contributions. For the plans where assets are below liabilities, the required pension contribution is the maximum of two components: the minimum funding contribution (MFC) and the deficit reduction contribution (DRC). The DRC as a percentage of firm funding is given by $min\,\{0.30, [0.30 - 0.25 * (\text{funding status} - 0.35)]\}$ for 1987 to 1994 and $min\,\{0.30, [0.30 - 0.40 * (\text{funding status} - 0.60)]\}$ for 1995 and later. The MFC is defined as the "normal cost" plus 10% of underfunding. The "normal cost" roughly corresponds to the present value of pension benefits accrued during the year.

In the main analysis, all matching observations are used, irrespectively of the industry, to avoid further reducing the size of the sample. Robustness checks are provided that restrict the sample to manufacturing industries. Given that MC is divided by assets in the previous period, a firm needs to have available accounting data for at least two years prior to the year t before appearing in the sample in year t.

Compustat reports a fiscal year as ending in the previous calendar year if the fiscal-year-end month is before June. When matching contributions and accounting information to market information from CRSP, I undo this transformation and use the actual fiscal year end. Consequently, for some firms the available information stretches to 1999. Then, contributions and accounting variables for fiscal years ending in year $t-1$ are matched with annual and monthly returns between July of year t and June of year $t+1$. Annual returns are computed by compounding monthly returns from CRSP. In the end, returns range between July of 1991 and June of 2001.

Using Compustat and CRSP data, I construct Tobin's Q (definition in the appendix), beta, size, book-to-market (BM), momentum returns, accruals, and industry dummies. Beta in a given year is computed in December using monthly returns on at least twenty-four and at most sixty months of prior data. The market index is the CRSP value-weighted portfolio. Size is market capitalization. BM is book value of equity at the end of the year, as in Fama and French (1993), divided by market capitalization in December (details in the appendix). Momentum is captured by the compound return over the twelve months before year t. Accruals are computed as in Sloan (1996) (the definition is in the appendix). All of these variables are winsorized at the first and ninety-ninth percentiles. Industry dummies are created using three-digit SIC codes. These variables are constructed at the end of year $t-1$ and matched to returns between July of year t and June of year $t+1$, except for size and momentum that are computed in June of year t.

The analysis uses three indices of financial constraints. The appendix provides the details on their construction. The first index is the one conceived by Whited and Wu (2006), which I label WW index. These authors fit the shadow value of external funds from a structural model of investment onto a number of observable variables. Being based on the estimation of the Euler equation, the resulting index is free from the endogeneity problem that I described above. The second measure of financial constraints is the Kaplan and Zingales index (KZ index), as developed by Lamont, Polk, and Saa-Requejo (2001). This score is obtained by running an ordered logit of Kaplan and Zingales's (1997) scale of financing constraints onto observable characteristics. Although the estimation is originally performed on a sample of forty-nine firms, the literature has extensively applied this measure to classify firms in large samples. Thirdly, I use the Cleary (1999) index in the version by Hennessy and Whited (2007). Cleary uses discriminant analysis based on observable company characteristics to construct a Z-score for the firm's likelihood of increasing or decreasing dividend payments. In synthesis, the KZ index identifies as constrained those firms with low cash stock, low cash flow, and high leverage. Firms ranking high with the WW index are small, rely heavily on equity financing, have low cash flow, and are slow growing firms in fast-growing industries. Firms appearing constrained by the Cleary index are slow-growing, have low profit margins, and have few resources to cover their debt burdens. The three financial constraints indices are constructed using accounting information from the end of year $t-1$ and are matched to returns between July of year t and June of year $t+1$. The indices are defined so that higher levels denote firms that are *a priori* more financially constrained. To be included in the sample a firm needs to have sufficient data to construct the three measures.[14]

The purpose of adopting three indices is to provide evidence that is robust to alternative measures of financial frictions. Hennessy and Whited (2007) distinguish two dimensions of financial constraints. The first dimension is the firm's need for external funds, as measured by the ratio of first best investment to internal resources. The second dimension is the cost of external funds, which is the additional cost the firm would incur if it used external rather internal finance. According to these authors, the three indices capture the need for external funds. However, the results in the paper are also robust to other classifications schemes based on size and dividend payments, which according to Hennessy and Whited (2007) are better proxies for the cost of external funds.[15]

To capture potential agency conflicts between managers and shareholders, I use the corporate governance index (G) by Gompers, Ishii, and Metrick (2003).[16] These authors use the data from the Investor Responsibility Research Center (IRRC) providing twenty-four distinct corporate-governance provisions on about 1,500 firms since 1990. This measure is meant to capture the balance of power between managers and shareholders in a corporation. The index adds one for each anti-takeover provision adopted by the firm. Hence, higher levels of the index imply less shareholder protection. Returns between July of year t and June of year $t+1$ are matched with G from the latest publication of the Investor Responsibility Research Center as of year t. In order not to excessively reduce the sample size, the availability of the G index is a selection criterion only for

[14] To preserve the size of the sample, I do not condition on positive sales growth when constructing the financial constraints indices.

[15] These results are not reported, but are available upon request.

[16] The index is available on Professor Metrick's website.

the part of the paper that focuses on the price impact of agency conflicts (Sections IV and V). As robustness checks, I consider two other governance indices by Bebchuck, Cohen, and Ferrell (2004) and Bebchuck and Cohen (2005). The first measure is called Entrenchment index and is based on six out of the original twenty-four provisions. These are four constitutional provisions that prevent a majority of shareholders from having their way (staggered boards, limits to shareholder bylaw amendments, supermajority requirements for mergers, and supermajority requirements for charter amendments), and two readiness provisions that boards put in place to prevent hostile takeovers (poison pills and golden parachutes). The second measure focuses exclusively on the presence of staggered boards.

Excluding the selection criterion based on G, the sample consists of 6,119 firm-year observations on 1,198 firms. After imposing the availability of G, the sample is restricted to 4,107 firm-year observations on 834 firms. Table I reports summary statistics on the main variables in the analysis. Concerning MC, the average contribution in the sample is 0.1% of total assets. This number includes the observations for which contributions are zero (about 71%). Conditioning on nonzero observations, the average MC is about 0.24%. Except for the correlation between the KZ and Cleary indices (59%), the correlations among the financial constraints indices tend to be low, suggesting that the three classification schemes capture different financial frictions. The WW index is strongly negatively correlated with size (-63%), likely because it depends on total assets with a minus sign (see the appendix for its definition).

One question that may arise is whether the firms that incur mandatory contributions are different under some dimension from the rest of the sample. Specifically, one may wonder whether they qualify as distressed companies. If that was the case, the empirical results that follow could not be generalized to the universe of listed firms and the interpretation of the price reactions would take a different connotation, specifically in terms of costs of financial distress. Table II reports descriptive statistics on a number of characteristics for firms that pay and do not pay contributions. Also, among the firms with positive MC, I consider those with contributions above and below the conditional median of MC (that is, about 0.09%). It appears that the top contribution firms rank as most constrained according to the three FC indices. Instead, there is no apparent difference in the governance index across MC groups. There is a hump-shaped relationship between MC and firm size, with the smallest firms in the top MC group. A similar pattern concerns firm age and the coverage by analysts and rating agencies. The relationship between MC and BM is instead U-shaped, with the top MC firms displaying the highest BM. While it is the case that high MC

firms have the lowest cash flows, highest leverage, and are least inclined to pay dividends, the difference with respect to the other groups is not substantial and does not suggest a condition of financial distress. Relative distress of high contribution firms does not appear likely also in the light of their sales growth, which attains the same level as for zero contributions firms. If top MC firms rank lowest in terms of the operating margin (IMARG), the evidence on interest coverage (COVER) and assets liquidity (SLACK) does not suggest financial distress for these companies. In summary, while high MC firms appear to be more constrained and smaller, there is no evidence that they are financially distressed. In the empirical tests that follow, this heterogeneity in size is carefully controlled for.

III. Financial Constraints

A. Price Reactions To Mandatory Contributions

As argued in Section I, the assumption underlying the analysis in this paper is that at least a part of mandatory contributions represents a cash flow shock. The goal is measuring the differential impact of financial frictions on firm value by estimating the price reaction to mandatory contributions across different levels of financial constraints and agency conflicts.

The idea of looking at the price reaction over the year following the contributions to the pension plan is suggested by the evidence in Franzoni and Marin (2006). However, these authors use the degree of underfunding in the pension plan to predict future returns. They do not use pension contributions, as they are not available in Compustat. Hence, the conclusion in their paper that the market is surprised by mandatory contributions is only indirect.[17] Fortunately, the availability of mandatory contributions in Rauh's (2006) data allows one to directly verify that contributions significantly predict returns.

Table III reports the estimates from a shorter version of equation 1, in which the two variables involving I_{it} do not appear. The regression is estimated on pooled data. To ease the interpretation, the variable MC has been standardized. I control for investment opportunities through the funding status (FS) and Tobin's Q, in order to identify the price reaction associated with the part of the cash flow shock that is orthogonal to information about the firm's future investment prospects. The different columns in the table correspond to different sets of controls. All models include time effects. In column 1, there is a significant impact of mandatory contributions on returns over the

[17] Their story is that the pension deficit in a given year triggers mandatory contributions in the future, which hit investors by surprise. They argue that the recorded abnormal returns are the evidence of this surprise.

next year. However, this specification does not include controls for asset pricing effects. In columns 2 through 4, where I add controls for the stock beta, size, BM, momentum, accruals, and industry effects, mandatory contributions remain a significant predictor of returns. In column 5, I allow for potentially time-varying risk premia by interacting the asset pricing controls with time effects. The slope on MC stays significant. In this case, a one-standard deviation increase in contributions (about 0.2% of total assets) reduces annual returns by about 1.9%. By including firm fixed effects in column 6, the identification comes from firm level variation of mandatory contributions. In this case, an increase in contributions at the firm level causes a significant drop in returns relative to the average return earned by the firm over the sample. In all specifications, the controls for investment opportunities (FS and Q) are not significant. While not reported, the slopes on the asset pricing controls are consistent with previous studies.

Overall, Table III suggests that mandatory contributions significantly predict returns over the following year. These results confirm Franzoni and Marin's (2006) evidence of cross-sectional return predictability and slow reaction to pension information. Moreover, they legitimate the use of mandatory contributions as a cash flow shift that is partly unpredicted by the market, namely a source of cash flow shocks.

B. Price Reactions and Financial Constraints

Next, the analysis directly addresses the issue of whether the market's response to mandatory contributions varies across firms displaying different degrees of financial constraints. As mentioned above, the purpose is to assess to what extent investors impound into prices the fact that a shift in internal funds obliges financially constrained firms to face costly external finance and to give up positive NPV investment projects.

In the specification from equation 1, the interaction between the index I_{it} and mandatory contributions is replaced by the product between MC and dummy variables denoting firms in different financial constraints groups. In particular, for each of the three indices of financial constraints (FC indices), the group of mostly constrained firms contains those companies that in year $t-1$ rank in the top 25% of the distribution of the index (Hi FC). The medium group contains firms between the 25th and 75th percentiles (Med FC). Hence, the coefficients on these two interactions measure the incremental price reaction to a given amount of contributions for firms in one of the two groups relative to those in the bottom quarter of the distribution, for which the price reaction is captured by the slope on MC. The financial constraints index also appears directly in the regression.

Table IV reports the estimated coefficients. All models include time effects. The variable MC is standardized, so that the coefficient on the interactions captures the incremental price reaction of a one-standard deviation shift in contributions for different groups of constrained firms. The FC indices are standardized too.

In the first specification the asset pricing controls are beta, BM, and size. For the three FC indices, constrained firms (Hi FC) display significantly larger price reactions (in absolute value) than unconstrained firms. For example, a one-standard deviation increase in contributions causes an annual price decline that is significantly larger by 2.65% for constrained firms than for unconstrained firms according to the KZ index. Strikingly, the magnitude of the effects for the top group is very similar across FC indices, suggesting that the three measures are likely to capture the same phenomenon.[18]

Incidentally, only the WW index has a significantly positive direct effect on returns. Specifically, the returns of firms that are one standard deviation apart in terms of the WW differ by 4.44% annually. This evidence is consistent with the results in Whited and Wu (2006) who find a significant premium on a financial constraints factor that is constructed using their score. Moreover, the lack of a significant direct effect of the KZ index on returns reflects the evidence in Lamont, Polk, and Saa-Requejo (2001), who fail to find a systematic difference in returns between constrained and unconstrained firms as classified by the KZ index. The mixed evidence on the existence of a financial constraints factor that is present in other papers, and confirmed in Table IV, does not interfere with the conclusions of this work. Here, the focus is on return reactions after a cash flow shock that is idiosyncratic in nature. Hence, these shocks are not candidates for the type of systematic variation that generates a factor in returns.

