# Interest Rates and Investment Redux

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**Abstract**

The well-documented 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 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—enable us to construct a firm-specific measure of the user cost of capital based on the marginal cost of external finance as determined in the market for long-term corporate debt. Our results imply a robust and quantitatively important effect of the user cost of capital on the firm-level investment decisions. According to our estimates, a one-percentage-point increase in the user cost of capital is associated with the reduction in the average rate of capital spending between 100 to 125 basis points.

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# 1 Introduction

The notion that business spending on fixed capital falls when interest rates rise is a theoretically unambiguous relationship that lies at the heart of the monetary transmission mechanism. Nevertheless, the presence of a robust negative relationship between investment expenditures and real interest rates—or the user cost of capital more generally—has been surprisingly difficult to document in actual data (e.g., Abel and Blanchard [1986] and Schaller [2002]). Similarly, the magnitude of the response of investment to changes in corporate tax policies is a key parameter that fiscal policy makers rely on when determining the costs and benefits of altering the tax code. With the exception of Cummins, Hassett, and Hubbard [1994, 1996], whose methodology utilizes firm-level variation in investment expenditures within a context of a "natural" experiment, researchers have had a difficult time identifying the relationship between capital formation and changes in corporate tax policy (e.g., Schaller [2002] and Chirinko, Fazzari, and Meyer [1999, 2004]).[1]

The empirical difficulties associated with estimating the effects of changes in interest rates and corporate tax policies on business fixed investment are often blamed on a lack of identification. At the macroeconomic level in particular, long-term interest rates—through monetary policy actions—and corporate tax obligations—through investment tax credits or partial expensing allowances—are often lowered when investment spending is weak.[2] The endogeneity between both monetary and fiscal policy actions and the macroeconomy may imply a positive relationship between investment expenditures and 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 900 large U.S. nonfinancial corporations to interest rates on their publicly-traded debt. Covering the last three decades, this new data set enables us to evaluate and to quantify empirically the relationship between firms' investment decisions and fluctuations in the *firm-specific* user cost of capital based on marginal financing costs as measured by the changes in secondary market prices of firms' outstanding bonds.

Our results indicate that investment expenditures are highly sensitive—both economically and statistically—to movements in firm-specific user cost of capital. The sensitivity of capital formation to changes in the user cost is robust to the inclusion of various measures of investment opportunities emphasized by frictionless neoclassical models.

[1]For extensive surveys of this topic, see Auerbach [1983] and Chirinko [1993]; see also Hassett and Hubbard [1997] and Devereux, Keen, and Schiantarelli [1994].

[2]Partial expensing allowances permit firms to deduct a portion of their newly purchased capital goods from their taxable income. In that sense, both an investment tax credit (ITC) and an expensing allowance raise the firm's after-tax income when the firm purchases capital goods. The two tax policies, however, differ in that under partial expensing, the firm is not allowed to claim any future depreciation allowances for its expensed capital, whereas under an ITC, such a restriction is partly or wholly absent.

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 subsample period, we have also merged our firm-level monthly bond yields with market-based 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 926 publicly-traded firms in the U.S. nonfarm nonfinancial corporate sector covering the period 1973 to 2004. The distinguishing feature of these firms 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. We now turn to the construction of our key variables: firm-specific interest rates and the associated user cost of capital and key income and balance sheet variables.

### 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 for a significant fraction of 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 two years, 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).

To ensure that we are measuring long-term financing costs of different firms at the same point in their capital structure, we limited our sample to only senior unsecured issues. For the securities carrying the senior unsecured rating and with market prices in both the

Table 1: Summary Statistics of Bond Characteristics

| Variable | Mean | StdDev | Min | Median | Max |
|---|---|---|---|---|---|
| # of bonds per firm/month | 3.39 | 4.16 | 1.00 | 2.00 | 57.00 |
| Mkt. Value of Issue[a] ($mil.) | 285.3 | 322.5 | 1.21 | 210.5 | 6,771.1 |
| Maturity at Issue (years) | 14.1 | 9.5 | 2.0 | 10.0 | 50.0 |
| Duration (years) | 6.41 | 2.91 | 0.01 | 6.03 | 29.5 |
| S&P Credit Rating | - | - | D | A3 | AAA |
| Coupon Rate (pct) | 7.67 | 2.13 | 0.00 | 7.42 | 16.63 |
| Nominal Yield (pct) | 8.00 | 2.44 | 1.39 | 7.67 | 24.06 |
| Real Yield[b] (pct) | 4.83 | 1.81 | -3.47 | 4.71 | 15.27 |
| Credit Spread[c] (bps) | 149 | 135 | 0 | 105 | 1000 |

