Interest Rates and Investment Redux

Simon Gilchrist[*] Fabio M. Natalucci[†] Egon Zakrajšek[‡]

May 19, 2006

Abstract

The empirical difficulties associated with estimating the effects of changes in interest rates and corporate tax policy on business fixed investment are often blamed on a lack of identification. In this paper, we study the effect of variation in interest rates on investment spending, employing a large new panel data set that links yields on outstanding corporate bonds to the issuer income and balance sheet statements. The bond price data, based on trades in the secondary market, allow us to construct firm-specific measures of the marginal cost of external finance. Our results imply a robust and quantitatively important effect of real interest rates on the firm-level investment decisions. According to our estimates, a one-percentage-point increase in real interest rates is associated with the reduction in the average rate of capital spending between 70 to 130 basis points.

JEL Classification: E22, E44, E62
Keywords: Investment, interest rates, external financing costs

[*]Boston University and NBER.
[†]Federal Reserve Board.
[‡]Corresponding author: Banking Analysis Section, Division of Monetary Affairs, Federal Reserve Board, 20th Street & Constitution Avenue, NW, Washington D.C., 20551. Tel: (202) 728-5864; Email: egon.zakrajsek@frb.gov

1 Introduction

The notion that investment falls when interest rates rise is a theoretically unambiguous relationship that lies at the heart of the monetary transmission mechanism. This negative relationship, however, is surprisingly difficult to document in actual data; see, for example, Abel and Blanchard (1986) and Schaller (2002). Similarly, the effect of taxes on investment is a key relationship that fiscal policy makers rely on when determining the costs and benefits of alternative tax policies. With the exception of Cummins et al. (1994, 1996), who adopt a natural experiments approach using firm-level data, researchers have had a difficult time identifying the relationship between capital formation and corporate tax policy (c.f. Schaller (2002) and Chirinko et al. (1999, 2004)).[1]

The empirical difficulties associated with estimating the effect of interest rates and tax changes on business fixed investment are often blamed on a lack of identification. In particular, at the macroeconomic level, interest rates and taxes are often lowered when investment spending is weak. Thus, endogeneity between policy and the macroeconomy may imply a positive relationship between investment and various components of the user cost of capital.

In this paper, we revisit this apparent and long-standing empirical anomaly. We do so by constructing a new data set that links income and balance sheet information for more than 1,100 large U.S. nonfinancial corporations to interest rates on their publicly-traded debt. Covering the last thirty years, this new data set enables us to evaluate and quantify empirically the relationship between firms' investment decisions and fluctuations in the *firm-specific* marginal financing costs as measured by the changes in secondary market prices of firms' outstanding bonds.

Our results indicate that business fixed investment is highly sensitive—both economically and statistically—to movements in firm-specific real interest rates. The interest-sensitivity of capital formation is robust to the inclusion of various measures of investment opportunities emphasized by frictionless neoclassical models as well as firm-level measures of expected default risk.

The remainder of the paper is organized as follows. In Section 2, we review the existing evidence—at both the macro and micro levels—on the link between financing costs and investment spending. Section 3 describes our new data set and highlights its key feature. Section 4 outlines our empirical methodology and presents our benchmark results. In Section 5, we focus on the 1990–2004 period, a part of our sample characterized by a fully developed market for both investment- and speculative-grade corporate debt. For this sub-sample period, we have also merged our firm-level monthly bond yields with market-based

[1] For extensive surveys, see Auerbach (1983) and Chirinko (1993); see also Hassett and Hubbard (1997) and Devereux et al. (1994).

measures of expected default risk widely used by financial markets participants. In addition, we utilize our interest rate data to construct neoclassical user cost of capital at the firm level, taking into account depreciation, expected capital gains (or losses), movements in the relative price of capital, and tax treatment of investment and capital income.

2 Data Description

Our data set is an unbalanced panel of more than 1,100 publicly-traded firms in the U.S. nonfarm nonfinancial corporate sector covering the period 1973 to 2004. The distinguishing feature of the firms in our sample is that a part of their long-term debt—in many cases, a significant portion—is in form of bonds that are actively traded in the secondary corporate cash market. For these firms, we have linked monthly market prices of their outstanding securities to annual income and balance sheet statements from Compustat. For the last decade and a half of the sample period, we have also linked this data to option-theoretic measures of default risks that are widely used by market participants. We now turn to the construction of our key variables: firm-specific interest rates, key income and balance sheet variables, and measures of expected default risk.

2.1 Bond Yields

We obtained month-end market prices of outstanding long-term corporate bonds from the Lehman/Warga (LW) and Merrill Lynch (ML) databases. These two data sources include prices of nearly all dollar-denominated bonds publicly issued in the U.S. corporate cash market. The ML database is a proprietary data source of *daily* bond prices that starts in 1997. Focused on the most liquid securities, bonds in the ML database must have a remaining term-to-maturity of at least one year, a fixed coupon schedule, and a minimum amount outstanding of $100 million for below investment-grade and $150 million for investment-grade issuers. By contrast, the LW database of *month-end* bond prices has a somewhat broader coverage and is available from 1973 through mid-1998 (see Warga (1991) for details). For securities with market prices in both the LW and LM databases, we spliced the option-adjusted yields at month-end—a component of the bond's yield that is not attributable to embedded options—across the two data sources.

To ensure that we are measuring financing costs of different firms at the same point in their capital structure, we limited our sample to only senior unsecured issues. To covert the monthly nominal bond yields into real terms, we employed a simplifying assumption that the expected inflation in period t is equal to the last period's realized annual core CPI inflation. Specifically, letting i^k_{jt} denote the nominal yield (in percent per annum) on bond k of firm j at the end of month t, we computed the corresponding real yield r^k_{jt} according

Table 1: Summary Statistics of Bond Characteristics

Variable	Mean	Std. Dev.	Min	Median	Max
# of bonds per firm/month	3.28	4.01	1.00	2.00	57.00
Mkt. Value of Issue[a] ($mil.)	266.9	298.1	1.2	197.8	6,771.1
Maturity[b] (years)	13.8	9.4	2.0	10.0	50.0
Effective Duration	6.58	2.89	0.01	6.28	19.54
Composite Rating (S&P)	-	-	D	A3	AAA
Coupon Rate (%)	7.83	2.19	0.00	7.59	17.50
Nominal Yield (%)	8.52	2.89	0.17	8.05	35.31
Real Yield[c] (%)	4.96	2.60	-4.07	4.74	29.99

Panel Dimensions

Obs. $= 374,747$ $\quad N = 6,293$ bonds

Min. Tenure = 1 Median Tenure = 45 Max. Tenure = 302

NOTES: Sample period: Monthly data from January 1973 to December 2004 ($T = 382$).
[a]Market value of the outstanding issue deflated by the CPI.
[b]Maturity at issue date.
[c]Nominal yield less the percent change in the previous month's core CPI from twelve months prior.

to

$$r_{jt}^k = i_{jt}^k - 100 \times \ln\left(\frac{\text{CPI}_{t-1}}{\text{CPI}_{t-13}}\right),$$

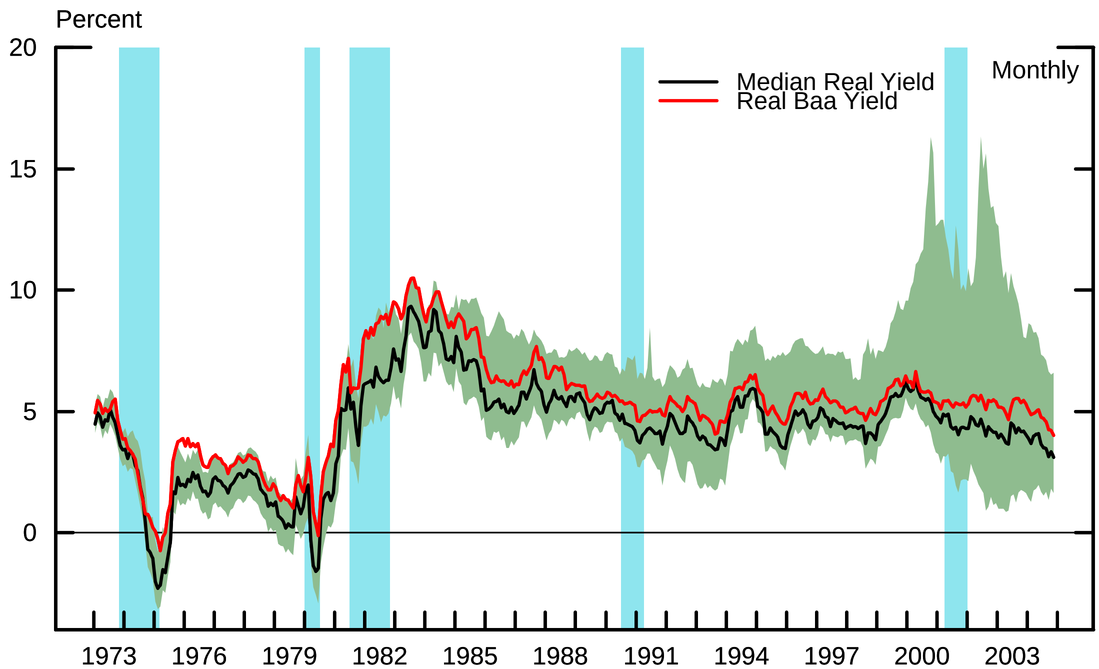
where CPI denotes the level of the Consumer Price Index, excluding its food and energy components.