The models 2 through 4 include larger sets of controls. In all cases, the main inference does not change: price reactions are significantly larger for constrained firms according to the three indices. In particular, specification 4 allows the slopes on the asset pricing controls (Beta, BM, Size, momentum, accruals, and industry effects) to change in each year. In this case, the incremental effect of being in the group of constrained firms is even more significant than for previous models. For example, a one-standard deviation increase in contributions causes an additional negative annual return of about 3% for constrained firms according to the KZ index.

Finally, by including firm fixed effects, model 5 achieves identification through firm specific

[18] For completeness, the difference in slopes between the top and medium groups is significant only for the KZ index.

variation in contributions and financially constrained status. In this case, the constrained group displays a significant incremental return reaction only with the WW index. The lack of significant effects for the other indices is probably the result of insufficient firm level variation of the degree of financial constraints.

The overall evidence from Table IV suggests that the market differentiates between constrained and unconstrained firms when reacting to a given drop in net worth. In the case of constrained firms, the price reaction is significantly more negative. These results are insensitive to the definition of the FC index that I use. Although not reported, similar evidence is obtained when the constrained status is defined on the basis of firm size and dividend payments.

The fact that the heterogeneity in investment opportunities is absorbed through the inclusion of the funding status assures that the estimated effect does not result from spurious correlation between the FC indices and marginal firm profitability. Instead, the preferred interpretation hinges on the incremental value that internal liquidity has for firms that face costly external finance. Along these lines, one can perform a back of the envelope calculation to compute the market's assessment of the costs of external finance and lost investment from a cash flow shock. Assuming that all the price reaction takes place within the year, in model 4 of Table IV, the effect of a one-standard deviation increase in MC (about 0.24% of total assets) reduces market value of constrained firms by 3% according to the KZ index. Given the average ratio of assets to market capitalization for constrained firms, this coefficient suggests that a drop in net worth equal to 1% of the firm's market value would trigger a negative return of about 3.6% for the average constrained firm.[19] According to this interpretation, the market's perception of the shadow price of internal funds for constrained firms is about 3.6 times the dollar value of the cash flow shock.

Consistent with this interpretation, one cannot reject the hypothesis that the effect of a one-standard deviation increase in contributions causes market value to fall by 0.24% for unconstrained firms (slope on MC). Given what said above, this effect represents a one-to-one price response to a cash flow drop. In other words, as one would expect, the cost of financial constraints for these companies is zero and the magnifying effect on price reaction is not present.

One might wonder whether the magnitude of the estimated return reaction to a drop in internal resources reflects realistic investment behavior. While an accurate investigation of this issue would require the calibration of a fully specified investment model, which is outside the scope of this

[19] The average ratio of assets to market capitalization for the KZ constrained group is about 3.46. Then, a cash flow drop of 0.24% of total assets represents about 0.83% of market value for the average firm. Hence, given the 3% estimate, the effect of a decrease in cash flow by 1% of market value is equal to (3/0.83)% = 3.6%.

paper, a simple back-of-the-envelope calculation provides some intuition. As argued, one possible reading of the results in Table IV is that a cash flow shock of 1% of market capitalization reduces market value by about 3.6 times that amount for constrained firms. Assuming for simplicity that the incremental decrease is due to forsaken investment opportunities, rather than additional cost of capital to existing projects, one can compute the internal rate of return (IRR) of these projects. In this case, the present value of the investments that would have been generated by an initial outlay equal to 1% of market capitalization is estimated at about 2.6% of current capitalization. By making plausible assumptions on the growth rate of the firm's cash flows and cost of capital, one can conclude that the IRR of the forsaken projects ranges between about 11.6% and about 20.8%. These numbers suggest that the estimated return reactions are consistent with realistic investment behavior.[20]

From the point of view of asset pricing, the empirical results confirm the theoretical presumption that financially constrained firms are more sensitive to shocks to their net worth than unconstrained firms. In this sense, the evidence provides a foundation for the empirical work that either points out a magnified sensitivity to aggregate factors for firms that rely on internal finance (e.g., Perez-Quiros and Timmermann (2000)) or searches a financial constraints factor (e.g., Lamont, Polk, and Saa-Requejo (2001), Gomes, Yaron, and Zhang (2006), Whited and Wu (2006)). While this paper cannot take a stance on the existence of a separate financial constraints factor, as it examines idiosyncratic shocks to net worth, the by-product of the analysis is the confirmation of the findings in Whited and Wu (2006).

C. Robustness Checks

The first of set of robustness checks adds further controls to the original specification, estimates the model on different samples, and re-defines the variables. The results are in Table V that reports the slopes on the FC index, mandatory contributions, and the interaction between MC and the top financial constraints group dummy for the three FC indices. As a benchmark, the first row contains the preferred specification from Table IV (model 4).

[20] One has to compute the internal rate of return of an investment of 1% of market capitalization, whose infinite stream of cash flows has a present value (PV) that equals 2.6% of capitalization. One can use the formula $PV = CF/(r-g)$, where r is the firm's cost of equity capital and g is the nominal growth rate of these cash flows (CF). Assuming a growth rate that equals the observed sales growth for constrained firms (7%), risk free interest rate in this sample of 5%, beta of 1 (which corresponds to observation), and equity premium between 3% and 5%, and using CAPM for the cost of capital, one concludes that the IRR of these projects ranges between 11.6% and 20.8%.

The discussion in Section I points out the need to control for investment opportunities. The rationale is that the same cash flow shock can impact value differently depending on a firm's investment prospects. In this sense, the total price reaction may also depend on the interaction between investment opportunities and marginal contributions. Omitting this variable can cause a bias in the estimation of the effect of interest.[21] Fortunately, one can avoid this problem by controlling for the interaction between MC and a proxy for investment opportunities, for which I use either FS or Tobin's Q. The estimates in rows 2 and 3 reveal that the inclusion of this potentially omitted variable does not affect the main results, which are even stronger in some specifications. Incidentally, although not reported, the slopes on the interactions between MC and FS are never significant. Instead, the coefficients on $MC \times Q$ are negative and significant, consistent with interpretation that better investment prospects magnify the price reaction to the drop in cash flows.

It could be the case that the same cash flow shock generates heterogeneous returns reactions because of a discount rate effect. In particular, the present value of the forsaken investment projects may differ, because the cash flows are discounted at different rates. Then, if the index was correlated with the discount rate, the interaction between the FC index and MC contributions would capture this effect. However, it is likely the case that more constrained firms are more risky and have higher discount rates, which seems to be confirmed by the summary statistics in Table I. In such circumstances, the discount rate effect would actually weaken the magnifying effect of financial constraints for price reactions. To verify whether this story plays any role, I estimate models that include the interaction of MC with beta, BM, momentum, and size, as proxies for discount rates. The results are in rows 4, 5, 6, and 7. The coefficient on the interaction between MC and the FC index is not significantly impacted by the inclusion of any of these controls. On a related note, if beta is measured with error, leverage may be a more likely candidate to capture the discount rate effect. When one includes the interaction between MC and leverage in the specification, the estimate of interest loses significance for the KZ and Cleary indices, but not for the WW index (untabulated results). The insignificant results for the KZ and Cleary indices are likely due to their collinearity with book leverage (correlations of 73% and 71%, respectively). Instead, the significant estimate for the WW index, whose correlation with leverage is negligible, tends to rule out a discount rate effect operating through leverage.

The empirical strategy assumes that, for all firms, the market's reaction to mandatory contribu-

[21] See, for example, Section 4.3.3 in Wooldridge (2002).

tions in a given year is captured by the stock return in the following year. This approach could be invalidated if the speed at which information is impounded into prices differs across firms in a way that is correlated with the FC indices. For example, one could imagine that the information about a given amount of contributions spreads less rapidly for smaller firms. At the same time, small firms are likely to appear as financially constrained. In this case, the interaction between MC and the High FC dummy would be capturing price informativeness rather than financial constraints. To partly control for this possibility, I introduce interactions of MC with variables that proxy for the speed of information revelation such as firm size and the average number of analyst following the company over the year. Rows 7 and 8 of Table V report the results. The coefficients of interest are in general not affected. Only in the case of the WW index in row 7, the slope on the incremental impact for the high FC group loses significance. This fact depends on the high correlation between size and the WW index (see Table I). A test for the joint significance of the size control and the interaction of interest rejects the null of zero coefficients. Overall, these results suggest that the identification is not affected by the controls for information, which in turn do not have a significant impact on returns (estimates not reported).

The regression discontinuity approach that identifies mandatory contributions as a cash flow shift orthogonal to investment opportunities requires that one controls for the exact function that links investment opportunities to the funding status. So far, I have only included a linear function of the firm's aggregate funding status (FS). Following Rauh (2006), the specification in row 9 adds separate controls for the funding status of the over- and underfunded plans sponsored by the firm and for the second and third powers of these variables. In spite of the high correlation of the included variables with MC, there is still sufficient independent variation in mandatory contributions to identify the effects of interest, which remain significant.

Most studies on the impact of financial constraints on investment restrict the analysis to firms in the manufacturing industries. Imposing this further selection criterion limits the sample to 3,861 firm-year observations. The estimates in row 10 suggest that the effect of interest stays significant even in this smaller sample (at the 10% level for the WW index and with higher confidence for the other indices). In this sample, the impact of mandatory contribution on constrained firms according to the KZ and Cleary indices is even larger (-4.58% and -3.51%, respectively), suggesting that financial constraints are more relevant for manufacturing firms.

The analysis by subsamples (rows 11 and 12) reveals that the effect of financial constraints is stronger in the first part of the sample. Still, when classifying firms by the WW index, one can identify a significant additional price impact of contribution on constrained firms also in the second subsample. Other results below confirm that the effect is present in the later years as well.

The fact that Tobin's Q enters the original definition of the KZ index makes it more suitable to capture firms in need of external funds rather than companies that face costly external finance. Hence, following Baker, Stein, and Wurgler (2003), I define a version of the KZ index that excludes Q (KZ4). The estimates in the last row of Table V suggest that using KZ4 the impact of contributions on constrained firms remains significant, although it is slightly smaller.

In estimating the main specification, the choice has fallen on the use of annual as opposed to monthly returns. One motivation is that temporary variations in stock prices are more likely to cancel out over the yearly horizon, so that one can better identify the price reaction of interest. Another reason is that the monthly price reactions are not necessarily equally spread over the twelve months and across firms, which implies further noise in estimating the monthly effects. This noise is more likely to be averaged out in annual returns. However, if the error terms are independent at the monthly frequency and the price reaction is sufficiently constant across months and firms, the larger sample allowed by monthly observations guarantees more efficiency and better approximation of the asymptotic limits. This is especially important for the convergence of the standard errors, for which the asymptotics are defined in both the time-series and cross-sectional dimension. Hence, it is worthwhile considering higher frequency results as well. Table VI replicates the structure of the previous table, but uses monthly returns between July of year t and June of year $t+1$, rather than annual returns over the same horizon. As a further robustness check, I replace the interaction between MC and the dummies for the two groups of financially constrained firms with an interaction between MC and the FC index (the specification is therefore the same as in equation 1).[22] Moreover, time fixed effects are now at the monthly frequency too.[23] The general result is that the effect of interest remains overall statistically significant. Compared to Table V, the effects are slightly weaker for the WW index, and somewhat stronger for the other two indices. Remarkably, for two measures out of three (the WW and Cleary indices) the interaction between MC and the FC index is at least marginally significant also in the later subsample. Overall, the

[22]In this case, the interpretation of the coefficient on the interaction is the annual change in returns due to a one-standard deviation increase in the FC index for firms experiencing a one-standard deviation increase in mandatory contributions.

[23]A further difference between the specifications in Tables V and VI is that the latter does not include the interactions between the asset pricing controls and time effects. Accounting for these interactions reduces the degrees of freedom and makes the results less strong, but still statistically significant in most cases.

results from monthly returns confirm the evidence at the annual frequency.

D. Portfolio Analysis

Franzoni and Marin (2006) adopt a portfolio methodology that uses the aggregate degree of underfunding of a firm's pension plans to predict abnormal returns. The measure of funding status that they use is the difference between assets and liabilities in the pension plan divided by market capitalization. As explained above, they interpret the negative abnormal returns for the portfolio of severely underfunded firms as evidence that the market is surprised by the drop in cash flows due to the pension contributions.

In the present context, the portfolio approach provides an entirely different strategy to further corroborate the results presented so far. The hypothesis that price reactions to cash flow shocks should be larger for more financially constrained firms suggests that the abnormal return to the portfolio of underfunded firms can be magnified by further sorting on the FC indices. In particular, in each portfolio of firms with poor funding status, the firms that rank higher by the FC index should display returns that are more negative.