*Panel Dimensions*

Obs. $= 316,984$ $N = 5,800$ bonds

Min. Tenure $= 1$ Median Tenure $= 45$ Max. Tenure $= 229$

NOTES: Sample period: Monthly data from January 1973 to December 2005. Sample statistics are based on trimmed data (see text for details).

[a]Market value of the outstanding issue deflated by the CPI.

[b]Nominal yield less the percent change in previous month's core CPI from twelve months prior.

[c]Measured relative to comparable maturity Treasury yield (see text for details).

LW and LM databases, we spliced the option-adjusted effective yields at month-end—a component of the bond's yield that is not attributable to embedded options—across the two data sources. To calculate the credit spreads at each point in time, we matched the yield on *each* individual security issued by the firm to the estimated yield on the Treasury coupon security of the same maturity. The month-end Treasury yields were taken from the daily estimates of the U.S. Treasury yield curve reported in Gürkaynak, Sack, and Wright [2006]. To mitigate the effect of outliers on our analysis, we eliminated all observations with negative credit spreads and with spreads greater than 1,000 basis points. This selection criterion yielded a sample of 5,800 individual securities issued by the 926 firms in our sample during the 1973–2005 period.

Table 1 contains summary statistics for the key characteristics of bonds in our sample. Note that a typical firm has only a few senior unsecured issues outstanding at any point in time—the median firm, for example, has two such issues trading at any given month. Nevertheless, this distribution is highly positively skewed, and some firms can have more than fifty different senior unsecured 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 7.67 percent during our sample period, while the average total nominal return, as measured by the nominal effective yield, was 8.00 percent per annum. Reflecting the wide range of credit quality, the distribution of nominal yields is quite wide, with the minimum of about 1.40 percent and the maximum of more than 24 percent. In real terms, these bonds yielded 4.8 percent per annum, on average, during our sample period, with the standard deviation of 1.81 percent.[3] Relative to Treasuries, an average bond in our sample generated a return of about 150 basis points above the comparable-maturity risk-free rate, with the standard deviation of 135 basis points.

Figure 1 depicts the time-series evolution of the cross-sectional distribution of nominal yields for the bonds in our sample. For comparison, we also plotted the nominal yield on all corporate bonds carrying the Moody's Baa credit rating. 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 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 yields in our sample are consistently below the yield on the Baa-rated corporate bonds, a category of debt that sits at the bottom rung of the investment-grade ladder.

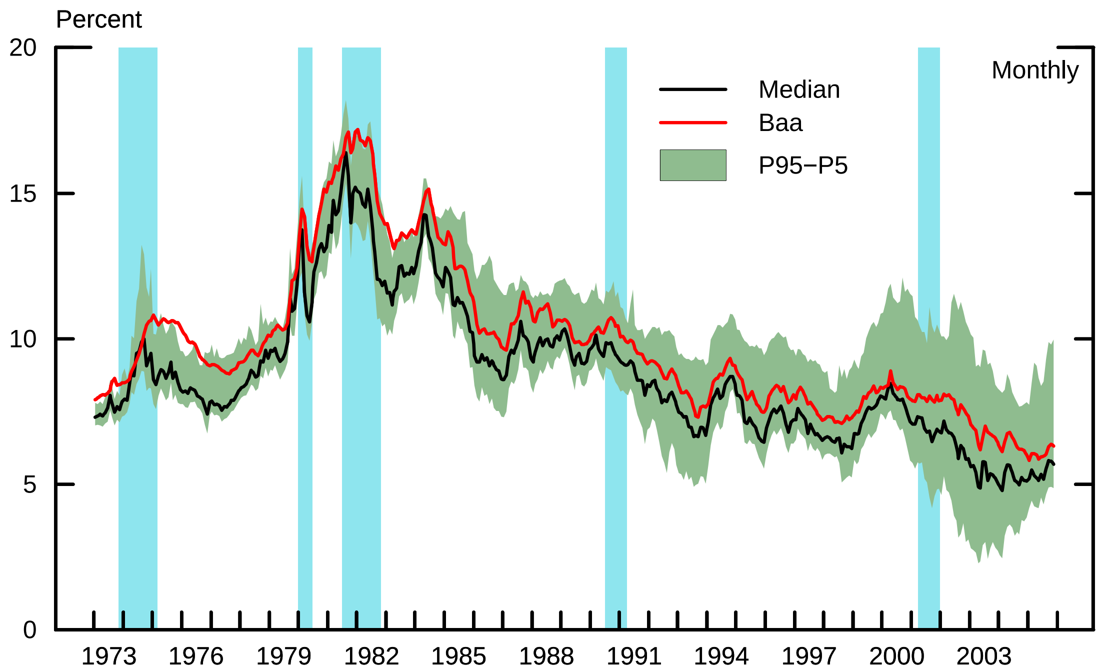
Second, the increase in the cross-sectional dispersion of corporate interest rates that began in the second half of the 1980s coincided with the deepening of the market for "junk-rated" corporate debt. The drift of the aggregate Baa yield towards the center of the cross-sectional distribution is another piece of evidence pointing to the increased ability of riskier firms to tap the corporate cash market. The amount of cross-sectional heterogeneity