Table 1 contains summary statistics for the key characteristics of bonds in our sample.[2] Note that a typical firm has only a few bond issues outstanding—the median firm, for example, has two bond issues trading at any given month. Nevertheless, this distribution is highly positively skewed, and some firms can have more than fifty different bond issues trading in the market at a point in time. The distribution of the real market values of these issues is similarly skewed, with the range running from $1.2 million to more than $6.7 billion. Not surprisingly, the maturity of these debt instruments is fairly long, with the average maturity at issue of about 14 years. Because corporate bonds typically generate significant cash flow in the form of regular coupon payments, the effective duration is considerably shorter, with both the average and the median duration of about 7.5 years. Although our sample spans the entire spectrum of credit quality—from "single D" to "triple A"—the median bond/month observation, at "A3," is solidly in the investment-grade category.

Turning to returns, the (nominal) coupon rate on these bonds averaged 7.83 percent

[2]To mitigate the effect of outliers on the sample statistics, we eliminated all observations with real interest rates in excess of 30 percent per annum.

Figure 1: The Evolution of Real Bond Yields



NOTES: This figure depicts the evolution of the cross-sectional distribution of real bond yields in our sample. The solid black line shows the median of the cross-sectional distribution of real yields, while the shaded green band shows a corresponding measure of cross-sectional dispersion, calculated as the difference between the 95th percentile and the 5th percentile of the distribution. The dashed red line shows the real aggregate yield on all Baa-rated corporate bonds. The shaded blue vertical bars denote the NBER-dated recessions.

during our sample period, while the average total nominal return, as measured by the nominal yield, was 8.52 percent per annum. Reflecting the wide range of credit quality, the distribution of nominal yields is fairly wide, with the minimum of 17 basis points and the maximum of more than 35 percent. In real terms, these bonds yielded about 5 percent per annum during our sample period, with the standard deviation of 2.6 percent.

Figure 1 depicts the time-series evolution of the cross-sectional distribution of real yields for the bonds in our sample (see Table 1). For comparison, we also plotted the real yield on all nonfinancial corporate bonds carrying the Moody's Baa credit rating, calculated using the same methodology as in the case of our bond-level data. Several features in Figure 1 are worth noting. First, as evidenced by the closeness of the 95th and 5th percentiles, there is surprisingly little cross-sectional dispersion in real yields until the second half of the 1980s. The narrowness of the distribution before the mid-1980s reflects in large part the fact that the secondary market for corporate debt during this time period was limited largely to investment-grade issues at the upper end of the credit-quality spectrum. Indeed, during this period, a significant majority of real yields in our sample are consistently below the real yield on the Baa-rated corporate bonds, a category of debt that sits at the bottom rung of the investment-grade ladder.

Second, the increase in the cross-sectional dispersion of real interest rates that began in the second half of the 1980s coincided with the deepening of the market for "junk-rated" corporate debt. The drift of the real Baa yield towards the center of the cross-sectional distribution is another piece of evidence pointing to the increased ability of riskier firms to tap the corporate cash market. The amount of cross-sectional heterogeneity in our sample is particularly apparent between 2000 and 2003, a period in which the effects of a cyclical downturn were compounded by a slew of corporate scandals. This combination of the cross-sectional heterogeneity in real financing costs with considerable cyclical fluctuations are factors that should enhance our ability to identify variation in the investment supply curve and thus help us to estimate more precisely the interest sensitivity of investment demand.

As noted above, effective duration varies widely across our sample of bonds. To ensure that differences in our measure of firm-specific financing costs are not influenced by term premiums, we subtracted from each bond yield an estimate of the term premium derived from the Treasury yield curve. Specifically, let d_{jt}^k denote the effective duration of bond k (issued by firm j) on day t and let d^* denote the "target" duration. Our estimate of the time-varying term premium around the target duration d^*, denoted by δ_{kt}, is given by

$$\delta_{kt} = y_t(d^*) - y_t(d_{jt}^k),$$

where $y_t(d)$ denotes the yield on day t on a zero-coupon Treasury security of maturity d. We set our target duration d^* equal to 7 years—around the median duration in our sample—and we used the daily (month-end) estimates Treasury yield curve from Gürkaynak, Sack, and Wright (2006) to compute the term premium δ_{kt}.

Because our income and balance sheet data are available only at an annual frequency, we converted the monthly bond yields to firm-level interest rates in two steps. First, we calculated an average bond yield for firm j in month t by averaging the duration-adjusted yields (both nominal and real) on the firm's outstanding bonds in that month, using market values of bond issues as weights:

$$i_{jt} = \sum_{k=1}^{B_{jt}} w_{jt}^k i_{jt}^k \quad \text{and} \quad r_{jt} = \sum_{k=1}^{B_{jt}} w_{jt}^k r_{jt}^k,$$

where B_{jt} denotes the number of outstanding bond issues of firm j at the end of month t, $0 < w_{jt}^k \leq 1$ is the weight for bond issue k, and i_{jt}^k and r_{jt}^k are the duration-adjusted nominal and real yields on bond k, respectively. To convert these firm-level rates to annual frequency, we then averaged the available monthly yields over the twelve months of the firm's fiscal year. For example, for a firm with fiscal year ending in December, the average interest rate in year t is calculated as an unweighted average of the available monthly yields from January through December of the same year. For a firm with fiscal year ending in, say, June, the average interest rate in year t is calculated as an unweighted average of the available monthly yields from July of year $t-1$ through June of year t.

2.2 Income and Balance Sheet Data

For 1,191 firms in the U.S. nonfarm nonfinancial corporate sector, we linked these firm-specific average market interest rates on long-term unsecured debt to income and balance sheet items from the annual Compustat data files. To ensure comparability with previous empirical work, we follow Gilchrist and Himmelberg (1998) in the construction of the standard variables (e.g., investment rate, sales-to-capital ratio, Tobin's Q, etc.) used in our analysis. Table 2 contains summary statistics for the key variables in our matched annual panel.[3]

Although our sample focuses on firms that have both equity and a portion of their long-term debt traded in capital markets, firm size—measured by sales or market capitalization—varies widely in our sample. Not surprisingly, though, most of the firms in our data set are quite large. The median firm has annual real sales of about $3.4 billion and a real market capitalization of more than $2.6 billion.

Despite the fact that firms in our sample generally have only a few senior unsecured bond issues trading at any given point in time, this publicly-traded debt represents a significant portion of their long-term debt. The ratio of the par value of traded bonds outstanding to the book value of total long-term debt on firms' balance sheet is, on average, almost 0.45, indicating that market prices on these outstanding securities likely provide an accurate gauge of the marginal financing costs. During our sample period, these financing costs averaged—in real terms—about 5.3 percent and were associated with an average annual investment rate (i.e., investment-to-capital ratio) of 21 percent.

Figure 2 compares the dynamics of investment in our sample with those of the U.S. economy as a whole. While our sample includes less than 1,200 firms, these firms tend to be large and, consequently, their investment pattern in the aggregate is broadly similar to the investment dynamics in the National Income and Product Accounts (NIPA). [To be continued.]

[3]To ensure that our results are not driven by a small number of outliers, we eliminated all firm/year observations that exceeded the thresholds specified in Gilchrist and Himmelberg (1998).