The main advantage of this strategy is that it does not rely on the availability of mandatory contributions. Hence, the sample enlarges considerably and covers a much longer period. However, this approach presents a number of shortcomings relative to the one that uses mandatory contributions. First, one cannot directly control for the amount of the cash flow shock through mandatory contributions. Indeed, the sort by the funding status is only indirectly a sort by the extent of the cash flow shock, as the formula determining contributions depends on other variables besides the funding status.[24] Second, one loses the possibility of controlling for investment opportunities through the funding status, which now becomes the main source of identification of price reactions. As a consequence, it may be the case that the sort by the FC index identifies firms with different marginal profitability, which in turn could justify differential price reactions to a given cash flow shock. Finally, and most relevant, it becomes impossible to separately identify the direct effect of financial constraints on returns from the indirect effect of financial constraints on the price reaction to a given cash flow shock. This identification is performed in equation 1 by the separate inclusion of the index I_{it} and the interaction $MC_{it} \times I_{it}$. In the portfolio analysis, one can only capture the combination of the direct effect and the interaction. Hence, if for example there is a financial

[24] In particular, in this application the measure of funding is the difference between assets and liabilities in all the plans sponsored by the firm, while the true formula to compute contributions is based on plan level data.

constraints factor in returns (that is, β_3 is positive in equation 1), it becomes more difficult to identify the magnifying impact of financial frictions on the price reaction to a given cash flow shock (that is, a negative β_3 in equation 1).[25]

With these caveats in mind, I extend the analysis in Franzoni and Marin (2006) to capture the incremental predictive power of financial constraints. The sample includes all defined benefit firms for which data on assets and liabilities in the pension plans are available in Compustat. Also, to be part of the sample, a firm needs to have data to construct the three FC indices. These requirements are satisfied by firms in fiscal years between 1980 and 2004. As in that paper, the pension funding ratio for year t (FR) is defined as the difference between aggregate assets and liabilities in the pension plans at the end of the fiscal year divided by market capitalization in December (the details on variable definitions are in the cited paper). In July of each year t, five quintiles are formed according to the distribution of FR in year $t-1$, conditioning on the firms with aggregate underfunding ($FR < 0$), and using NYSE breakpoints. Then, the stocks in each FR quintile are further sorted into three groups using the distribution of the FC index in year $t-1$. The three groups are: Low FC (bottom 20%), Medium FC (between 20% and 80%), and High FC (top 80%). Value-weighted portfolios are formed with the stock in each of the resulting fifteen groups. Monthly returns on these portfolios range between July 1981 and December 2005.

The results in Franzoni and Marin's paper suggest that underfunded firms are mostly mispriced relative to the Fama and French's (1993) three-factor model. The explanation is that underfunded companies tend to be value firms and, as such, earn a positive premium, which needs to be subtracted from their returns to identify the pension related overpricing. For this reason, and for the sake of brevity, I focus on the intercept (alpha) from the three-factor model.

As a reference, the first column of Table VII reports the alphas for the five portfolios that result from a single sort by FR. As expected, the most underfunded portfolio displays a significantly negative abnormal return of -0.66% monthly. No significant mispricing is detected for less severely

[25] In fact, one could think of identifying the interactive effect using a difference-in-differences approach. The interaction could be estimated from the difference in excess returns between two zero-investment portfolios. The first one is the portfolio that goes long in the High FC portfolio and short in the Low FC portfolio conditioning on the most underfunded stocks. The second portfolio results from the same difference restricted to the least underfunded stocks. Under some assumptions, the double difference filters out the direct effects of both underfunding and financial constraints leaving only the interaction. The crucial assumption, which is unlikely to be satisfied, is that the direct effect of financial constraints is the same across levels of underfunding. This stringent requirement makes this approach not worthwhile.

underfunded firms.

The rest of Table VII focuses on the sets of portfolios that result from the double sort by FR and one of the three FC indices. At each funding level, the table also provides the abnormal return of an arbitrage portfolio that is long in the Low FC and short in the High FC stocks. If financial constraints have an incremental impact on price reactions, one should expect these long-short portfolios to earn positive and significant alphas. This is especially the case for more underfunded firms that are required to make larger contributions.

The evidence from the WW index is mixed. At the highest level of underfunding, the largest alphas in absolute value are for the relatively unconstrained companies. Instead, in the second FR quintile, constrained firms display the most negative abnormal returns and the alpha for the arbitrage portfolio is marginally significant at 0.61% monthly. The results for the remaining portfolios are also mixed, with the long-short positions earning positive and insignificant alphas for FR quintiles three and four, and a significantly negative alpha for the fifth FR quintile.

These results are not surprising in the light of the previous discussion. The portfolio approach does not control for the direct effect of financial constraints on returns. So, the financial constraints factor identified by Whited and Wu (2006) is likely to play a role here. The higher returns of constrained firms in the first and fifth FR quintiles are probably the result of their positive loadings on the Whited and Wu factor. In spite of this effect, three arbitrage portfolios out of five earn positive alphas, providing some support for the hypothesis that financial constraints magnify the price reactions to cash flow shocks.

More convincing evidence comes from the KZ index. In this case, at all levels of underfunding, the spread portfolios earn positive returns. Remarkably, the alphas of these long-short positions are significant in four cases out of five. For example, in the quintile of the most underfunded firms, the differential return between unconstrained and constrained stocks is a striking 1.63% monthly. The previous argument on financial constraints factors in returns can probably explain why the KZ index better identifies the incrementally negative price reactions for constrained firms. Specifically, the lack of a significant risk premium associated with the KZ index in Lamont, Polk, and Saa-Requejo (2001) suggests that one should not expect firms that rank higher by this measure to earn higher returns.

Finally, the results for the Cleary index are also supportive of the magnifying role played by financial constraints for price reactions. Again, all long-short portfolios display positive abnormal returns. In the first two FR quintiles, the alphas are economically large and statistically significant.

Overall, the evidence in Table VII corroborates the conclusions from the previous analysis using a longer sample and a different methodology. Financially constrained firms tend to display significantly larger price reactions for a given level of pension underfunding, which in turn is related to contributions to the pension plan. Still, in interpreting these results, one has to keep in mind that the portfolio approach does not control for investment opportunities, which are potentially correlated with the financial constraints indices.

IV. Evidence of Overinvestment

As pointed out by Stein (2003), the evidence that investment is sensitive to cash flows is not by itself enough to discriminate between costly-external-finance and empire-building models. A theory nesting the two stories is likely to be a more appropriate description of reality. If that is the case, both under- and overinvestment are possible outcomes. It is then an empirical question to assess which situation is relevant in practice.

The previous section concluded that the market reacts more strongly to cash flow shocks that hit more financially constrained firms. These results are consistent with an underinvestment story, where investors anticipate that internal finance is more valuable for constrained companies. In what follows, I use the same framework to tackle the issue of overinvestment. The assumption is that the stronger the agency conflict, the higher is the managers' incentive to undertake wasteful projects. Then, by looking at price reactions to mandatory contributions, I test whether the market anticipates that some of these funds would be wasted in negative NPV investments by the firms where agency conflicts are more severe.

I focus on the type of conflicts that arise when firms raise anti-takeover defenses. Being shielded from the disciplinary power of the market for corporate control, managers in these companies have more leeway to pursue their own interests as opposed to maximizing shareholders' value. Gompers, Ishii, and Metrick (2003) show that, over the long-run, firms with more anti-takeover provisions have worse operating performance and earn lower returns than the ones where shareholders have more rights. Using the governance index developed by these authors, I test whether price reactions to mandatory contributions vary across different levels of shareholders' protection. The hypothesis is that the same cash flow drop should be less costly for firms with weaker shareholders' rights, as part of it would be wasted in negative NPV projects.

Gompers, Ishii, and Metrick admit that their study cannot make strong claims about causality, because of the lack of random assignment. Specifically, weak firms could self-select into the group of companies with low shareholders' protection. Then, the positive link between good corporate governance and future performance would be the outcome of this endogeneity. Fortunately, this criticism is not likely to apply to the present set-up, where one can control for investment opportunities through the firm's funding status, as explained in Section I.

As the base model, I estimate a version of equation 1 where I_{it} is replaced by one of the three indices of corporate governance (G):

$$R_{i,t+1} = \beta_0 + \beta_1 MC_{it} + \beta_2 MC_{it} G_{it} + \beta_3 G_{it} + \beta_4 FS_{it} + \beta_5 X_{it} + \varepsilon_{it}, \tag{2}$$

Higher levels of G, imply worse governance. Hence, if agency conflicts determine overinvestment, one should expect β_2 to be positive.

Given the evidence in Section III that financial constraints matter for price reactions, it makes sense to control for these effects. Hence, I nest the costly-external-finance and empire-building models into one specification by including the FC index and its interaction with mandatory contributions. That is to say, I also run the regression:

$$\begin{aligned} R_{i,t+1} \quad = \quad & \beta_0 + \beta_1 MC_{it} + \beta_2 MC_{it} G_{it} + \beta_3 G_{it} + \beta_4 FS_{it} + \beta_5 X_{it} \\ & + \gamma_1 MC_{it} FC_{it} + \gamma_2 FC_{it} + \varepsilon_{it}, \end{aligned} \tag{3}$$

where FC stands for one of the three FC indices.

By imposing the availability of the governance indices, the sample is restricted to 4,107 firm-year observations on 834 firms. The average size in this smaller sample is \$3.3 billions, that is about 40% larger than the original sample. The average values of the WW, KZ, and Cleary indices are 0.56, 0.57, and 0.82. This means that, relative to the original sample, the firms in this smaller sample appear to be slightly less financially constrained, at least according to the first two measures (see Table I).

Table VIII uses Gompers, Ishii, and Metrick's (2003) G index. Column 1 of reports the estimates for equation 2.[26] The results provide some evidence in favor of differential price reactions across levels of agency conflicts. The slope on the interaction between G and MC is significant at the 10% level. One additional anti-takeover provision reduces the absolute value of the negative annual return due to a one-standard deviation increase in contributions by about 0.58%.

In the interpretational framework based on overinvestment, one could argue that the market believes that free cash flows amounting to 1% of capitalization would lead to wasteful projects

[26]The model includes all the control variables from the preferred specification (model 4) in Table IV.

totalling about 1.40% of firm value for each anti-takeover provision.[27] This figure becomes 14% for the average firm, which adopts about ten anti-takeover provisions. Still, given that the direct drop in price due to cash contributions of 1% of capitalization is about 17.6% (that is, 7.27%/0.41%), the total price reaction for the average firm in the year following an increase in MC is negative, amounting to about -3.6% annually (that is, $14\% - 17.6\%$).

One might wonder why the slope on MC is much higher than the magnitude it would have if there was a one-to-one direct effect of contributions on returns. The answer lies in the fact that the coefficient on MC originates from the integration over all levels of financial constraints for the firms in the sample, according to the interpretation provided in the previous section. Therefore, the resulting estimate includes the magnifying impact of financial constraints for the average firm.

Also noteworthy, the estimation of equation 2 tends to confirm the results in Gompers, Ishii, and Metrick (2003). Like these authors, I find a negative association between average returns and the governance index controlling for other known determinants of asset prices. The slope on G in column 1 of Table VIII is negative and significant at the 10% level. The interpretation they provide for this finding hinges on the market being repeatedly surprised over these years by the poor performance of firms in which the balance of power is tilted in favor of managers.

The reader might find an apparent contradiction between the forward looking behavior on investors' side that I assume in this paper and Gompers, Ishii, and Metrick's preferred explanation of their finding, which relies on the market's ignorance of the negative consequences of bad governance. Indeed, one can hold the two stories together by arguing that the impact of poor governance was far beyond what investors were anticipating in terms of magnitude and scope at the beginning of the sample. In this story, while investors foresee some of the negative implications of the agency conflicts, it takes them a number of years to learn about the full dimension of the phenomenon. One could test this explanation by verifying that, in a later sample, the interactive effect of agency conflicts and cash flow shocks on returns is still present, while the direct of effect of G is gradually disappearing.

Columns 2, 3, and 4 of Table VIII report estimates of equation 3 for the three FC indices. Despite the inclusions of the additional controls, in all three specifications, the effect of the interaction between the governance index and mandatory contributions is still present. In two specifications out of three (KZ and Cleary), the positive slope on the G index remains significant. Given the

[27]As argued earlier, the standard deviation of MC is about 0.24% of total assets, which for the average firm in this sample is 0.41% of market value.

smaller sample, the impact of financial constraints is significantly estimated only for the WW and KZ indices. In the first case, the significance level for the interaction between financial constraints and MC is 10%. Like before, there is a positive risk premium attached to the WW index. The results are mostly significant with the KZ index. In this case, the interactive effects of both the governance index and financial constraints attain 5% significance. Overall, the data suggest that the market cares about both financial constraints and corporate governance when reacting to cash flow shocks.