[3]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 $i_{jt}^k$ 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 $r_{jt}^k$ according to

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

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

Figure 1: The Evolution of Corporate Bond Yields




NOTES: This figure depicts the evolution of the cross-sectional distribution of nominal bond yields in our sample. The solid black line shows the market-value-weighted median of the cross-sectional distribution of yields, while the shaded green band shows a corresponding measure of cross-sectional dispersion, calculated as the difference between the market-value-weighted 95th percentile (P95) and the market-value-weighted 5th percentile (P5) of the distribution. The dashed red line shows the aggregate yield on all Baa-rated corporate bonds. The shaded blue vertical bars denote the NBER-dated recessions.

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 external 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.

## 2.2 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-specific long-term interest rates, along with sectoral information on the remaining variables, to construct the user cost of capital for firm $j$ in period $t$—denoted by $C^K_{jt}$—according to:

$$C^K_{jt} = \frac{P^K_{st}}{P^Q_{st}} \left[ (1-\tau_t) i_{jt} + \delta_{st} - E_t \left( \frac{\Delta P^K_{s,t+1}}{P^K_{st}} \right) \right] \left[ \frac{1 - \mathrm{ITC}_t - \tau_t z_t}{1-\tau_t} \right]. \tag{1}$$

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, $P^K_{st}/P^Q_{st}$ denotes the price of investment goods in sector $s$ relative to the price of output in the same sector; $\delta_{st}$ is the time-varying rate of fixed capital depreciation in sector $s$; and $E_t(\Delta P^K_{s,t+1}/P^K_{st})$ denotes any expected capital gains (or losses )stemming from the purchase of investment goods.[4] The last term in equation 1 captures tax considerations—assumed to be common to all firms—associated with the purchase of physical capital. That is, $\mathrm{ITC}_t$ is the tax credit rate allowed on investment expenditures; $\tau_t$ is the corporate tax rate faced by firm $j$ in period $t$; and $z_t$ is the present value of the depreciation deduction that can be subtracted from income for tax purposes.[5]

The key component of the user cost that varies across firms is the post-tax nominal interest rate (interest being tax deductible) $(1-\tau_t)i_{jt}$. As discussed above, however, effective duration varies widely across our sample of bonds (see Table 1). To ensure that differences in our firm-specific measure of the user cost 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 $d^k_{jt}$ 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 $\delta_{kt}$, is given by

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

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 of the zero-coupon Treasury yield curve from Gürkaynak, Sack, and Wright [2006] to compute the term premium $\delta_{kt}$.