Table 2: Summary Statistics for Key Variables

Variable	Mean	Std. Dev.	Min	Median	Max
Sales[a] ($bil.)	8.36	16.78	$<.00$	3.41	245.0
Mkt. Capitalization[b] ($bil.)	8.27	18.98	$<.00$	2.62	297.7
Par Value to L-T Debt[c]	0.44	0.25	$<.00$	0.41	1.00
Real Interest Rate[d] (%)	5.51	3.08	-2.42	5.04	29.92
Investment to Capital[e]	0.21	0.14	$<.00$	0.18	1.00
Sales to Capital[f]	3.66	3.26	0.13	2.81	24.81
Profits to Capital[g]	0.46	0.36	-0.20	0.37	2.50
Tobin's Q[h]	1.49	0.78	0.45	1.26	15.25

Panel Dimensions

Obs. $= 9,993$ $N = 1,131$ firms

Min. Tenure = 1 Median Tenure = 6 Max. Tenure = 32

NOTES: Sample period: Annual data from 1973 to 2004 ($T = 32$). In variable definitions, $x_n(t)$ denotes the Compustat data item n in period t.

[a]The real value of sales in period t: $x_{12}(t)$ deflated by the CPI.

[b]The real market value of common shares outstanding at the end of period t: $x_{25}(t) \times x_{199}(t)$ deflated by the CPI.

[c]The ratio of the par value of all of the firm's traded bonds to the book value of its total long-term debt ($x_9(t)$).

[d]Duration adjusted annual real yield on the firm's outstanding bonds (see text for details).

[e]The ratio of gross investment in period t to net property, plant, and equipment at the end of period $t-1$: $x_{30}(t)/x_8(t-1)$.

[f]The ratio of (net) sales in period t to net property, plant, and equipment at the end of period $t-1$: $x_{12}(t)/x_8(t-1)$.

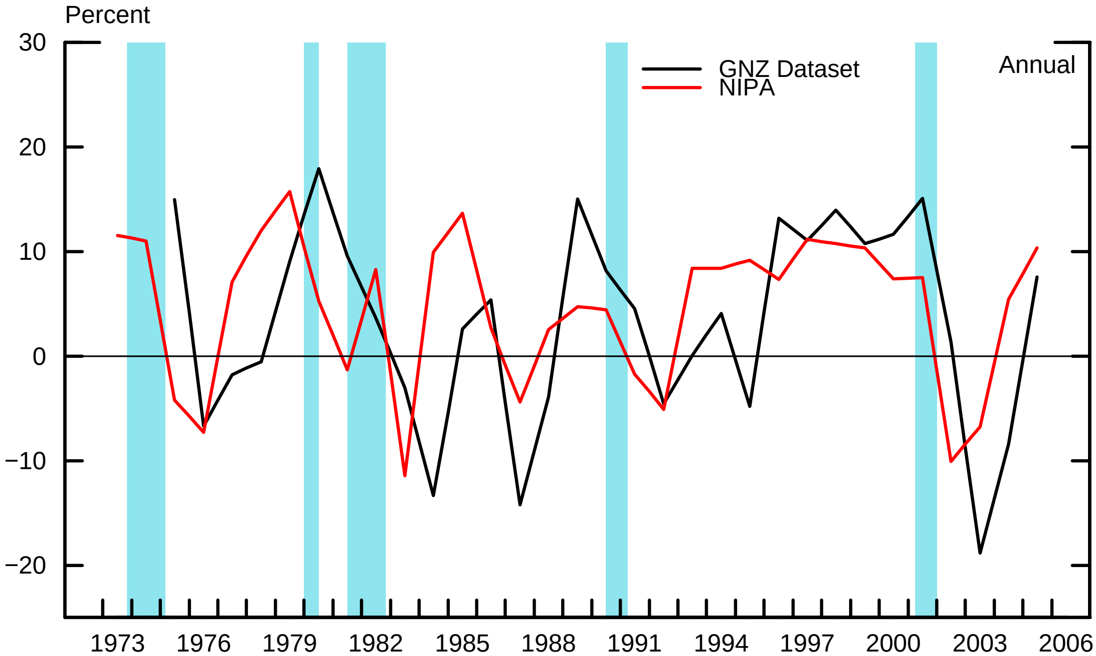
[g]The ratio of operating income (loss) in period t to net property, plant, and equipment at the end of period $t-1$: $x_{13}(t)/x_8(t-1)$.

[h]The ratio of the sum of the market value of common shares outstanding and the book value of total liabilities at the end of period t to the book value of total assets at the end of period t: $[x_{25}(t) \times x_{199}(t) + x_{181}(t)]/x_6(t)$.

2.3 Default Risk

Corporate bond yields are influenced importantly by likelihood of default. Since 1991, we have firm-specific measures of expected default risk, which allows us to control for default risk in our empirical analysis. Our measure of the probability that a firm will default within a certain period of time comes from the Moody's/KMV Corporation (MKMV). The theoretical underpinnings for these probabilities of default are provided by the seminal work of Merton (1973, 1974). According to this option-theoretic approach, the probability that a firm will default on its debt obligations at any point in the future is determined by three major factors: the market value of the firm's assets, the standard deviation of the stochastic process for the market value of assets (i.e., asset volatility), and the firm's leverage. These

Figure 2: The Growth of Business Fixed Investment



NOTES: The solid black line shows the growth rate of the aggregate real capital expenditures for the firms in our sample. The dashed red line shows the growth rate of real business fixed investment measured by the NIPA. Both variables are in chain-weighted (2000=100) dollars. The shaded blue vertical bars denote the NBER-dated recessions.

three factors are combined into a single measure of default risk called *distance to default*, defined as

$$\begin{bmatrix}\text{Distance}\\ \text{to Default}\end{bmatrix} = \frac{\begin{bmatrix}\text{Mkt. Value}\\ \text{of Assets}\end{bmatrix} - \begin{bmatrix}\text{Default}\\ \text{Point}\end{bmatrix}}{\begin{bmatrix}\text{Mkt. Value}\\ \text{of Assets}\end{bmatrix} \times \begin{bmatrix}\text{Asset}\\ \text{Volatility}\end{bmatrix}}.$$

In theory, the default point should equal to the book value of total liabilities, implying that the distance to default compares the net worth of the firm with the size of a one-standard-deviation move in the firm's asset value.[4] The market value of assets and the volatility of assets, however, are not directly observable, so they have to be computed in order to calculate the distance to default. Assuming that the firm's assets are traded, the market value of the firm's equity can be viewed as a call option on the firm's assets with the strike price equal to the current book value of the firm's total debt.[5] Using this insight,

[4]Empirically, however, MKMV has found that most defaults occur when the market value of the firm's assets drops to the value equal to the sum of the firm's current liabilities and one-half of long-term liabilities (i.e., Default Point = Current Liabilities + 0.5 × Long-Term Liabilities), and the default point is calibrated accordingly.

[5]The assumption that all of the firm's assets are traded is clearly inappropriate in most cases. Neverthe-

MKMV "backs out" the market value and the volatility of assets from a proprietary variant of the Black-Scholes-Merton option pricing model, employing the observed book value of liabilities and the market value of equity as inputs; see Crosbie and Bohn (2003) for details.

In the final step, MKMV transforms the distance to default into an expected probability of default—the so-called expected default frequency (EDF)—using an empirical distribution of actual defaults. Specifically, MKMV estimates a mapping relating the likelihood of default over a particular horizon to various levels of distance to default, employing an extensive proprietary database of historical defaults and bankruptcies in the United States.[6] These EDFs are calculated monthly and in our case measure the probability that a firm will default on its debt obligations over the subsequent 12 months. We used EDFs as of the last month of the firm's fiscal year when merging MKMV data to the annual Compustat data files.

It should be noted that MKMV does not disclose how the mapping between the distance to default and the EDF is computed. However, these timely, forward-looking measures of default risk are widely used by financial market participants when assessing credit risk. One clear advantage of EDFs over the traditional measures of default risk based, for example, on credit ratings stems from the fact that the dynamics of EDFs are driven primarily by the movements in equity values. As a result, EDF-based measures of credit risk have the ability to react more rapidly to deterioration in the firm's credit quality as well as to reflect more promptly changes in aggregate economic conditions.

Figure 3 depicts the evolution of the cross-sectional distribution of the expected year-ahead default frequencies for the firms in our sample.