Next, I examine the robustness of these results to alternative definitions of the governance index. As in Masulis, Wang, and Xie (2007), I extend the analysis to two other measures of agency conflicts. First, I consider the index developed by Bebchuck, Cohen, and Ferrell (2004) who focus on a subset of six provisions out of the twenty-four originally used by Gompers, Ishii, and Metrick (2003). The authors show that this parsimonious index has a stronger association with long-run returns than G. I label it G2. The second extension comes from Bebchuck and Cohen (2005) who single out staggered boards as a key provision in forecasting returns. The dummy variable for the presence of a staggered board is labeled G3. Both G2 and G3 draw on the Investor Responsibility Research Center data and follow the same timing convention as G. The correlations of G2 and G3 with G are 73% and 49% respectively, while the correlation between G2 and G3 is 63%. The governance indices are constructed for firms without dual class shares. Hence, I restrict the sample to these companies. As a further robustness check, I focus on manufacturing industries (SIC codes 2000 through 3999). These requirements limit the sample to 2,305 observations on 435 firms.

Table IX replicates the structure of Table VIII replacing G with G2 (Panel A) and G3 (Panel B). The results in Panel A are comparable in terms of patterns of significance to those obtained with the G index. Again, the interaction between the governance index and MC is mostly significant when one controls for the effect of financial constraints through the KZ index. Across models, one additional anti-takeover provision in the G2 index attenuates the drop in value caused by a one standard deviation increase in MC by about 1% of market capitalization. For the average firm with G2 equal to about three, these estimates suggest that a cash flow of 1% of market value is translated into negative NPV projects totalling about 7.3% of capitalization (that is, $1\% \times 3/0.41$). Consistent with prior studies, the index by Bebchuck, Cohen, and Ferrell (2004) is negatively related to returns with more significant coefficients than G. As in Table VIII, the magnifying effect of financial constraints is captured only by the WW and KZ indices.

The evidence from Panel B is supportive of differential price reactions across levels of agency conflicts, but only when controlling for the impact of financial constraints. Most likely, these controls reduce the residual variance and contribute to increase estimation efficiency. The G3 index is negatively related to future returns, confirming the evidence in Bebchuck and Cohen (2005). Having a staggered board reduces the negative impact on market value of a drop in cash flow. According to the overinvestment interpretation and using the estimates in column 2, free cash flows amounting to 1% of capitalization are perceived as causing a waste of resources totalling 4.2% of market value in firms with a staggered board (that is, 1.74%/0.41).

The main conclusion from Tables VIII and IX is that the market's reaction to a cash flow shock varies across levels of the governance indices. If one interprets these indices as capturing the extent of agency conflicts between managers and shareholders, the results seem to support an interpretation based on overinvestment. In this view, the loss in value due to a drop in cash is attenuated by the fact that some of these resources would be used for wasteful projects. Moreover, the effect of financial constraints remains present alongside the agency conflicts story. Hence, a description of reality that possibly emerges from the data is one where the market considers both over- and underinvestment as relevant outcomes of firm behavior.

V. Underinvestment vs. Overinvestment

The empirical evidence suggests that the market differentiates firms across the financial constraints and agency conflicts dimensions when reacting to a cash flow shock. The proposed interpretation of these findings relates the value of the internal resources for a company to the distortions generated by costly external finance and empire-building projects. Although these distortions have opposite implications for investment, they can be nested in a model that admits both under- and overinvestment (see Stein (2003)). The question, then, is which type of frictions is more prevalent from an empirical point of view.

The present set-up allows one to partially address this question for the firms that are included in this study. Specifically, for each company one can compute the aggregate price reaction to a given cash flow shock. This return comprises three elements: the direct effect of the decrease in cash, which is obviously negative; the magnifying impact of financial constraints (also negative); and the attenuating impact of agency conflicts, which is positive. If no distortion was present, the price reaction would be negative and limited to the first effect. However, given that the direct effect of a one-standard deviation increase in MC for the average firm is a tiny 0.41% over a year (see the discussion in Section IV), the sign is effectively determined by the comparison of financial constraints and agency conflicts.[28] For some firms anticipated overinvestment could be so high that the effect of agency conflicts dominates and the resulting reaction is positive. Economically, one can interpret a positive return after a drop in cash as suggesting that the market believes that these funds would trigger a destruction of value, which exceeds their amount. In this case, the evidence unambiguously suggests that overinvestment is the prevalent phenomenon from investors' point of view. On the other hand, if the aggregate price reaction is negative (and it exceeds 0.41% in absolute value), one can conclude that the market gauges underinvestment as prevailing.

Concretely, I take equation 3 as the reference model because it allows for the effect of both financial constraints and agency conflicts. In this specification, the impact of MC, the cash flow shock, is measured through its direct effect (β_1), and through the interactions with the FC index and the governance index. Then, by using the estimates of the relevant parameters and firm measures of the FC index and G, I construct a composite index (CI) that reflects the aggregate reaction to a one-standard deviation shock for each firm:

$$CI_{it} = \hat{\beta}_1 + \hat{\beta}_2 G_{it} + \hat{\gamma}_1 FC_{it} + 0.41. \tag{4}$$

I draw the estimates of the relevant slopes from columns 2, 4, and 6 of Table VIII. In these specifications, $\hat{\beta}_1$ does not measure the impact of the cash flow shock in the absence of frictions. In fact, the FC indices are standardized. So, $\hat{\beta}_1$ measures the impact of MC for a firm at the average level of the FC index and with zero anti-takeover provisions. I add 0.41 to CI in order to filter out the direct effect of a cash flow shock in the absence of frictions. The sign of CI for a given firm in a given year reveals if under- or overinvestment is the prevalent distortion for that firm. The magnitude of CI measures the price reaction to a one-standard deviation increase in mandatory contributions.

Figure 1 plots the distributions of CI for firm-year observations using the three different FC indices. In all cases, the average and median CI are negative, suggesting that underinvestment is the prevalent distortion for the majority of firms in the sample. In particular, with the WW index 65% of firm-year observations display a negative total price reaction to a cash flow shock. The number drops to 57% with the KZ index. Finally, in the case of the Cleary index, 72% of the observations have negative CI. Given that the Cleary index does not produce significant coefficients in Table VIII, one should trust more the CI indices that depend on the other two measures of financial constraints.

[28]The reader should remember that mandatory contributions, which represent the cash flow shock, are measured in standard deviation units.

The fact that underinvestment is the main distortion in equilibrium is consistent with previous evidence (see Stein (2003)). For example, McConnell and Muscarella (1985) find that the average price reaction to the announcement of capital expenditures is positive, consistent with the marginal project having a positive NPV. Instead, it is negative in the oil industry where, according to Jensen (1986), there was systematic overinvestment in those years.

Most likely, the CI indices in this study provide a lower bound for the incidence of underinvestment. The reason is that the defined benefit firms with data on mandatory contributions, to which the sample is restricted, are on average larger than the firms in the market. In particular, while the average capitalization in this sample is \$2,387 millions, the average size in the market in the same years is about \$1,044 millions.

To conclude, one can investigate the nature of the firms that have different exposures to the two distortions. Specifically, I rank the firm-year observations into three groups by the levels of the CI indices (bottom 25%, intermediate two quarters, and top 25% of the distribution). According to the proposed interpretation of the CI index, the observations in the bottom group are the most exposed to financial constraints and the least to agency conflicts. The opposite is true for the top group.

Table X provides summary statistics on several firm characteristics for the three CI groups and the three FC indices. By construction, there is positive correlation between CI and G, while the correlation is negative between CI and the FC indices. More interesting, the observations that rank higher by CI have a better funding status (FS) and, as a consequence, pay lower mandatory contributions (MC). This fact could be evidence that the firms that are more prone to overinvestment (high CI) are those whose cash flows are not committed to the pension plan, consistent with Jensen's (1986) free cash flow hypothesis. No consistent pattern is found for the yearly returns after the ranking. Instead, high CI firms are relative winners in terms of past-twelve-month returns. As one would expect, firms that are more subject to overinvestment are larger (by both market capitalization and total assets) and older. The same companies are growth rather value. The firms for which financial constraints are the relevant distortion have lower analyst coverage and are less likely to have debt ratings, consistent with asymmetric information explanations of financial frictions. Although not consistent across FC indices, the pattern of capital expenditures across CI groups tends to be hump-shaped. In interpreting this finding, one could argue that costly external finance limits investment of low CI firms, while at high CI levels managers tend to divert funds to other uses than capital investment. As one would expect, firms that underinvest have the lowest R&D expenditures. The evidence of positive correlation between CI and cash flows is consistent with the reliance of financially constrained firms on internal funds and with the free cash flow theories of overinvestment. Except for the WW index, the firms that are more subject to underinvestment are more leveraged. This result largely depends on the definitions of the FC indices that load positively on leverage. The pattern for cash holdings is ambiguous, as different FC indices yield different correlations of cash with CI. Instead, high CI firms are consistently more likely to pay dividends and their total amount is higher. Somewhat counterintuitive from the point of view of the marginal profitability of investment, the link between Tobin's Q and CI is positive. However, this is the result of the univariate analysis and it disappears in multivariate regressions that control for other firm characteristics such as size (results not reported). The fact that the correlation between firm sales growth and CI is positive possibly justifies the availability of funds for overinvestment. No clear ranking can be made in terms of the ratio of current assets to liabilities, interest coverage, and a measure of financial slack. Finally, the firms that are more likely to overinvest make higher margins on their sales, possibly implying that profitability provides manager with the funds to engage in empire-building activities.

Overall, one can interpret this exercise as suggesting that underinvestment is the prevailing distortion in the market. Consistent with general theoretical predictions, the descriptive statistics reveal that firms that overinvest are larger, older, have higher coverage by analysts and rating agencies, and benefit from larger cash flows than constrained companies.

VI. Conclusions

The projection onto firm value of a negative cash flow shock incorporates the face value of the lost funds but also the shadow value of this cash for the firm. If the company depends on internal resources for its investments, i.e. it is financially constrained, then the cost of the cash flow shock exceeds its face value. On the other hand, if the managers would have used these funds for wasteful projects (i.e. empire building), then the impact on firm value is attenuated relative to the amount of the shock. The paper studies the market's assessment of the shadow value of internal resources.

Mandatory contributions to defined benefit pension plans allow the researcher to compare cash flow shocks keeping investment opportunities constant. Using the identification device developed by Rauh (2006), one can control for marginal profitability by using the pension plan funding status, while focusing on mandatory contribution as a source of cash flow shock.

By studying price reactions to pension contributions across different levels of financial constraints and potential agency conflicts, the paper aims at estimating the market's assessment of these two departures from a frictionless model. Financial constraints are measured with three indices that are commonly used in the literature: the Whited and Wu (2006) index, the Kaplan and Zingales (1997) index, and the Cleary (1999) index. To capture agency conflicts, I use the Gompers, Ishii, and Metrick (2003) index of corporate governance. As in these authors' work, the assumption is that bad governance gives the managers opportunities to pursue empire-building projects.

The data reveal that the market reacts significantly more strongly to a cash flow shock that hits a financially constrained firm, irrespectively of the chosen measure of constraints. An approximate calculation suggests that a drop in cash amounting to 1% of capitalization causes an additional drop in value by about 3.6% for a constrained firm relative to the price reaction for an unconstrained one. This difference is interpreted as the cost of external funds and the value of forsaken investment projects for a financially constrained company. Simple descriptive evidence seems to suggest that these estimates are consistent with realistic investment behavior. The results survive after a battery of robustness checks.

The other set of results concerns the market's assessment of the implications of agency conflicts. In this case, the price reactions to a given drop in cash are attenuated by the presence of agency conflicts, as measured by the governance index. Keeping everything else constant, a drop in cash of 1% of market capitalization is estimated to prevent the average firm from wasting between 4.6% and 14% of its value, depending on the chosen measure of the quality of governace. These estimates probably overstate the true effect, as the sample is tilted towards large firms for which overinvestment is more of a concern. In any case, the evidence suggests that the market anticipates that agency conflicts bring about a significant destruction of value.