Because our income and balance sheet data are available only at an annual frequency,

[4]The 15 sectors (based on 2-digit NAICS) are: mining, utilities, construction, nondurable goods manufacturing, durable goods manufacturing, wholesale trade, retail trade, transportation and warehousing, information technology, professional services, administrative services, healthcare, arts and entertainment, accommodation and food services, and other services. In terms of sectoral composition, our sample—measured by sales or capital expenditures—is dominated by manufacturing and information technology firms: firms engaged in nondurable goods manufacturing account for about 30 percent of all investment spending; firms engaged in durable goods manufacturing account for 25 percent; and information technology firms account for about 19 percent of all investment during our sample period.

[5]The tax parameters were taken from the Federal Reserve's FRB/US model of the U.S. economy.

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 nominal yield on the firm's outstanding bonds in that month, using market values of bond issues as weights:

$$i_{jt} = \sum_{k=1}^{B_{jt}} w_{jt}^k i_{jt}^k,$$

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 $i_{jt}^k$ is the duration-adjusted nominal yield on bond $k$. 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 $t$.

### 2.3 Income and Balance Sheet Data

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.[6]

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 \$4.5 billion and a real market capitalization of more than \$1.7 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, 0.35, indicating that market prices on these outstanding securities likely provide an accurate gauge of the marginal financing costs.

[6]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 of Income and Balance Sheet Variables

| Variable | Mean | StdDev | Min | Median | Max |
|---|---|---|---|---|---|
| Sales[a] ($bil.) | 10.77 | 22.01 | < .01 | 4.47 | 253.6 |
| Mkt. Capitalization[b] ($bil.) | 5.40 | 11.85 | < .01 | 1.77 | 172.5 |
| Par Value to L-T Debt[c] | 0.35 | 0.23 | < .01 | 0.31 | 1.00 |
| Investment to Capital[d] | 0.18 | 0.11 | 0.01 | 0.16 | 1.00 |
| Sales to Capital[e] | 4.07 | 3.57 | 0.11 | 2.99 | 24.8 |
| Profits to Capital[f] | 0.53 | 0.43 | -0.49 | 0.41 | 2.97 |
| Tobin's $Q$[g] | 1.53 | 0.80 | 0.44 | 1.30 | 15.3 |

*Panel Dimensions*

Obs. $= 8,076$ $N = 926$ firms

Min. Tenure = 1 Median Tenure = 6 Max. Tenure = 33

NOTES: Sample period: Annual data from 1973 to 2005. Sample statistics are based on trimmed data, and real (i.e., inflation-adjusted) variables are expressed in 2000 dollars (see text for details). In variable definitions, $x_n(t)$ denotes the Compustat data item $n$ in period $t$.

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

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

[c]The ratio of the par value of all of the firm's senior unsecured bonds from the LW/ML database to the book value of its total long-term debt ($x_9(t)$).

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

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

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

[g]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 $t$: $[x_{25}(t) \times x_{199}(t) + x_{181}(t)]/x_6(t)$.

# 3 Empirical Specification of Investment Equation

In this section, we describe our empirical methodology. We regress investment on measures of economic fundamentals and our measure of the user cost of capital calculated using firm-specific long-term interest rates. In addition to our measures of the user cost and investment fundamentals, we control for firm and time fixed effects in our regression analysis. Fixed time effects capture a common investment component owing to macroeconomic influences working through either output or interest rates, while fixed firm effects are included to control for heterogeneity in the average investment rate across firms. Such heterogeneity may arise either because the average level of fundamentals differs, or because 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 including lagged investment rate among the explanatory variables.

Our baseline empirical investment equation is given by

$$\frac{I_{jt}}{K_{j,t-1}} = \beta Z_{jt} + \theta C^K_{jt} + \eta_j + \lambda_t + \epsilon_{jt}, \tag{2}$$

where $Z_{jt}$ is a variable that measures future investment opportunities (i.e., economic fundamentals), $C^K_{jt}$ is the firm-specific user cost of capital, $\eta_j$ is the firm-specific fixed effect, and $\lambda_t$ is a time dummy. In our baseline case, we assume that $\epsilon_{jt}$ is orthogonal to current and past values of $Z_{jt}$ and $C^K_{jt}$.