2.4 User Cost of Capital

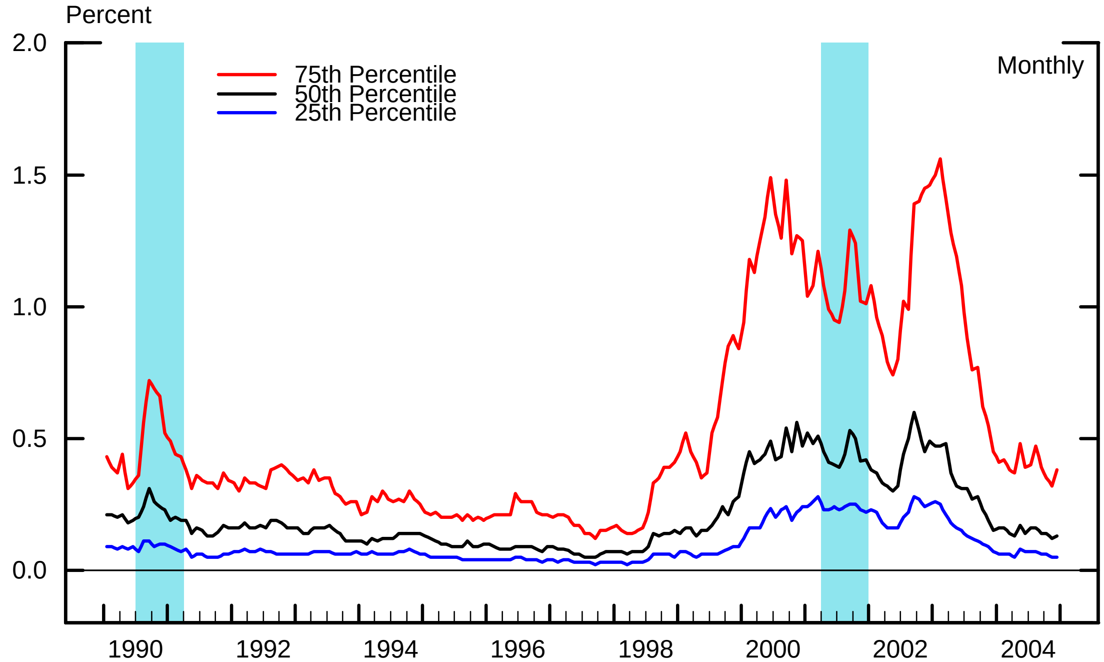
The incentive to purchase physical capital depends not only on the financial costs, but also on the price of investment goods relative to the price of output, the rate at which capital depreciates, any expected gains or losses associated with capital purchases, and the tax treatment of both capital purchases and the capital income. These factors were summarized in the expression for the user cost of capital, derived in the seminal work of Hall and Jorgenson (1967).

We use our firm-level interest rates and industry-level (2-digit SIC) information on the remaining variables to construct the user cost of capital for firm j in period t—denoted by

less, as shown by Ericsson and Reneby (2004), this approach is still valid provided that at least one of the firm's securities (e.g., equity) is traded.

[6] The MKMV's mapping of distances to default to EDFs restricts the probability estimates to the range between 0.02 percent and 20 percent because of sparse data beyond these points.

Figure 3: The Evolution of Year-Ahead Expected Default Frequencies



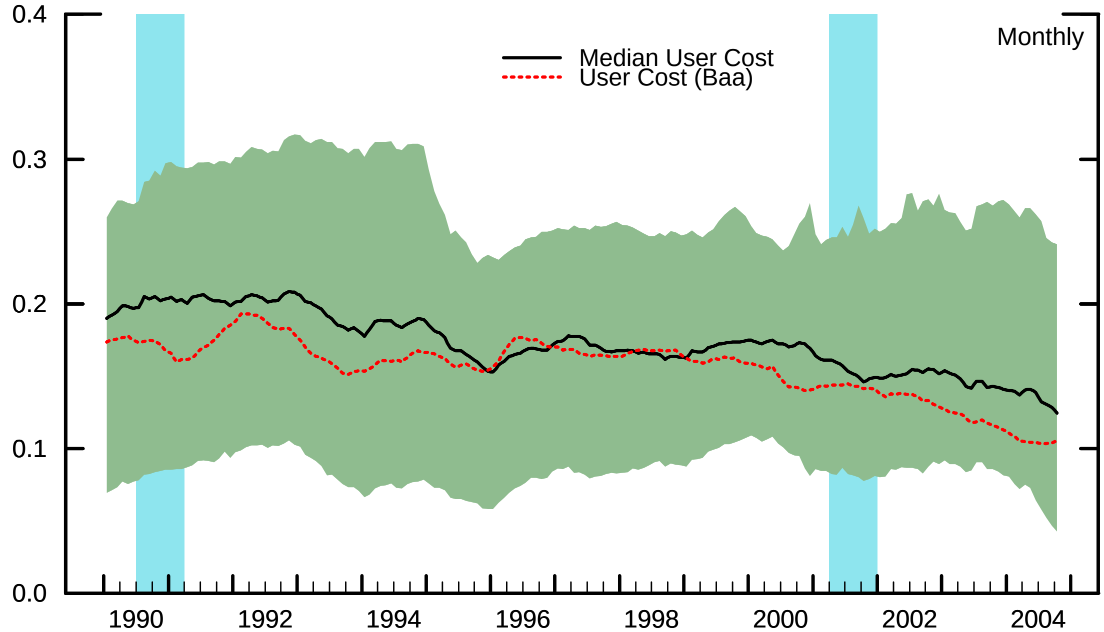
NOTES: This figure depicts the time-series of the 25th, 50th, and 75th percentile of the cross-sectional distribution of year-ahead expected defaults frequencies (EDFs) for the firms in our sample. The shaded blue vertical bars denote the NBER-dated recessions.

C_{jt}^K—according to:

$$C_{jt}^K = \frac{P_{st}^K}{P_{st}^Q}\left[(1-\tau_t)i_{jt} + \delta_{st} - E_t\left(\frac{\Delta P_{s,t+1}^K}{P_{st}^K}\right)\right]\left[\frac{1-\mathrm{ITC}_t - \tau_t z_{st}}{1-\tau_t}\right]. \tag{1}$$

Equation 1 combines the effects of the relative price of investment goods, the rate of return on financial assets, the depreciation rate, the capital gains term, and lastly, the tax considerations. Specifically, P_{st}^K/P_{st}^Q denotes the price of investment goods in industry s relative to the price of output in the same industry; δ_{st} is the time-varying rate of fixed capital depreciation in industry s; $E_t(\Delta P_{s,t+1}^K/P_{st}^K)$ denotes any expected capital gains stemming from the purchase of investment goods; ITC_t is the tax credit rate allowed on investment expenditures; τ_t is the corporate tax rate faced by firm j in period t (assumed to be common across firms); and z_{st} is the present value of the depreciation deduction that can be subtracted from income for tax purposes. Appendix A contains a detailed description of the construction of these industry-level components of the user cost. The component of the user cost that varies across firms is the post-tax nominal interest rate (interest being tax deductible) $(1-\tau_t)i_{jt}$, where i_{jt} is the average duration-adjusted nominal yield on firm j's

Figure 4: The Evolution of the User Cost of Capital



NOTES: This figure depicts the evolution of the cross-sectional distribution of the user cost of capital for the firms in our sample. The solid black line shows the median of the cross-sectional distribution of user cost, while the shaded green band shows a corresponding measure of cross-sectional dispersion, calculated as the difference between the 95th percentile and the 5th percentile of the distribution. The dashed red line shows the aggregate user cost of capital computed using the rate of return on all Baa-rated corporate bonds. The shaded blue vertical bars denote the NBER-dated recessions.

bonds in period t.

Figure 4 shows the evolution of the cross-sectional distribution of the user cost of capital for the firms in our sample. For comparison, we also plotted the aggregate user cost of capital, calculated under the assumption that the required aggregate real rate of return on financial assets is equal to the yield on Baa-rated corporate debt.