Finally, I compute the market's joint reaction to the two types of distortions in order to assess whether under- or overinvestment prevails in this sample. This exercise suggests that underinvestment is the prevalent phenomenon for at least 58% of the firm-years in the sample. This estimate provides a lower bound for more representative samples where the average firm is smaller and more subject to financial constraints. Consistent with the theory, the firms that are more likely to overinvest appear to be larger, older, more covered by analysts and rating agencies, and to produce more cash flows than the companies for which underinvestment is the main distortion.

From the point of view of corporate finance, this work makes a comprehensive statement concerning the two main departures from a frictionless model of investment. By studying the market's reaction to a shift in internal funds, the paper concludes that both under- and overinvestment significantly affect value in equilibrium. Also, it provides a broad quantification of the economic importance of these phenomena and their distribution across the spectrum of listed firms.

Throughout the paper, it is the market's point of view that is considered. The assessment of these distortions to investment may not be accurate if investors are affected by some type of irrationality. For example, investors may temporarily misvalue the company and opportunistic managers can cater this mispricing (e.g., Stein (1996), Baker and Wurgler (2002), Baker, Stein, and Wurgler (2003), Polk and Sapienza (2007)). However, the paper does not take a stance on these issues. The goal is the estimation of investors' perception of the value of internal resources for different types of firms, irrespectively of rational behavior.

Obviously, the conclusions rely on the assumption that the chosen indices do actually capture the distortions of interest. Some of the descriptive evidence in the paper suggests that the characteristics of the firms that are differently exposed to these phenomena conform to a priori theoretical predictions concerning the identity of firms that under/overinvest. Future research should probably follow these companies in time to find further evidence of distortion in their investment and assess the consequences on their performance.

To conclude, it is worth mentioning the potential asset pricing implications of the results in the paper. A number of studies find evidence of factors related to financial constraints in asset prices (Perez-Quiros and Timmermann (2000), Lamont, Polk, and Saa-Requejo (2001), Gomes, Yaron, and Zhang (2006), and Whited and Wu (2006)). By showing that a given shock has different impact on firms with different exposure to financial frictions, the evidence in the paper provides support for the existence of financial constraints factors in returns. Also important, some literature has used measures of the quality of corporate governance in the firm to predict returns (Gompers, Ishii, and Metrick (2003), Bebchuck, Cohen, and Ferrell (2004), Bebchuck and Cohen (2005), Cremers and Nair (2005), and Cremers, Nair, and John (2005)). This work confirms governance as a relevant asset pricing variables, because the reactions to a given cash flow shock vary across levels of the governance indices. Finally, the paper confirms previous evidence in finding a significant premium in returns only for the Whited and Wu (2006) index. An open question is why the other indices fail to generate factors in spite of the fact that they identify significant reactions to a cash flow shock. In addressing this question, future research should probably study price reactions to shocks that, unlike pension contributions, are systematic in nature.

Appendix

A. Variable Definitions

In this section I provide definitions for the variables that are used in the empirical analysis and that are not already defined in Section II.

Book-to-Market. Book-to-market (as in Fama and French (1993)) is book value of equity at the end of the fiscal year (BE) divided by market capitalization in December. BE is stockholders' equity, plus balance sheet deferred taxes (Compustat item 74) and investment tax credit (item 208) (if available), plus post-retirement benefit liabilities (item 330) (if available), minus the book value of preferred stock. Depending on availability, I use redemption (item 56), liquidation (item 10), or par value (item 130) (in that order) for the book value of preferred stock. I calculate stockholders' equity used in the above formula as follows. I prefer the stockholders' equity number reported by Moody's (data on Ken French's website), or Compustat (item 216). If neither one is available, we measure stockholders' equity as the book value of common equity (item 60), plus the book value of preferred stock. If common equity is not available, I compute stockholders' equity as the book value of assets (item 6) minus total liabilities (item 181).

Tobin's Q. Tobin's Q is defined as: [Total Assets (item 6) + CRSP December Market Equity - Common Equity (item 60) - Deferred Taxes (item 74)]/Total Assets.

Accruals. Following Sloan (1996), accruals are defined as: [(One-year change in current assets (item 4)- One-year Change in cash (item 1)) - (One-year change in current liabilities (item 5) - One-year change in debt included in current liabilities (item 34) - One-year change in income taxes payable (item 71) - Depreciation and amortization expense (data 14)]/[Average of total assets over year t and year $t-1$].

Kaplan-Zingales Index. The index is defined normalizing variables by total assets as in Hennessy and Whited (2007):

$$-1.001909CF + 3.139193TLTD - 39.36780TDIV - 1.314759CASH + 0.2826389Q,$$

where CF is the ratio of cash flow (data 18 + data 14) to total assets (item 6), $TLTD$ is the ratio of total long-term debt (item 9) to total assets, $TDIV$ is the ratio of total dividends (item 19 + item 21) to book assets, $CASH$ is the ratio of cash (item 1) to total assets, and Q is Tobin's Q.

Whited-Wu Index. The index is:

$$0.938407 - 0.091CF - 0.062DIVPOS + 0.021TLTD - 0.044LNTA + 0.102ISG - 0.035SG,$$

in which $DIVPOS$ is an indicator that equals one if the firm pays dividends, and zero otherwise, SG is own-firm real sales growth, ISG is three-digit industry sales growth, and $LNTA$ is the natural log of total assets.[29]

Cleary index. The index is constructed as in Hennessy and Whited (2007) on the basis of Table II in Cleary (1999):

$$-0.11905CURAT - 1.903670TLTD + 0.00138COVER + 1.45618IMARG + 2.03604ISG$$
$$-0.04772SLACK,$$

in which $CURAT$ is the ratio of current assets to current liabilities; the numerator of $COVER$ is earnings before interest and taxes (item 13 - item 14), the denominator of $COVER$ is (interest expense (item 15) + preferred dividend payments (item 19))/ (1 - Tax Rate), where Tax Rate is: item 16/(item 13 - item 14 - item 15); $IMARG$ is the ratio of net income (item 18) to sales (item 12); the numerator of $SLACK$ is: cash + short-term investments + 0.5*inventory + 0.7*accounts receivable - short term loans, and denominator of SLACK is net fixed assets. Cash plus short term investments is item 1; inventory is item 3; accounts receivable is item 2; short terms loans is item 196; and net fixed assets is item 8. Extreme values of these variables are winsorized as in Cleary (1999). I multiply the Cleary index by -1 so that it is increasing in the likelihood of facing costly external finance.

The three financial constraints indices are winsorized at the first and ninety-nineth percentile.

Other variables that appear in the paper are: Age, which is the number of years since the company first appeared in the CRSP monthly file; Rating, a dummy variable equalling one if the firm's debt is rated according to item 280; Analysts, which is the average number of analysts following the firm in a given year according to IBES data (if the firm does not appear in the data set for a given year the variable is set to zero); Capital Expenditures (CAPEX), item 128; Research and Development, item 46.

[29]The constant in the Whited-Wu index has been kindly provided by Toni Whited.

B. Price Reactions in a One-Period Model of Investment

Following Kaplan and Zingales (1997), one can describe the firm's optimal choice of investment in a simple one-period model. In this model, the firm maximizes profits with respect to investment I. The cost of capital is normalized to one. The firm's problem is then:

$$\begin{aligned} Max\ \Pi &= F(I) - C(E,k) - I, \\ s.t.\ I &= W + E \end{aligned} \tag{A-1}$$

where E is the amount of external funds, W represents internal resources, and $C(E,k)$ is the cost of external funds, which depends on the wedge between internal and external finance k. As in Kaplan and Zingales (1997), the following natural assumptions are made:

$$\begin{aligned} F_1 &> 0,\ F_{11} < 0 \\ C_1 &> 0,\ C_2 > 0,\ C_{12} > 0 \end{aligned} \tag{A-2}$$

Now, the firm's constrained maximization problem can be solved as an unconstrained problem by replacing the constraint into the objective function:

$$\underset{E}{Max}\ F(W+E) - C(E,k) - (E+W). \tag{A-3}$$

The first order condition is then:

$$F_1 - C_1 - 1 = 0. \tag{A-4}$$

In this simple model, the value of the firm is equal to the firm's profit Π. Hence, if one wants to compute the price reaction to a shift internal resources, one needs to differentiate Π^*, that is the profit evaluated at the optimal level of investment, with respect to W. In order to do that, one can use the envelope theorem and disregard the change in optimal investment:

$$\frac{\partial \Pi^*}{\partial W} = F_1 - 1, \tag{A-5}$$

which is positive given the first order condition in equation A-4.

In this model, financial constraints are captured by the additional cost of external finance k and by the availability of internal resources W. Then, to compute the differential price reaction across levels of financial constraints, one has to take the derivative of the expression in A-5 with respect to either k or W. Let us first focus on k:

$$\frac{\partial^2 \Pi^*}{\partial W \partial k} = F_{11} \frac{dE^*}{dk}, \tag{A-6}$$

where $\frac{dE^*}{dk}$ is the derivative of the optimal level of external finance relative to k. This derivative can be obtained from implicit differentiation of the first order condition and it is

$$\frac{dE^*}{dk} = \frac{C_{12}}{F_{11} - C_{11}}, \tag{A-7}$$

which is negative due to the assumptions in A-2 and the fact that the denominator is negative by the second order condition for a maximum. Hence, this result along with $F_{11} < 0$ implies that the derivative in equation A-6 is positive. This conclusion proves that, in the context of this simple model, the increase in value following an increase in internal resources is larger for firms that are more constrained according to the cost of external funds k.

Analogously, one can show that

$$\frac{\partial^2 \Pi^*}{\partial^2 W} = F_{11} \frac{dE^*}{dW} \tag{A-8}$$

and

$$\frac{dE^*}{dW} = \frac{-F_{11}}{F_{11} - C_{11}} < 0 \tag{A-9}$$

So that, $\frac{\partial^2 \Pi^*}{\partial^2 W} > 0$, which proves that the increase in value following an increase in internal resources is larger for firms that have larger availability of internal funds (i.e., less constrained).

The intuition is the same for both results, and it is straightforward. An increase in internal funds allows constrained firms to increase their investment. Given that the marginal productivity of capital is decreasing ($F_{11} < 0$), this additional investment is more valuable for more constrained firms for which the initial investment is lower to begin with and it is, therefore, marginally more productive.

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Table I: Summary Statistics. The table reports summary statistics and correlations for the main variables in the analysis. The variables in the table are: mandatory contributions in year t over total assets in year $t-1$ (MC); the aggregate funding status in year t over total assets in year $t-1$ (FS); Tobin's Q at the end of year t; the stock's beta in December of year t; market capitalization in June of year $t+1$ (size, measured in $ millions); book-to-market (BM) in December of year t; past-twelve-months cumulative returns (Mom) in June of year $t+1$; Accruals over total assets in December of year t; the Whited and Wu (2006) index (WW); the Kaplan and Zingales (1997) index (KZ); the Cleary (1999) index (Cleary); and the Gompers, Ishii, and Metrick (2003) index of corporate governance (G). Higher levels of the financial constraints indices denote more financially constrained firms. The G index is measured using the latest publication of the Investor Responsibility Research Center as of year $t+1$. The sample is limited to the period and firms for which mandatory contributions are available (1990-1999), as collected by Rauh (2006). Excluding the selection criterion based on G, the sample consists of 6,119 firm-year observations on 1,198 firms. Conditioning on the availability of G, the sample is restricted to 4,107 firm-year observations on 834 firms.