3 Empirical Specification of Investment Equation

In this section, we describe our empirical methodology. We regress investment rate on measures of economic fundamentals and measures of financing costs based on the firm-specific information described above. In addition to our measures of financing costs and fundamentals, we control for fixed firm and time effects in our regression analysis. Time effects capture a common investment component owing to macroeconomic factors working through either output or interest rates.[7] Fixed firm effects are included to control for firm-level heterogeneity in the average investment rate of firms. Such heterogeneity may arise either because the mean level of fundamentals differs, or the cost of investing differs across firms in some systematic way not captured by our empirical proxies. Finally, for the sake of robustness, we also allow for serial correlation in the investment process by adding lagged investment rate to our set of explanatory variables.

Our baseline empirical investment equation is

$$\frac{I_{jt}}{K_{j,t-1}} = \beta' Z_{jt} + \theta_1 r_{jt} + \eta_j + \lambda_t + \epsilon_{jt}, \tag{2}$$

where Z_{jt} is a vector of variables that measures future investment opportunities (i.e., fundamentals), r_{jt} is the firm-specific interest rate describe above, η_j is the firm-specific fixed effect, and λ_t is a time dummy. In our baseline case, we assume that ϵ_{jt} is orthogonal to current and past values of Z_{jt} and r_{jt}. We also consider instrumental-variables versions of these regressions, using lagged values as instruments.

In addition to the baseline regression, which measures firm-specific variation in financing costs using the measured interest rate r_{jt}, we also consider investment regressions that explicitly control for expected default probabilities:

$$\frac{I_{jt}}{K_{j,t-1}} = \beta' Z_{jt} + \theta_1 \mathrm{EDF}_{j,t-1} + \theta_2 r_{jt} + \eta_j + \lambda_t + \epsilon_{jt}. \tag{3}$$

Equation 3 decomposes the effect of variation in real bond yields on investment into two distinct terms: the effect of variation in the expected default probability on investment, and the effect of real interest rates on investment, controlling for expected default. Finally, we consider empirical specifications of the investment process that replace the firm-specific real interest rate with its user-cost-of-capital equivalent:

$$\frac{I_{jt}}{K_{j,t-1}} = \beta' Z_{jt} + \theta_1 \mathrm{EDF}_{j,t-1} + \theta_2 C^K_{jt} + \eta_j + \lambda_t + \epsilon_{jt}. \tag{4}$$

Our baseline regressions control for firm-fixed effects by using a standard within-firm transformation. We also wish to capture the persistence of the investment process by

[7] Because our sample of firms covers the entire nonfarm, nonfinancial corporate sector, we also interacted the fixed time effect with a sector-indicator variable based on 2-digit SIC codes to allow aggregate shocks to differ across industrial sectors. All the results reported in the paper are robust to this alternative specification. Moreover, according to the Schwarz Bayesian Information Criterion (BIC), specifications with aggregate time effects are preferred to the specifications with sector-specific time effects, as they involve a considerably smaller number of parameters.

considering regressions of the form:

$$\frac{I_{jt}}{K_{j,t-1}} = \rho \left(\frac{I_{j,t-1}}{K_{j,t-2}} \right) + \beta' Z_{jt} + \theta r_{jt} + \eta_j + \lambda_t + \epsilon_{jt}. \tag{5}$$

As shown by Holtz-Eakin, Newey, and Rosen (1988) and Arellano and Bond (1991), the within-firm transformation to eliminate fixed firm effect leads to inconsistent parameter estimates once lagged dependent variable is included in the regression. Accordingly, we consider a first-differenced transformation of equation 5:

$$\Delta \left(\frac{I_{jt}}{K_{j,t-1}} \right) = \rho \Delta \left(\frac{I_{j,t-1}}{K_{j,t-2}} \right) + \beta' \Delta Z_{jt} + \theta \Delta r_{jt} + \Delta \lambda_t + \Delta \epsilon_{jt}. \tag{6}$$

First differencing, however, induces a moving average error term $\Delta\epsilon_{jt}$, which requires us to use lagged values of Z_{jt}, $I_{jt}/K_{j,t-1}$, and r_{jt} as instruments. We consider similar transformations for equations 3 and 4.

For each of these specifications, we measure investment fundamentals using either the current sales to capital ratio ($S_{jt}/K_{j,t-1}$) or the beginning-of-period t value of Tobin's Q ($Q_{j,t-1}$). Each of these two variables has advantages and disadvantages as a measure of economic fundamentals. Because Tobin's Q is based on stock prices, it is forward looking and thus provides a market-based measure of the discounted present value of future profit opportunities. Indeed, as shown by Hayashi (1982), with quadratic adjustment costs of investment, constant returns to scale in production, and perfect competition in product markets, Tobin's Q is a sufficient statistic for investment. The Q-theory of investment, therefore, implies that the firm-specific interest rates or default probabilities should not have any additional explanatory power in investment regressions; that is, a strict interpretation of the Q-theory implies that $\theta = 0$.

Q-theory, however, may not hold for a variety of reasons. For example, the assumption of both constant returns to scale and perfect competition has been questioned extensively in the empirical literature. In addition, Erickson and Whited (2000) emphasize that measurement error in Tobin's Q may reduce its explanatory power in investment regressions. And finally, as shown by Gilchrist and Himmelberg (1995, 1998), non-market based measures of fundamentals tend to perform at least as well, if not better, than market based measures of fundamentals such as Tobin's Q. In particular, if production is Cobb-Douglas, the sales-to-capital ratio measures the current marginal profitability of capital, even if firms have some degree of market power, which violates the perfect competition, constant returns to scale assumptions required by the standard Q theory. The drawback to using the sales-to-capital ratio as a measure of marginal profitability of capital is that it is not explicitly forward looking. However, assuming that the economic fundamentals follow a martingale process, the current value of the sales-to-capital ratio summarizes the future path of the marginal product of capital and, therefore, provides an accurate measure of expected investment opportunities.

In addition to measuring economic fundamentals using either the sales-to-capital ratio or Tobin's Q, we also consider regressions that include the lagged value of cash flow—defined as operating income relative to capital ($\Pi_{j,t-1}/K_{j,t-2}$)—in our vector of explanatory variables Z_{jt}. Lagged cash flow may be a significant predictor of capital spending either because it proxies for future investment opportunities or because it measures internal funds available for investment purposes. If credit markets are imperfect, firms may find that internal funds are a cheaper source of finance than external funds. In such case, investment demand will respond to fluctuations in cash flow even after controlling for future investment opportunities. Because our bond prices provide a direct measure of the cost of external finance, we expect that, to the extent that cash flow does measure available liquidity through internal funds, the explanatory power of cash flow for investment should diminish once we include our firm-specific measures of financing costs in the investment regression.[8]

4 Results

In this section, we present our empirical results. We begin with the baseline specification described in equation 2. The baseline specification is estimated over the full sample period, that is, 1974–2004. We then consider the effect of including the expected default probability in the investment regression, as well as the effect of replacing the real interest rate with the firm-specific measure of the user cost of capital. These regressions are estimated over the sub-sample period, 1991–2004, a period for which we have available measures of the expected default risk.

Table 3 contains our baseline regression results. Columns 1 and 2 report results of investment regressions using the sales-to-capital ratio as our primary measure of fundamentals, whereas columns 3 and 4 report results that use Tobin's Q as a measure fundamentals. Columns 2 and 4 include cash flow ($\Pi_{j,t-1}/K_{j,t-2}$) as an additional explanatory variable.

According to entries in Table 3, the firm-specific real interest rate is an economically important and statistically significant explanatory variable for investment in all four specifications. Depending on the specification, the coefficients on the interest rate vary between -0.878 and -0.636. Based on these estimates, a one-percentage-point increase in the firm's external financing cost implies a reduction in the rate of investment between 64 to 88 basis

[8] Under the alternative explanation, emphasized by Cooper and Ejarque (2001), cash flow serves as a proxy for investment opportunities because firms have some degree of market power. Under this interpretation, the inclusion of firm-specific interest rates in the investment regression equation should not reduce the explanatory power of cash flow.

Table 3: Investment and Interest Rates
(Level Specification)

	Dependent Variable: I_t/K_{t-1}			
Variable	(1)	(2)	(3)	(4)
S_t/K_{t-1}	0.028	0.021	-	-
	(0.003)	(0.003)		
Q_{t-1}	-	-	0.052	0.036
			(0.007)	(0.006)
Π_{t-1}/K_{t-2}	-	0.111	-	0.131
		(0.012)		(0.012)
r_t	-0.878	-0.699	-0.828	-0.636
	(0.131)	(0.131)	(0.126)	(0.122)
$\Pr > W_\lambda$ [a]	$< .001$	$< .001$	$< .001$	$< .001$
BIC[b]	-0.904	-0.941	-0.864	-0.917
Adj. R^2	0.509	0.527	0.489	0.516

Panel Dimensions

Estimation Period: 1974–2004

Obs. $= 7,954$ $N = 788$ (firms) $\overline{T} = 10.1$ (years)

NOTES: All specifications include fixed firm effects (η_j), fixed time effects (λ_t), and are estimated with OLS. Heteroscedasticity- and autocorrelation-consistent asymptotic standard errors are computed according to Arellano (1987) and are reported in parentheses.

[a] p-value for the robust Wald test of the absence of fixed time effects.

[b] Schwarz Bayesian Information Criterion (smaller is better).

points. The size of the coefficient, as well as its explanatory power, are roughly the same regardless of whether we use sales-to-capital ratio or Tobin's Q as a measure of economic fundamentals, although including cash flow in the regression reduces the size of the interest rate coefficient somewhat. Note that cash flow is a significant explanatory variable in the investment regression estimated in levels.

Table 4 reports the first-differenced specification of the baseline model that also includes the lagged investment rate. The results again indicate that the firm-specific interest rate is an economically and statistically significant predictor of investment spending, with the estimate that range between -1.1 to -1.3. Taking into account the lagged dependent variable, the response of investment to the interest rate is given by $\theta/(1-\rho)$, which for specification in column 1, equals -1.9 (p-value of $< .001$). Thus, controlling for the persistence of the investment process and first differencing roughly doubles the estimated response of investment to interest rates relative to the that reported in Table 3. By contrast, the estimated coeffi-

Table 4: Investment and Interest Rates
(Dynamic First-Difference Specification)

Variable	Dependent Variable: $\Delta(I_t/K_{t-1})$ (1)	(2)	(3)	(4)
$\Delta(I_{t-1}/K_{t-2})$	0.323	0.295	0.328	0.276
	(0.028)	(0.029)	(0.027)	(0.028)
$\Delta(S_t/K_{t-1})$	0.037	0.034	-	-
	(0.006)	(0.005)		
ΔQ_{t-1}	-	-	0.040	0.040
			(0.007)	(0.007)
$\Delta(\Pi_{t-1}/K_{t-2})$	-	0.031	-	0.065
		(0.019)		(0.017)
Δr_t	-1.295	-1.212	-1.093	-1.185
	(0.365)	(0.341)	(0.335)	(0.312)
$\Pr > \lvert m_1 \rvert^a$	$< .001$	$< .001$	$< .001$	$< .001$
$\Pr > \lvert m_2 \rvert^b$	0.518	0.574	0.521	0.523
$\Pr > J_N{}^c$	1.00	1.00	1.00	1.00

Panel Dimensions

Estimation Period: 1975–2004
Obs. $= 6,994$ $N = 777$ (firms) $\overline{T} = 9.0$ (years)

NOTES: All specifications include the first difference of fixed time effects ($\Delta\lambda_t$) and are estimated with GMM using a one-step weighting matrix; see Arellano and Bond (1991). The instrument set includes lags 2 to 4 of $(I_{jt}/K_{j,t-1})$, $(S_{jt}/K_{j,t-1})$, and r_{jt} and lags 1 to 4 of $(\Pi_{j,t-1}/K_{j,t-2})$. Heteroscedasticity- and autocorrelation-consistent asymptotic standard errors are reported in parentheses.

[a] p-value for the test of first-order serial correlation of the first-differenced residuals.

[b] p-value for the test of second-order serial correlation of the first-differenced residuals.

[c] Hansen (1982) test of the overidentifying restrictions. This test uses the minimized objective of the corresponding two-step GMM estimator.

cients on the fundamentals—measured by either sales-to-capital ratio or Tobin's Q—are of the same order of magnitude as in the levels regression. The magnitude of the coefficient on cash flow, however, is halved, and cash flow ceases to be a statistically significant predictor of investment in the specification that includes sales-to-capital ratio as the measure of fundamentals.

We now consider the investment regression specification augmented to include the expected default probability. These regressions are estimated over the sub-sample period 1991 to 2004. Table 5 contains estimates of the levels specification, while Table 6 reports results

Table 5: Investment, Interest Rates, and Default Risk
(Level Specification)

	Dependent Variable: I_t/K_{t-1}					
Variable	(1)	(2)	(3)	(4)	(5)	(6)
S_t/K_{t-1}	0.031	0.031	0.026	-	-	-
	(0.003)	(0.003)	(0.003)			
Q_{t-1}	-	-	-	0.047	0.046	0.036
				(0.007)	(0.006)	(0.006)
EDF_{t-1}	-	-0.446	-0.395	-	-0.415	-0.368
		(0.120)	(0.119)		(0.122)	(0.120)
Π_{t-1}/K_{t-2}	-	-	0.088	-	-	0.100
			(0.014)			(0.014)
r_t	-0.680	-0.436	-0.340	-0.717	-0.492	-0.382
	(0.141)	(0.158)	(0.160)	(0.140)	(0.157)	(0.157)
$\Pr > W_\lambda{}^a$	$< .001$	$< .001$	$< .001$	$< .001$	$< .001$	$< .001$
BIC^b	-0.975	-0.984	-1.014	-0.925	-0.932	-0.969
Adj. R^2	0.575	0.580	0.592	0.553	0.557	0.573

Panel Dimensions

Estimation Period: 1991–2004

Obs. $= 4,817$ $N = 717$ (firms) $\overline{T} = 6.7$ (years)

NOTES: All specifications include fixed firm effects (η_j), fixed time effects (λ_t), and are estimated with OLS. Heteroscedasticity- and autocorrelation-consistent asymptotic standard errors are computed according to Arellano (1987) and are reported in parentheses.

[a] p-value for the robust Wald test of the absence of fixed time effects.

[b] Schwarz Bayesian Information Criterion (smaller is better).

from the first-differenced specification that includes a lagged dependent variable. For comparison purposes, we re-estimated the baseline regression that omits the expected default probability and then report regression results that include the expected default probability.

For the specifications that omit the expected default probability, our estimates for the smaller sub-sample are consistent with those obtained over the full sample period in both the levels and first-differenced specifications. For example, in the levels regression that includes sales-to-capital ratio as the measure of fundamentals, the coefficient estimate on the interest rate is -0.68 for the sub-sample (Column 1, Table 5) and -0.878 for the full sample (Column 1, Table 3). Similarly, in the regression that includes Tobin's Q as the measure of fundamentals, the coefficient estimate on the interest rate in the levels regression is -0.717 in the sub-sample and -0.828 in the full sample. Across the levels specifications, the associated standard error is on the order of 0.13, implying that the interest rate is a statistically significant explanatory variable for investment across all levels specifications and sub-samples. In the first-differenced regression, we also obtain very similar coefficient estimates across the full sample and sub-sample. Not surprisingly, standard errors on the interest rate coefficients estimated over the 1991–2004 sub-sample period are somewhat larger for the first-differenced specifications, a likely reflection of the smaller sample size. Nevertheless, the interest rate remains quantitatively large and statistically significant across sub-samples and in both the levels and first-differenced specifications.

According to the estimation results reported in Table 5, the coefficient on the expected default risk is quantitatively large and statistically significant—a one-percentage-point increase in the expected year-ahead probability of default reduces the investment rate roughly 40 basis points. This result is consistent with standard investment theory, which implies that a higher probability of default lowers the present discounted value of investment opportunities and thus the value to increased investment. Including the expected default probability reduces the coefficient on the interest rate by a factor of one-third relative to the regression that does not include the expected default probability. Nevertheless, the firm-specific real interest rate remains an economically and statistically significant explanatory variable for investment, even after controlling for expected default.

According to the estimates from the first-differenced specification reported in Table 6, the expected year-ahead default probability again has a negative effect on investment, although the coefficient is no longer statistically significant. This is perhaps not too surprising, given that the EDFs, with the exception of specific periods, are relatively constant over the 1991–2004 period. In addition, by first differencing, we eliminate the levels information and increase the noise-to-signal ratio in the data, which has an adverse effect on the precision of our coefficient estimates.