	Mean	Median	St. dev.	Min	Max
MC	0.001	0.000	0.002	0.000	0.023
FS	0.029	0.013	0.050	-0.095	0.265
Q	1.513	1.264	0.875	0.338	10.744
Beta	0.921	0.904	0.490	-1.279	3.760
Size	2387	566	4511	1	36471
BM	0.816	0.643	0.874	0.026	21.564
Mom	0.134	0.094	0.361	-0.831	3.400
Acc	-0.044	-0.043	0.064	-0.492	0.342
WW	0.600	0.597	0.089	0.384	0.846
KZ	0.631	0.693	1.173	-5.952	5.501
Cleary	0.821	0.824	0.524	-1.248	4.698
G	9.819	10.000	2.759	2.000	18.000

Correlations	MC	FS	Q	Beta	Size	BM	Mom	Acc	WW	KZ	Cleary
MC	1.00										
FS	-0.31	1.00									
Q	-0.07	0.02	1.00								
Beta	0.02	-0.11	0.07	1.00							
Size	-0.09	0.03	0.42	0.02	1.00						
BM	0.08	-0.04	-0.40	-0.04	-0.21	1.00					
Mom	0.01	0.00	0.13	0.05	0.11	-0.15	1.00				
Acc	0.02	-0.03	0.08	0.00	-0.03	-0.02	-0.04	1.00			
WW	0.15	-0.08	-0.21	-0.06	-0.63	0.20	0.00	0.02	1.00		
KZ	0.12	-0.22	-0.33	0.12	-0.11	0.15	0.01	-0.02	0.16	1.00	
Cleary	0.06	-0.08	-0.28	0.03	-0.09	0.10	-0.05	-0.06	-0.02	0.59	1.00
G	-0.01	0.09	0.00	0.09	0.03	-0.08	0.02	-0.03	-0.09	0.01	0.00

Table II: Descriptive Statistics By MC Groups. The table reports sample means and medians of different firm characteristics by levels of mandatory contributions (MC). The first two columns use all firm-year observations for which MC=0 (4,366 observations). The third and fourth columns use the bottom 50% of observations for which MC>0 (877 observations). The fifth and sixth columns use the top 50% of observations for which MC>0 (876 observations). The median of MC, conditioning on MC>0, is about 0.09%. The firm characteristics are: mandatory contributions over assets (MC, in percent), the funding status over assets (FS, in percent), the financial constraints indices (WW, KZ, Cleary), the governance index (G), the stock beta, market capitalization in June of year t (Size, in $ millions), book-to-market (BM), Tobin's Q in December of year $t-1$, firm's age, a dummy for firms whose bonds are rated, the average number of analysts following the firm in the twelve months up to June of year t, cash flows over assets (CF), total long term debt over assets (TLTD), total dividends over assets (TDIV), a dummy that equals one if the firm pays dividends (DIVPOS), total cash over assets (CASH), capital expenditures over assets, research and development expenses over assets, the firm's reals sales growth in year $t-1$ (SG), the coverage ratio for interests and preferred dividends (COVER), the ratio of net income to net sales (IMARG), a measure of financial slack (SLACK). See the text and the appendix for the details on variable definitions.

			MC>0			
	MC=0		Bottom 50%		Top 50%	
	Mean	Median	Mean	Median	Mean	Median
MC (%)	0.00	0.00	0.03	0.02	0.46	0.28
FS (%)	3.69	1.90	2.21	0.52	-0.51	-0.55
WW	0.60	0.59	0.58	0.57	0.64	0.64
KZ	0.52	0.59	0.73	0.82	1.09	1.08
Cleary	0.79	0.81	0.88	0.86	0.92	0.89
G	9.81	10.00	9.80	10.00	9.91	10.00
Beta	0.88	0.86	1.05	1.03	1.00	0.98
Size	2462	614	3259	967	1137	204
BM	0.80	0.64	0.74	0.60	0.96	0.71
Q	1.55	1.28	1.52	1.30	1.33	1.18
Age	28.36	25.00	29.80	26.00	24.82	22.00
Rating	0.46	0.00	0.61	1.00	0.38	0.00
Analyst	7.76	4.17	9.21	7.00	5.21	2.67
CF	0.11	0.10	0.10	0.10	0.09	0.09
TLTD	0.37	0.39	0.41	0.41	0.43	0.41
TDIV	0.02	0.02	0.02	0.02	0.01	0.01
DIVPOS	0.83	1.00	0.81	1.00	0.66	1.00
CASH	0.07	0.03	0.06	0.03	0.07	0.03
CAPEX	0.08	0.07	0.07	0.06	0.07	0.06
RD	0.02	0.00	0.02	0.00	0.02	0.01
SG	0.05	0.03	0.04	0.02	0.05	0.03
COVER	24.03	5.53	21.83	5.31	25.54	4.88
IMARG	0.05	0.05	0.04	0.04	0.02	0.03
SLACK	1.21	0.63	1.00	0.68	1.19	0.79

Table III: Return Reactions To Mandatory Contributions. The table reports estimates from the OLS regressions of annual returns (in percent) on the amount of mandatory contributions to total assets (MC, in standard deviation units), and controls. The dependent variable is the compound return between July of year t and June of year $t+1$. All specifications include: the aggregate funding status over assets (FS), Tobin's Q, and year fixed effects. All models but (1) control for the stock beta, book-to-market (BM), and size. Models (3) to (6) include accruals over assets and past-twelve-months returns (Mom). Models (4) to (6) include industry dummies computed using three-digit SIC codes. In Model (5), the variables beta, BM, size, Acc, Mom, and industry dummies are interacted with year fixed effects. Model (6) includes firm fixed effects. All of the control variables are measured in December of year $t-1$, except for size and Mom that are computed in June of year t. Annual returns range between July of 1991 and June of 2001. The sample consists of 6,119 observations on 1,198 firms. Standard errors are clustered by firm and year according to the procedure in Thompson (2006).

	(1)	(2)	(3)	(4)	(5)	(6)
MC	-1.56	-1.42	-1.47	-2.04	-1.86	-1.62
	(-3.49)	(-3.04)	(-3.20)	(-3.93)	(-3.60)	(-2.65)
FS	-6.83	-1.96	-4.58	-16.24	-11.79	-9.80
	(-0.58)	(-0.18)	(-0.43)	(-1.37)	(-0.87)	(-0.69)
Q	-0.09	-1.13	-0.88	-0.53	-0.38	-3.25
	(-0.05)	(-0.60)	(-0.50)	(-0.30)	(-0.24)	(-1.30)
Beta, BM, Size	No	Yes	Yes	Yes	Yes	Yes
Acc, Mom	No	No	Yes	Yes	Yes	Yes
Industry FE	No	No	No	Yes	Yes	Yes
Controls*Time	No	No	No	No	Yes	No
Firm FE	No	No	No	No	No	Yes
Time FE	Yes	Yes	Yes	Yes	Yes	Yes

Table IV: Return Reactions By Financial Constraints Groups. The table reports estimates from OLS regressions of annual returns (in percent) on the amount of mandatory contributions to total assets (MC, in standard deviation units), the interactions between mandatory contributions and dummy variables that denote firms classified into the high (top 25%) and medium (between 25% and 75%) groups of the financial constraints index for year $t-1$, and a number of control variables. The controls that are common to all specifications are: the stock beta, book-to-market (BM), size, the aggregate funding status over assets (FS), Tobin's Q, an index of financial constraints, and year fixed effects. The financial constraints indices are: the Whited and Wu (2006) index, the Kaplan and Zingales (1997) index, and the Cleary (1999) index. Higher levels of the financial constraints indices denote more financially constrained firms. The three indices have been standardized. Models (2) to (5) include accruals over assets (Acc) and past-twelve-months returns (Mom). Models (3) to (5) include industry dummies computed using three-digit SIC codes. In Model (4), the variables beta, BM, size, Acc, Mom, and industry dummies are interacted with year fixed effects. Model (5) includes firm fixed effects. All of the control variables are measured in December of year $t-1$, except for size and Mom that are computed in June of year t. Annual returns range between July of 1991 and June of 2001. The sample consists of 6,119 observations on 1,198 firms. Standard errors are clustered by firm and year according to the procedure in Thompson (2006). T-statistics are reported in parentheses.

	(1)			(2)			(3)			(4)			(5)		
	WW	KZ	Cleary	WW	KZ	Cleary	WW	KZ	Cleary	WW	KZ	Cleary	WW	KZ	Cleary
FC Index	4.44	0.30	-0.19	4.11	0.12	-0.11	3.68	0.17	-0.29	4.44	0.40	-0.23	13.18	5.71	-3.96
	(3.68)	(0.26)	(-0.24)	(3.77)	(0.10)	(-0.14)	(3.01)	(0.15)	(-0.37)	(2.16)	(0.39)	(-0.31)	(2.58)	(2.26)	(-3.19)
MC	0.11	-0.13	0.13	0.05	-0.20	0.02	-0.60	-0.71	-0.54	-0.01	-0.33	-0.42	3.28	0.87	0.30
	(0.12)	(-0.14)	(0.26)	(0.06)	(-0.21)	(0.03)	(-0.68)	(-0.68)	(-0.83)	(-0.01)	(-0.34)	(-0.73)	(3.66)	(0.63)	(0.29)
Med FC*MC	-1.33	-0.44	-1.86	-1.34	-0.45	-1.66	-1.15	-0.49	-1.65	-1.38	-0.59	-1.51	-2.77	-0.40	-0.15
	(-1.27)	(-0.34)	(-2.24)	(-1.26)	(-0.36)	(-2.09)	(-1.14)	(-0.39)	(-2.03)	(-1.31)	(-0.53)	(-2.20)	(-1.74)	(-0.24)	(-0.13)
Hi FC*MC	-2.73	-2.65	-2.37	-2.64	-2.58	-2.44	-2.28	-2.66	-2.50	-2.85	-3.02	-2.44	-4.57	-1.79	-0.58
	(-2.42)	(-2.36)	(-2.18)	(-2.26)	(-2.40)	(-2.40)	(-2.03)	(-2.44)	(-2.42)	(-2.15)	(-2.78)	(-2.91)	(-3.61)	(-0.94)	(-0.30)
Acc, Mom		No			Yes			Yes			Yes			Yes	
Industry FE		No			No			Yes			Yes			Yes	
Controls*Time		No			No			No			Yes			No	
Firm FE		No			No			No			No			Yes	

Table V: Return Reactions By Financial Constraints Groups: Further Controls. The table reports estimates from OLS regressions of annual returns (in percent) on the amount of mandatory contributions to total assets (MC, in standard deviation units), the interactions between mandatory contributions and dummy variables that denote firms classified into the high (top 25%) and medium (between 25% and 75%) groups of the financial constraints index for year $t-1$, and a number of control variables. The controls that are common to all specifications are: the stock beta, book-to-market (BM), size, accruals over assets (Acc), past-twelve-months returns (Mom), the aggregate funding status over assets (FS), Tobin's Q, an index of financial constraints (FC), industry dummies computed using three-digit SIC codes, and year fixed effects. All the controls (except for Q, FS, and the FC index) are interacted with year fixed effects. The reported estimates refer to the financial constraint index, MC, and the interaction between MC and the dummy for the high financial constraints group. The financial constraints indices are: the Whited and Wu (2006) index, the Kaplan and Zingales (1997) index, and the Cleary (1999) index. Higher levels of the financial constraints indices denote more financially constrained firms. The three indices have been standardized. Model 1 is the base specification. Model 2 includes an interaction between MC and FS. Model 3 includes an interaction between MC and Tobin's Q. Model 4 includes an interaction between MC and beta. Model 5 includes an interaction between MC and BM. Model 6 includes an interaction between MC and prior-twelve-month returns (Mom). Model 7 includes an interaction between MC and market capitalization. Model 8 includes an interaction between MC and the number of analysts following the firm. Model 9 includes separate controls for the surplus and deficit in the pension plans normalized by total assets, and the second and third powers of these variables. Model 10 restricts the sample to manufacturing firms (SIC codes 2000 through 3999). Model 11 estimates the base model in the years before 1996. Model 12 estimates the base model in the years after 1995. Model 13 is like model 1, but Tobin's Q is excluded from the KZ index. All of the control variables are measured in December of year $t-1$, except for size and Mom that are computed in June of year t. Annual returns range between July of 1991 and June of 2001. Standard errors are clustered by firm and year according to the procedure in Thompson (2006). T-statistics are reported in parentheses.