The inclusion of the expected default probability to the first-differenced specification again reduces the size of the coefficient on the interest rate (column 1 vs. 2 and column 4 vs. 5). Nevertheless, the estimated interest rate effect remains economically large—taking into account the dynamics of the investment process, the response of investment to the interest rate is now -1.25 (p-value of 0.07) in the specification that includes sales-to-capital ratio (column 2). The dynamic multiplier from the regression that does not include the expected default probability is -2.21 (p-value of 0.008).

A striking result from the first-differenced specification estimated over the sub-sample period is that the coefficient on cash flow is essentially zero in all cases. This is consistent with our earlier finding in which the cash flow effect was statistically insignificant in the first-differenced specification estimated over the entire sample period and using sales-to-capital ratio as a measure of fundamentals (see column 2 in Table 4). Thus, although cash flow matters for investment in the levels regression, it is not a particularly robust result. In

Table 6: Investment, Interest Rates, and Default Risk
(Dynamic First-Difference Specification)

	Dependent Variable: $\Delta(I_t/K_{t-1})$					
Variable	(1)	(2)	(3)	(4)	(5)	(6)
$\Delta(I_{t-1}/K_{t-2})$	0.410	0.398	0.381	0.394	0.383	0.372
	(0.039)	(0.038)	(0.039)	(0.036)	(0.036)	(0.039)
$\Delta(S_t/K_{t-1})$	0.037	0.038	0.036	-	-	-
	(0.009)	(0.008)	(0.007)			
ΔQ_{t-1}	-	-	-	0.037	0.035	0.037
				(0.007)	(0.007)	(0.007)
ΔEDF_{t-1}	-	-0.207	-0.208	-	-0.296	-0.299
		(0.185)	(0.176)		(0.171)	(0.169)
$\Delta(\Pi_{t-1}/K_{t-2})$	-	-	0.002	-	-	-0.006
			(0.024)			(0.020)
Δr_t	-1.306	-0.753	-0.817	-1.033	-0.724	-0.841
	(0.455)	(0.396)	(0.378)	(0.402)	(0.359)	(0.353)
$\Pr > \lvert m_1 \rvert^a$	$< .001$	$< .001$	$< .001$	$< .001$	$< .001$	$< .001$
$\Pr > \lvert m_2 \rvert^b$	0.645	0.694	0.716	0.539	0.597	0.621
$\Pr > J_N{}^c$	0.957	0.983	0.767	0.978	0.971	0.735

Panel Dimensions

Estimation Period: 1992–2004

Obs. $= 3,977$ $\quad N = 707$ (firms) $\quad \overline{T} = 5.6$ (years)

NOTES: All specifications include the first difference of fixed time effects ($\Delta\lambda_t$) and are estimated with GMM using a one-step weighting matrix; see Arellano and Bond (1991). The instrument set includes lags 2 to 4 of $(I_{jt}/K_{j,t-1})$, $(S_{jt}/K_{j,t-1})$, and r_{jt} and lags 1 to 4 of $(\Pi_{j,t-1}/K_{j,t-2})$ and $(\text{EDF}_{j,t-1})$. Heteroscedasticity- and autocorrelation-consistent asymptotic standard errors are reported in parentheses.

[a] p-value for the test of first-order serial correlation of the first-differenced residuals.

[b] p-value for the test of second-order serial correlation of the first-differenced residuals.

[c] Hansen (1982) test of the overidentifying restrictions. This test uses the minimized objective of the corresponding two-step GMM estimator.

particular, the effect tends to become insignificant once we include firm-specific information regarding the cost of capital. The lack of explanatory power for cash flow is consistent with the notion that the actual real interest rate observed across firms is a better measure of the cost of external finance than a liquidity variable such as cash flow.

We now turn to the regressions that include the user cost of capital rather than the real interest rate as the relevant explanatory variable. We focus on the 1991–2004 sub-sample period and consider specifications with and without the expected default probability as an additional explanatory variable. Table 7 contains the results from the levels regressions,

Table 7: Investment and User Cost of Capital
(Level Specification)

	Dependent Variable: I_t/K_{t-1}					
Variable	(1)	(2)	(3)	(4)	(5)	(6)
S_t/K_{t-1}	0.032	0.031	0.027	-	-	-
	(0.003)	(0.003)	(0.003)			
Q_{t-1}	-	-	-	0.049	0.046	0.036
				(0.007)	(0.007)	(0.006)
EDF_{t-1}	-	-0.544	-0.466	-	-0.541	-0.460
		(0.112)	(0.110)		(0.112)	(0.108)
Π_{t-1}/K_{t-2}	-	-	0.090	-	-	0.102
			(0.014)			(0.013)
C_t^K	-0.270	-0.182	-0.153	-0.251	-0.165	-0.136
	(0.073)	(0.069)	(0.065)	(0.079)	(0.074)	(0.067)
$\Pr > W_\lambda{}^a$	$< .001$	$< .001$	$< .001$	$< .001$	$< .001$	$< .001$
BIC^b	-0.964	-0.981	-1.013	-0.911	-0.927	-0.966
Adj. R^2	0.570	0.578	0.592	0.547	0.555	0.572

Panel Dimensions

Estimation Period: 1991–2004

Obs. $= 4,817$ $N = 717$ (firms) $\overline{T} = 6.7$ (years)

NOTES: All specifications include fixed firm effects (η_j), fixed time effects (λ_t), and are estimated with OLS. Heteroscedasticity- and autocorrelation-consistent asymptotic standard errors are computed according to Arellano (1987) and are reported in parentheses.

[a] p-value for the robust Wald test of the absence of fixed time effects.

[b] Schwarz Bayesian Information Criterion (smaller is better).

while Tables 8 and 9 present results for the first-differenced specification with and without the lagged dependent variable. To allow for the possibility that the user cost of capital is subject to measurement error, Tables 8 and 9 instrument the user cost with the firm-specific (nominal) interest rate rather than the actual user cost.

The use of neoclassical user cost in place of the real interest rate in the levels regression implies a relatively small effect of the user-cost-of-capital on investment—the point estimate is -0.27 for the regression that includes sales-to-capital ratio and -0.25 for the regression that includes Tobin's Q. Adding the expected year-ahead default probability to these two specifications reduces the user-cost coefficients to -0.18 and -0.17, respectively. The coefficients on the expected default probability are also substantially larger in these specifications compared with those that included the real interest rate directly. Nevertheless, the user cost is statistically significant predictor of investment spending in all specifications reported in

Table 8: Investment and (mismeasured) User Cost of Capital
(First-Difference Specification)

	Dependent Variable: $\Delta(I_t/K_{t-1})$					
Variable	(1)	(2)	(3)	(4)	(5)	(6)
$\Delta(S_t/K_{t-1})$	0.037	0.037	0.037	-	-	-
	(0.003)	(0.003)	(0.003)			
ΔQ_{t-1}	-	-	-	0.050	0.049	0.048
				(0.006)	(0.005)	(0.006)
ΔEDF_{t-1}	-	-0.352	-0.332	-	-0.307	-0.304
		(0.104)	(0.105)		(0.102)	(0.103)
$\Delta(\Pi_{t-1}/K_{t-2})$	-	-	0.035	-	-	0.005
			(0.012)			(0.011)
ΔC_t^K	-0.782	-0.656	-0.623	-0.794	-0.689	-0.685
	(0.190)	(0.152)	(0.194)	(0.174)	(0.177)	(0.177)
$\Pr > W_{\Delta\lambda}{}^a$	$< .001$	$< .001$	$< .001$	$< .001$	$< .001$	$< .001$
BIC^b	-1.827	-1.837	-1.841	-1.752	-1.759	-1.756

Panel Dimensions

Estimation Period: 1992–2004

Obs. $= 3,992$ $\quad N = 708$ (firms) $\quad \overline{T} = 5.6$ (years)

Notes: All specifications include the first difference of fixed time effects ($\Delta\lambda_t$) and are estimated using 2SLS. The first difference of the user cost of capital ΔC_{jt}^K is instrumented with a current and a lagged level of the firm-specific nominal interest rate i_{jt}. Heteroscedasticity- and autocorrelation-consistent asymptotic standard errors are computed according to Arellano (1987) and are reported in parentheses.

[a] p-value for the robust Wald test of the absence of fixed time effects.

[b] Schwarz Bayesian Information Criterion (smaller is better).

Table 7.

The results from the first-differenced specifications in Tables 8 and 9 accord much better with the estimated interest rate coefficients reported earlier. In particular, the coefficient on the user cost of capital is between -0.62 and -0.78 in the static specification and between -0.75 and -1.1 in the dynamic specification. Because we are using firm-specific interest rates to instrument the user cost, these results arguably provide a better comparison of the magnitude of the investment response to firm-specific variation in the user cost of capital. Indeed, these estimates of the response of investment to the user cost of capital are very similar to the firm-level user-cost elasticities reported by Cummins, Hasset, and Hubbard (1994). Finally, we again note that in the first-differenced dynamic specification, cash flow has a statistically insignificant effect on investment across all specifications.

Table 9: Investment and (mismeasured) User Cost of Capital
(Dynamic First-Difference Specification)

	Dependent Variable: $\Delta(I_t/K_{t-1})$					
Variable	(1)	(2)	(3)	(4)	(5)	(6)
$\Delta(I_{t-1}/K_{t-2})$	0.420	0.412	0.391	0.383	0.372	0.387
	(0.040)	(0.038)	(0.040)	(0.034)	(0.034)	(0.040)
$\Delta(S_t/K_{t-1})$	0.031	0.033	0.031	-	-	-
	(0.009)	(0.008)	(0.007)			
ΔQ_{t-1}	-	-	-	0.035	0.034	0.038
				(0.007)	(0.007)	(0.007)
ΔEDF_{t-1}	-	-0.271	-0.296	-	-0.180	-0.292
		(0.152)	(0.143)		(0.151)	(0.140)
$\Delta(\Pi_{t-1}/K_{t-2})$	-	-	-0.006	-	-	0.031
			(0.024)			(0.023)
ΔC_t^K	-1.100	-0.753	-0.759	-1.378	-1.199	-1.030
	(0.447)	(0.342)	(0.309)	(0.369)	(0.308)	(0.269)
$\Pr > \lvert m_1 \rvert^a$	$< .001$	$< .001$	$< .001$	$< .001$	$< .001$	$< .001$
$\Pr > \lvert m_2 \rvert^b$	0.580	0.646	0.672	0.461	0.502	0.565
$\Pr > J_N{}^c$	0.970	0.970	0.801	0.989	0.962	0.434

Panel Dimensions

Estimation Period: 1992–2004
Obs. $= 3,977$ $N = 707$ (firms) $\overline{T} = 5.6$ (years)

NOTES: All specifications include the first difference of fixed time effects ($\Delta\lambda_t$) and are estimated with GMM using a one-step weighting matrix; see Arellano and Bond (1991). The instrument set includes lags 2 to 4 of ($I_{jt}/K_{j,t-1}$), ($S_{jt}/K_{j,t-1}$), and $i_{j,t}$ and lags 1 to 4 of ($\Pi_{j,t-1}/K_{j,t-2}$) and ($\text{EDF}_{j,t-1}$). Heteroscedasticity- and autocorrelation-consistent asymptotic standard errors are reported in parentheses.

[a] p-value for the test of first-order serial correlation of the first-differenced residuals.

[b] p-value for the test of second-order serial correlation of the first-differenced residuals.

[c] Hansen (1982) test of the overidentifying restrictions. This test uses the minimized objective of the corresponding two-step GMM estimator.

5 Conclusion

In this paper, we exploit a newly available data set linking firm-specific bond prices to balance sheet and income statement data in order to study the effect of variation in interest rates on investment spending. The bond price data obtained from secondary markets allow us to construct firm-specific measures of the marginal cost of external finance. In addition, this data allow us to measure expected default probabilities at the firm-level.

In contrast to macroeconomic results that find little, if any, systematic relationship between interest rates and capital spending, our estimation results imply a robust and quantitatively important effect of interest rates on investment at the firm-level. According to our estimates, a one-percentage-point rise in real interest rates is associated with the reduction in the average rate of investment somewhere between 70 to 130 basis points. These results imply that investment is highly responsive to changes in interest rates. Consistent with previous studies documented in Cummins, Hasset and Hubbard (1994), our results also imply a strong link between the user cost of capital and investment spending at the firm level. These findings have important implications for the conduct of both monetary and fiscal policy.

References

Abel, A. B., and O. J. Blanchard (1986): "The Present Value of Profits and Cyclical Movements in Investment," *Econometrica*, 54, 249–273.

Arellano, M. (1987): "Computing Robust Standard Errors for Within-Group Estimators," *Oxford Bulletin of Economics and Statistics*, 49, 431–434.

Arellano, M., and S. Bond (1991): "Some Tests of Specification for Panel Data: Monte Carlo Evidence and an Application to Employment Equations," *Review of Economic Studies*, 58, 277–297.

Auerbach, A. J. (1983): "Taxation, Corporate Financial Policy, and the Cost of Capital," *Journal of Economic Literature*, 21, 905–940.

Chirinko, R. S. (1993): "Business Fixed Investment Spending: A Critical Survey of Modeling Strategies, Empirical Results, and Policy Implications," *Journal of Economic Literature*, 31, 1875–1911.

Chirinko, R. S., S. M. Fazzari, and A. P. Meyer (1999): "How Responsive is Business Capital to User Cost? An Exploration with Micro Data," *Journal of Public Economics*, 74, 53–80.

——— (2004): "That Elusive Elasticity: A Long-Panel Approach to Estimating the Capital-Labor Substitution Elasticity," Working Paper Washington University in St. Louis.

Cooper, R., and J. ao M. Ejarque (2001): "Exhuming Q: Market Power vs. Capital Market Imperfections," NBER Working Paper No. 8182.

Crosbie, P. J., and J. R. Bohn (2003): "Modeling Default Risk," Research Report, Moody's|K·M·V Corporation.

Cummins, J. G., K. A. Hassett, and R. G. Hubbard (1994): "A Reconsideration of Investment Behavior Using Tax Reforms as Natural Experiments," *Brookings Papers on Economic Activity*, 2, 1–59.

——— (1996): "Tax Reforms and Investment: A Cross-Country Comparison," *Journal of Public Economics*, 62, 237–273.

Devereux, M. M., M. J. Keen, and F. Schiantarelli (1994): "Corporate Tax, Investment, and the Role of Tax Asymmetries," *Journal of Public Economics*, 53, 395–418.

Erickson, T., and T. M. Whited (2000): "Measurement Error and the Relationship between Investment and q," *Journal of Political Economy*, 108, 1027–1057.

Ericsson, J., and J. Reneby (2004): "A Note on Contingent Claims Pricing with Non-Traded Assets," *Finance Letters*, 2, No. 3.

Gilchrist, S., and C. P. Himmelberg (1995): "The Role of Cash Flow in Reduced-Form Investment Equations," *Journal of Monetary Economics*, 36, 541–572.

——— (1998): "Investment: Fundamentals and Finance," in *NBER Macro Annual*, ed. by B. S. Bernanke, and J. J. Rotemberg, pp. 223–274. The MIT Press, Cambridge.

Gürkaynak, R. S., B. Sack, and J. H. Wright (2006): "The U.S. Treasury Yield Curve: 1961 to Present," Mimeo, Federal Reserve Board.

Hall, R. E., and D. W. Jorgenson (1967): "Tax Policy and Investment Behavior," *American Economic Review*, 57, 319–414.

Hansen, L. P. (1982): "Large Sample Properties of Generalized Method of Moment Estimators," *Econometrica*, 50, 1029–1054.

Hassett, K. A., and R. G. Hubbard (1997): "Tax Policy and Investment," in *Fiscal Policy: Lessons from Economic Research*, ed. by A. J. Auerbach, pp. 339–385. MIT Press, Cambridge.

Hayashi, F. (1982): "Tobin's Marginal Q and Average Q: A Neoclassical Interpretation," *Econometrica*, 50, 215–224.

Holtz-Eakin, D., W. K. Newey, and H. S. Rosen (1988): "Estimating Vector Autoregression With Panel Data," *Econometrica*, 56, 1371–1395.

Merton, R. C. (1973): "A Rational Theory of Option Pricing," *Bell Journal of Economics and Management Science*, 4, 141–183.

——— (1974): "On the Pricing of Corporate Debt: The Risk Structure of Interest Rates," *Journal of Finance*, 29, 449–470.

Schaller, H. (2002): "Estimating the Long-Run User Cost Elasticity," Working Paper No. 02-31, Dept. of Economic MIT.

Warga, A. D. (1991): "A Fixed Income Database," Mimeo, University of Houston.