	WW			KZ			Cleary			
	FC	MC	Hi*MC	FC	MC	Hi*MC	FC	MC	Hi*MC	Obs.
1. Base Model	4.44	-0.01	-2.85	0.40	-0.33	-3.02	-0.23	-0.42	-2.44	6,119
	(2.16)	(-0.01)	(-2.15)	(0.39)	(-0.34)	(-2.78)	(-0.31)	(-0.73)	(-2.91)	
2. Ctrl. FS*MC	4.40	0.77	-2.96	0.37	-0.15	-2.80	-0.19	-0.13	-2.07	6,119
	(2.12)	(0.92)	(-2.48)	(0.38)	(-0.13)	(-2.49)	(-0.25)	(-0.18)	(-1.88)	
3. Ctrl. Q*MC	4.46	1.40	-2.83	0.45	1.80	-3.34	-0.25	1.49	-2.68	6,119
	(2.16)	(0.82)	(-2.14)	(0.44)	(1.31)	(-3.24)	(-0.34)	(1.18)	(-3.39)	
4. Ctrl. Beta*MC	4.47	1.35	-3.47	0.39	0.05	-3.02	-0.23	-0.03	-2.44	6,119
	(2.17)	(0.74)	(-2.08)	(0.38)	(0.05)	(-2.67)	(-0.31)	(-0.04)	(-2.89)	
5. Ctrl. BM*MC	4.43	0.16	-2.80	0.39	-0.20	-2.93	-0.22	-0.18	-2.33	6,119
	(2.15)	(0.20)	(-2.06)	(0.39)	(-0.20)	(-2.77)	(-0.29)	(-0.28)	(-2.94)	
6. Ctrl. Mom*MC	4.44	-0.01	-2.85	0.40	-0.30	-3.02	-0.23	-0.37	-2.46	6,119
	(2.16)	(-0.01)	(-2.23)	(0.39)	(-0.31)	(-2.75)	(-0.31)	(-0.53)	(-2.92)	
7. Ctrl. Size*MC	4.44	-0.20	-2.67	0.41	-0.48	-2.97	-0.23	-0.55	-2.38	6,119
	(2.14)	(-0.14)	(-1.27)	(0.40)	(-0.57)	(-2.80)	(-0.31)	(-1.12)	(-2.78)	
8. Ctrl. Analyst*MC	4.44	0.69	-3.48	0.40	-0.20	-3.06	-0.23	-0.38	-2.46	6,119
	(2.16)	(0.53)	(-2.05)	(0.39)	(-0.23)	(-2.87)	(-0.31)	(-0.79)	(-2.87)	
9. Powers of FS	4.54	1.05	-3.02	0.56	0.57	-3.38	-0.26	0.31	-2.31	6,119
	(2.17)	(1.05)	(-2.18)	(0.56)	(0.65)	(-2.40)	(-0.34)	(0.38)	(-2.19)	
10. Mfg. Firms	7.19	-0.45	-2.45	0.08	0.42	-4.58	0.48	-0.17	-3.51	3,861
	(2.55)	(-0.96)	(-1.77)	(0.08)	(0.28)	(-2.57)	(0.48)	(-0.18)	(-3.01)	
11. Year $<$ 1996	3.66	-0.63	-2.14	1.96	-0.52	-3.41	-1.07	-0.89	-2.88	3,547
	(3.97)	(-0.87)	(-1.95)	(3.71)	(-0.54)	(-2.59)	(-1.08)	(-1.40)	(-2.69)	
12. Year $>$ 1995	5.49	0.98	-4.05	-1.98	-0.20	-2.10	0.90	0.15	-1.66	2,572
	(1.10)	(4.27)	(-1.82)	(-1.17)	(-0.13)	(-1.22)	(1.14)	(0.16)	(-1.16)	
13. KZ4				0.37	-0.89	-2.39				6,119
				(0.33)	(-1.08)	(-2.22)				

Table VI: Return Reactions By Financial Constraints Groups: Monthly Returns. The table reports estimates from OLS regressions of monthly returns (in percent) on the amount of mandatory contributions to total assets (MC, in standard deviation units), the interactions between mandatory and an index of financial constraints for year $t-1$ (FC), and a number of control variables. The controls that are common to all specifications are: the stock beta, book-to-market (BM), size, accruals over assets (Acc), past-twelve-months returns (Mom), the aggregate funding status over assets (FS), Tobin's Q, an index of financial constraints, industry dummies computed using three-digit SIC codes, and year×month fixed effects. The reported estimates refer to the financial constraint index, MC, and the interaction between MC and FC. The financial constraints indices are: the Whited and Wu (2006) index, the Kaplan and Zingales (1997) index, and the Cleary (1999) index. Higher levels of the financial constraints indices denote more financially constrained firms. The three indices have been standardized. Model 1 is the base specification. Model 2 includes an interaction between MC and FS. Model 3 includes an interaction between MC and Tobin's Q. Model 4 includes an interaction between MC and beta. Model 5 includes an interaction between MC and BM. Model 6 includes an interaction between MC and prior-twelve-month returns (Mom). Model 7 includes an interaction between MC and market capitalization. Model 8 includes an interaction between MC and the number of analysts following the firm. Model 9 includes separate controls for the surplus and deficit in the pension plans normalized by total assets, and the second and third powers of these variables. Model 10 restricts the sample to manufacturing firms (SIC codes 2000 through 3999). Model 11 estimates the base model in the years before 1996. Model 12 estimates the base model in the years after 1995. Model 13 is like model 1, but Tobin's Q is excluded from the KZ index. The control variables in December of year $t-1$ (except for size and Mom that are computed in June of year t) are associated to monthly returns between July of year t and June of year $t+1$. Monthly returns range between July of 1991 and June of 2001. Standard errors are clustered by firm and month according to the procedure in Thompson (2006). T-statistics are reported in parentheses.

	WW			KZ			Cleary			
	FC	MC	FC*MC	FC	MC	FC*MC	FC	MC	FC*MC	Obs.
1. Base Model	0.21	-0.12	-0.08	0.08	-0.11	-0.11	0.05	-0.14	-0.10	77,099
	(1.33)	(-2.60)	(-1.94)	(1.08)	(-2.93)	(-3.00)	(0.81)	(-3.55)	(-2.70)	
2. Ctrl. FS*MC	0.21	-0.07	-0.08	0.08	-0.10	-0.10	0.05	-0.12	-0.10	77,099
	(1.30)	(-1.23)	(-1.98)	(1.06)	(-1.74)	(-2.55)	(0.78)	(-1.95)	(-2.35)	
3. Ctrl. Q*MC	0.21	-0.19	-0.08	0.08	-0.16	-0.10	0.05	-0.17	-0.10	77,099
	(1.32)	(-2.00)	(-1.89)	(1.07)	(-1.95)	(-2.95)	(0.80)	(-2.28)	(-2.64)	
4. Ctrl. Beta*MC	0.21	-0.12	-0.08	0.08	-0.16	-0.11	0.05	-0.17	-0.10	77,099
	(1.32)	(-1.36)	(-1.75)	(1.10)	(-2.49)	(-3.12)	(0.81)	(-2.67)	(-2.70)	
5. Ctrl. BM*MC	0.21	-0.10	-0.08	0.08	-0.10	-0.10	0.05	-0.12	-0.10	77,099
	(1.32)	(-2.08)	(-1.92)	(1.07)	(-2.12)	(-3.02)	(0.79)	(-2.54)	(-2.69)	
6. Ctrl. Mom*MC	0.21	-0.10	-0.08	0.08	-0.10	-0.11	0.05	-0.12	-0.10	77,099
	(1.33)	(-2.01)	(-1.88)	(1.10)	(-2.28)	(-2.85)	(0.82)	(-2.84)	(-2.61)	
7. Ctrl. Size*MC	0.21	-0.13	-0.07	0.08	-0.13	-0.10	0.05	-0.15	-0.10	77,099
	(1.30)	(-2.56)	(-1.62)	(1.09)	(-3.08)	(-2.95)	(0.81)	(-3.63)	(-2.61)	
8. Ctrl. Analyst*MC	0.21	-0.15	-0.06	0.08	-0.14	-0.10	0.05	-0.16	-0.10	77,099
	(1.32)	(-2.56)	(-1.28)	(1.09)	(-2.71)	(-2.77)	(0.80)	(-3.33)	(-2.59)	
9. Powers of FS	0.21	-0.06	-0.09	0.09	-0.08	-0.12	0.05	-0.10	-0.11	77,099
	(1.34)	(-0.95)	(-1.95)	(1.18)	(-1.17)	(-3.05)	(0.83)	(-1.62)	(-2.63)	
10. Mfg. Firms	0.33	-0.16	-0.07	0.10	-0.11	-0.18	0.02	-0.16	-0.14	48,803
	(1.83)	(-3.35)	(-1.51)	(1.26)	(-2.67)	(-5.25)	(0.25)	(-3.43)	(-3.75)	
11. Year < 1996	0.21	-0.16	-0.05	0.17	-0.14	-0.10	0.09	-0.17	-0.08	44,228
	(1.35)	(-3.15)	(-0.93)	(2.02)	(-3.25)	(-3.24)	(1.17)	(-3.71)	(-1.85)	
12. Year > 1995	0.17	-0.07	-0.11	-0.07	-0.08	-0.10	-0.01	-0.11	-0.11	32,871
	(0.50)	(-1.10)	(-1.64)	(-0.60)	(-1.24)	(-1.41)	(-0.08)	(-1.87)	(-1.96)	
13. KZ4				0.08	-0.11	-0.11				77,099
				(1.07)	(-2.78)	(-3.03)				

Table VII: Portfolio Analysis By Funding Ratio And Financial Constraints Groups. The table reports alphas from regressions of portfolio monthly excess returns (in percent) on the Fama and French (1993) three factors. The portfolios are formed by sorting stocks by the funding ratio (FR), for firms with FR<0, and by an index of financial constraints. FR is constructed as the ratio of Assets in the Pension Plans minus Projected Benefit Obligation to market capitalization in December. In July of year t, five quintiles are formed according to the distribution of FR in year $t-1$, for firms with FR<0, and using NYSE breakpoints. The first column reports results for portfolios that are formed value-weighting the returns in each FR quintile between July of year t and June of year $t+1$. For the rest of the table, the stocks in each FR quintile are further sorted into three groups using the distribution of an index of financial constraints (FC) in year $t-1$ for the stocks in that FR quintile. The three groups are: Low (bottom 20%), Medium (between 20% and 80%), High (top 80%). Value-weighted portfolios are formed with the stock in each of the resulting groups. The three FC indices are: the Whited and Wu (2006) index, the Kaplan and Zingales (1997) index, and the Cleary (1999) index. Higher levels of the FC indices denote more financially constrained firms. For each index, and for each FR quintile, the table also reports the alpha for the portfolio that goes long in the Low FC and short in the High FC stocks. Monthly returns range between July of 1981 and December of 2005. T-statistics are reported in parentheses.

	Alphas												
		WW				KZ				Cleary			
	No Index	Lo	Med	Hi	Lo - Hi	Lo	Med	Hi	Lo - Hi	Lo	Med	Hi	Lo - Hi
1: Most Underf.	-0.66	-0.58	-0.63	-0.16	-0.42	0.09	-0.69	-1.54	1.63	-0.06	-0.67	-1.17	1.12
	(-3.03)	(-2.23)	(-3.26)	(-0.48)	(-1.10)	(0.36)	(-2.96)	(-3.29)	(3.12)	(-0.21)	(-2.81)	(-2.93)	(2.40)
2	-0.01	0.12	-0.20	-0.49	0.61	0.26	-0.05	-0.79	1.04	0.35	-0.02	-0.46	0.81
	(-0.09)	(0.70)	(-1.33)	(-1.74)	(1.90)	(1.09)	(-0.30)	(-2.49)	(2.87)	(1.36)	(-0.12)	(-1.77)	(2.35)
3	0.20	-0.04	0.28	-0.27	0.23	0.21	0.15	-0.60	0.81	0.49	0.09	0.09	0.40
	(1.31)	(-0.24)	(2.03)	(-0.98)	(0.77)	(1.12)	(1.02)	(-2.30)	(2.73)	(1.86)	(0.62)	(0.36)	(1.12)
4	-0.02	-0.06	0.08	-0.22	0.16	0.59	-0.17	-0.61	1.21	0.28	0.01	-0.12	0.40
	(-0.14)	(-0.40)	(0.59)	(-0.94)	(0.55)	(2.69)	(-1.13)	(-2.14)	(3.23)	(1.09)	(0.07)	(-0.47)	(1.06)
5: Least Underf.	0.03	0.01	-0.07	0.72	-0.70	0.24	-0.10	-0.05	0.29	-0.09	0.04	-0.30	0.21
	(0.23)	(0.09)	(-0.52)	(2.26)	(-2.02)	(1.09)	(-0.71)	(-0.18)	(0.80)	(-0.38)	(0.30)	(-1.21)	(0.60)

Table VIII: Return Reactions And Corporate Governance Index. The table reports estimates from OLS regressions of annual returns (in percent) on the amount of mandatory contributions to total assets (MC, in standard deviation units), the Gompers, Ishii, and Metrick (2003) index of corporate governance (G), the interaction between G and MC, and a number of control variables. The controls that are common to all specifications are: the stock beta, book-to-market (BM), size, accruals over assets (Acc), past-twelve-months returns (Mom), the aggregate funding status over assets (FS), Tobin's Q, an index of financial constraints, industry dummies computed using three-digit SIC codes, and year fixed effects. All the controls (except for Q and FS) are interacted with year fixed effects. The models in columns two through four also include an index of financial constraints (FC) and the interaction between this index and MC. The financial constraints indices are: the Whited and Wu (2006) index, the Kaplan and Zingales (1997) index, and the Cleary (1999) index. Higher levels of the financial constraints indices denote more financially constrained firms. The three indices have been standardized. The G index is measured using the latest publication of the Investor Responsibility Research Center as of year t. The other control variables are measured in December of year $t-1$, except for size and Mom that are computed in June of year t. Annual returns range between July of 1991 and June of 2001. The sample consists of 4,107 observations on 834 firms. Standard errors are clustered by firm and year according to the procedure in Thompson (2006). T-statistics are reported in parentheses.

	(1)	(2)	(3)	(4)
	No FC	WW	KZ	Cleary
MC	-7.27	-7.05	-7.21	-7.32
	(-2.39)	(-2.22)	(-2.51)	(-2.44)
G	-0.30	-0.21	-0.30	-0.30
	(-1.88)	(-1.15)	(-1.93)	(-1.90)
MC*G	0.58	0.54	0.63	0.59
	(1.87)	(1.66)	(2.28)	(1.98)
FC		3.35	1.12	0.37
		(1.77)	(1.22)	(0.44)
MC*FC		-1.11	-1.39	-0.47
		(-1.82)	(-2.44)	(-0.72)

Table IX: Alternative Corporate Governance Indices. The table reports estimates from OLS regressions of annual returns (in percent) on the amount of mandatory contributions to total assets (MC, in standard deviation units), an index of corporate governance, the interaction between the governance index and MC, and a number of control variables. The two governance indices are: the Bebchuck, Cohen, and Ferrell (2004) index (G2) and the Bebchuck and Cohen (2005) index (G3), which is a dummy variable for firms with staggered boards. The controls that are common to all specifications are: the stock beta, book-to-market (BM), size, accruals over assets (Acc), past-twelve-months returns (Mom), the aggregate funding status over assets (FS), Tobin's Q, an index of financial constraints, industry dummies computed using three-digit SIC codes, and year fixed effects. All the controls (except for Q and FS) are interacted with year fixed effects. The models in columns two through four also include an index of financial constraints (FC) and the interaction between this index and MC. The financial constraints indices are: the Whited and Wu (2006) index, the Kaplan and Zingales (1997) index, and the Cleary (1999) index. Higher levels of the financial constraints indices denote more financially constrained firms. The three indices have been standardized. The governance index is measured using the latest publication of the Investor Responsibility Research Center as of year t. The other control variables are measured in December of year $t-1$, except for size and Mom that are computed in June of year t. The sample is restricted to manufacturing firms (SIC codes 2000 through 3999) without dual class shares. Annual returns range between July of 1991 and June of 2001. The sample consists of 2,305 observations on 435 firms. Standard errors are clustered by firm and year according to the procedure in Thompson (2006). T-statistics are reported in parentheses.

	(1)	(2)	(3)	(4)
		Panel A		
	No FC	WW	KZ	Cleary
MC	-5.38	-5.04	-4.60	-5.38
	(-3.48)	(-3.38)	(-3.31)	(-3.61)
G2	-1.46	-1.00	-1.47	-1.54
	(-2.45)	(-1.54)	(-2.57)	(-2.46)
MC*G2	1.23	1.38	1.39	1.28
	(1.58)	(1.87)	(2.26)	(1.73)
FC		6.47	0.29	-0.94
		(3.00)	(0.39)	(-0.87)
MC*FC		-1.54	-2.17	-0.71
		(-2.52)	(-2.78)	(-0.73)
		Panel B		
	No FC	WW	KZ	Cleary
MC	-3.58	-3.38	-2.83	-3.70
	(-14.30)	(-12.38)	(-6.22)	(-20.42)
G3	-2.73	-2.59	-2.82	-2.89
	(-2.01)	(-1.69)	(-2.10)	(-2.26)
MC*G3	0.81	1.74	1.44	1.24
	(1.14)	(3.11)	(2.54)	(6.64)
FC		6.52	0.36	-0.81
		(3.04)	(0.47)	(-0.74)
MC*FC		-1.80	-2.21	-0.86
		(-2.84)	(-2.78)	(-0.84)

Table X: Firm Characteristics By Levels Of The Composite Index. The table reports means of different variables by three groups of the composite index of price impact (CI). The groups include firms with low levels of the index (Low, bottom 25%), medium levels (Med, between 25% and 75%), and high levels (High, top 25%). Three versions of the index are constructed combining the Gompers, Ishii, and Metrick (2003) index of corporate governance (G) and one of three indices of financial constraints (FC). CI for firm i in year t is: $CI_{it} = \hat{\beta}_1 + \hat{\beta}_2 G_{it} + \hat{\gamma}_1 FC_{it} + 0.41$. The coefficients $\hat{\beta}_1$, $\hat{\beta}_2$, and $\hat{\gamma}_1$ refer to equation 3 in the text and are the slopes on MC, MC*G, and MC*FC, respectively, from the models in columns 2, 4, and 6 of Table VIII, depending on which FC index is considered. The financial constraints indices are: the Whited and Wu (2006) index, the Kaplan and Zingales (1997) index, and the Cleary (1999) index. Higher levels of the financial constraints indices denote more financially constrained firms. The variables for which the mean is reported are: the governance index, the financial constraints indices (WW, KZ, Cleary), the funding status over assets (FS, in percent), mandatory contributions over assets (MC, in percent), the annual return starting in July of year t (Ret12), the annual return up to June of year t (Ret_12), the stock beta, market capitalization in June of year t (Size, in $ millions), book-to-market (BM), accruals over assets, firm's age, a dummy for firms whose bonds are rated, the average number of analysts following the firm in the twelve months up to June of year t, capital expenditures over assets, research and development expenses over assets, cash flows over assets (CF), total long term debt over assets (TLTD), total dividends over assets (TDIV), total cash over assets (CASH), Tobin's Q in December of year $t-1$, a dummy that equals one if the firm pays dividends (DIVPOS), the logarithm of total assets (LNTA), the three-digit industry real sales growth (ISG), the firm's real sales growth in year $t-1$ (SG), the ratio of current assets to current liabilities (CURAT), the coverage ratio for interests and preferred dividends (COVER), the ratio of net income to net sales (IMARG), a measure of financial slack (SLACK). See the text and the appendix for the details on variable definitions.

	CI_WW			CI_KZ			CI_Cleary		
	Low	Med	High	Low	Med	High	Low	Med	High
G	6.704	9.863	12.866	6.926	9.913	12.542	6.277	9.931	13.158
WW	0.619	0.556	0.517	0.580	0.558	0.552	0.573	0.559	0.556
KZ	0.729	0.527	0.510	1.443	0.554	-0.262	0.761	0.570	0.392
Cleary	0.838	0.817	0.834	1.083	0.803	0.617	0.963	0.839	0.665
FS	2.360	3.152	3.749	1.969	3.040	4.367	2.750	2.816	4.029
MC	0.075	0.052	0.034	0.080	0.049	0.033	0.062	0.054	0.042
Ret12	12.901	12.755	12.707	13.157	13.346	11.265	12.073	13.253	12.544
Ret_12	10.719	13.415	14.277	12.232	13.425	12.736	11.346	13.363	13.751
Beta	0.966	0.950	1.004	0.983	0.958	0.972	0.909	0.973	1.016
Size	1,130	3,320	5,598	2,489	3,511	3,847	2,965	3,401	3,589
BM	0.815	0.676	0.607	0.865	0.677	0.553	0.771	0.702	0.599
Acc	-0.043	-0.047	-0.048	-0.046	-0.046	-0.045	-0.045	-0.046	-0.045
Age	22.334	34.553	40.126	27.066	33.542	37.395	27.884	33.908	35.837
Rating	0.389	0.650	0.848	0.609	0.651	0.626	0.543	0.653	0.688
Analysts	7.443	12.963	17.654	10.664	12.944	14.394	10.701	13.046	14.202
CAPEX	0.072	0.078	0.079	0.076	0.078	0.075	0.074	0.078	0.077
RD	0.018	0.019	0.022	0.013	0.020	0.025	0.015	0.019	0.024
CF	0.095	0.108	0.113	0.083	0.104	0.133	0.099	0.103	0.118
TLTD	0.380	0.393	0.408	0.524	0.382	0.285	0.429	0.399	0.348
TDIV	0.018	0.025	0.027	0.012	0.023	0.037	0.022	0.024	0.025
CASH	0.083	0.064	0.050	0.043	0.067	0.084	0.069	0.065	0.061
Q	1.501	1.638	1.663	1.405	1.554	1.927	1.569	1.565	1.739
DIVPOS	0.709	0.916	0.989	0.741	0.901	0.986	0.828	0.877	0.947
LNTA	6.380	7.430	8.177	7.257	7.415	7.323	7.183	7.412	7.404
ISG	0.061	0.044	0.032	0.049	0.041	0.050	0.030	0.039	0.072
SG	0.024	0.041	0.059	0.035	0.039	0.053	0.022	0.043	0.057
CURAT	2.172	1.717	1.535	1.679	1.749	1.965	1.906	1.743	1.748
COVER	30.737	20.389	15.610	20.889	21.249	23.792	25.321	21.156	19.520
IMARG	0.038	0.051	0.051	0.028	0.049	0.066	0.039	0.049	0.054
SLACK	1.484	1.087	0.694	0.907	1.025	1.399	1.286	1.069	0.930

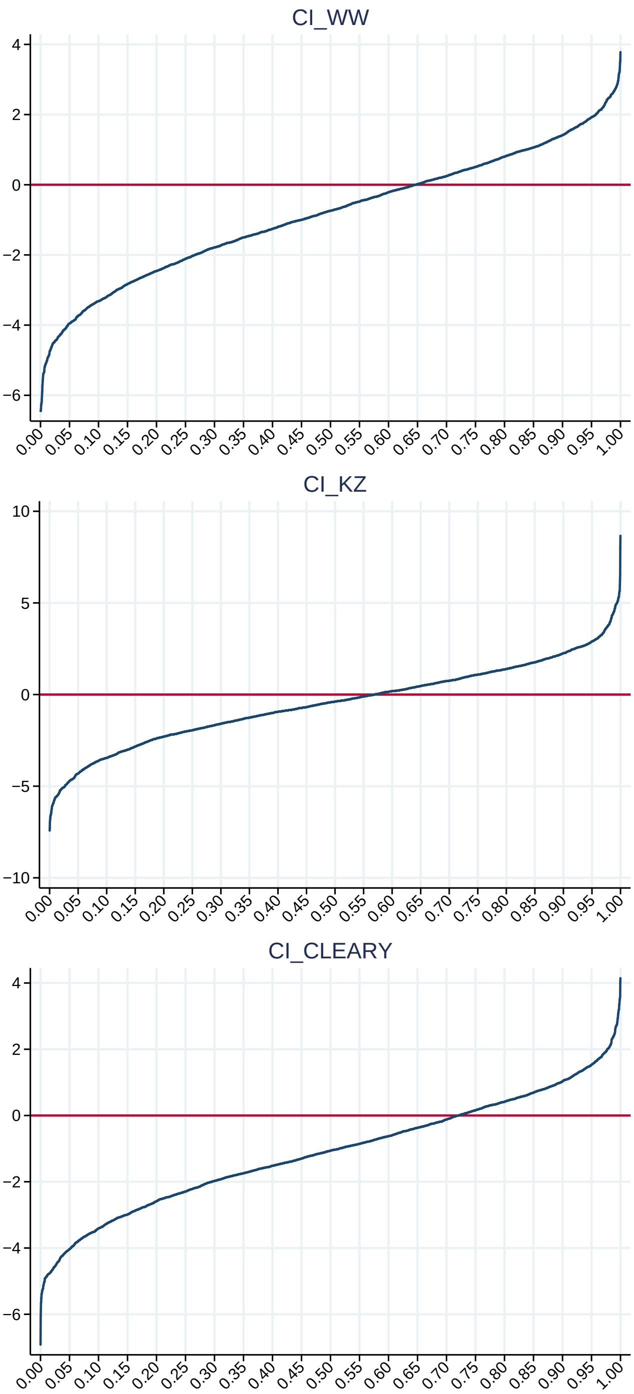

Figure 1: Distribution of the Composite Index. The graphs plot the cumulative distribution of the composite index of price impact (CI). For each level of CI (on the vertical axis) the graphs give the fraction of the firm-year observations for which CI is below that level (horizontal axis). Three versions of CI are constructed combining the Gompers, Ishii, and Metrick (2003) index of corporate governance (G) and one of three indices of financial constraints (FC). CI for firm i in year t is: $CI_{it} = \hat{\beta}_1 + \hat{\beta}_2 G_{it} + \hat{\gamma}_1 FC_{it} + 0.31$. The coefficients $\hat{\beta}_1$, $\hat{\beta}_2$, and $\hat{\gamma}_1$ refer to equation 3 in the text and are the slopes on MC, MC*G, and MC*FC, respectively, from the models in columns 2, 4, and 6 of Table VIII, depending on which FC index is considered. The financial constraints indices are: the Whited and Wu (2006) index, the Kaplan and Zingales (1997) index, and the Cleary (1999) index. Higher levels of the financial constraints indices denote more financially constrained firms. The sample consists of 4,107 observations on 834 firms between 1990 and 2000.