Our baseline regressions control for fixed firm effects by using a standard within-firm transformation. We also wish to control for lagged dependent variables by considering a dynamic specification of the form:

$$\frac{I_{jt}}{K_{j,t-1}} = \rho \frac{I_{j,t-1}}{K_{j,t-2}} + \beta Z_{jt} + \theta C^K_{jt} + \eta_j + \lambda_t + \epsilon_{jt}. \tag{3}$$

With a lagged dependent variable on the right-hand side, the within-firm regression does not provide consistent parameter estimates. We therefore consider a forward-mean-differenced transformation of equation 3:

$$\Delta_t^{T_j}\left(\frac{I_{jt}}{K_{j,t-1}}\right) = \rho \Delta_t^{T_j}\left(\frac{I_{j,t-1}}{K_{j,t-2}}\right) + \beta \Delta_t^{T_j}(Z_{jt}) + \theta \Delta_t^{T_j}\left(C^K_{jt}\right) + \Delta_t^{T_j}(\lambda_t) + \Delta_t^{T_j}(\epsilon_{jt}), \tag{4}$$

where $\Delta_t^{T_j}$ denotes the forward mean-differencing operator

$$\Delta_t^{T_j}(X_{it}) \equiv X_{it} - \left[\frac{1}{T_j - s}\right] \sum_{s=t+1}^{T_j} X_{is}.$$

This transformation induces a moving-average component into the original error term

$$\Delta_t^{T_j}(\epsilon_{jt}) = \epsilon_{jt} - \left[\frac{1}{T_j - s}\right] \sum_{s=t+1}^{T_j} \epsilon_{js},$$

which nevertheless preserves the validity of instruments in the sense that if $E[\epsilon_{jt} X_{it} \mid \eta_j, \lambda_t] = 0$, then $E[\Delta_t^{T_j}(\epsilon_{jt})\, X_{it} \mid \eta_j, \lambda_t] = 0$. Hence, assuming that for $s \geq 0$

$$E[\epsilon_{jt} Z_{j,t-s} \mid \eta_j, \lambda_t] = E[\epsilon_{jt} C^K_{j,t-s} \mid \eta_j, \lambda_t] = E\left[\epsilon_{jt} \frac{I_{j,t-1-s}}{K_{j,t-2-s}} \mid \eta_j, \lambda_t\right] = 0,$$

lagged values of $I_{jt}/K_{j,t-1}$, along with current and lagged values of $Z_{jt}$ and $C^K_{jt}$, are valid instruments in the presence of the transformed error term $\Delta_t^{T_j}(\epsilon_{jt})$.

In both the static and dynamic specifications, we measure investment fundamentals using either the current sales-to-capital ratio ($S_{jt}/K_{j,t-1}$) or the operating-income-to-capital ratio ($\Pi_{jt}/K_{j,t-1}$). Following Gilchrist and Himmelberg [1998], we construct a measure of the marginal product of capital for firm $j$ at time $t$ as

$$\mathrm{MPK}^{S}_{jt} = \phi_s \left( \frac{S_{jt}}{K_{j,t-1}} \right);$$

or

$$\mathrm{MPK}^{\Pi}_{jt} = \psi_s \left( \frac{\Pi_{jt}}{K_{jt-1}} \right),$$

where $\phi_s$ and $\psi_s$ are appropriately defined scaling factors that are specific to the sector $s$ for firm $j$. The scaling factors $\phi_s$ and $\psi_s$ capture the fact that sales-to-capital and operating-income-to-capital ratios vary substantially across sectors, while in equilibrium, the return on capital should be equalized across sectors. We then set $Z_{jt}$—our measure of economic fundamentals—equal to both measures of the marginal product of capital.

We also wish to consider a log-log specification of the relationship between investment, the user cost of capital, and fundamentals. Taking logs of $\mathrm{MPK}^{S}_{jt}$ is straightforward, but because operating income may be negative, we use

$$Z_{jt} = \log(0.5 + \mathrm{MPK}^{\Pi}_{jt})$$

to measure fundamentals for the log-log specification when using operating income as the measure of investment opportunities. One drawback of both measures is that they are not explicitly forward looking. However, under the assumption that economic fundamentals approximately follow an AR(1) process, the current value of the marginal product of capital summarizes its future path and may, therefore, provide a reasonable measure of future investment opportunities.

### 3.1 Liquidity Shocks as Instruments

An important concern when estimating equation 3 is the potential endogeneity between investment and the firm-specific interest rate that is used to construct the user cost of capital. At the macroeconomic level, endogeneity between interest rates and investment is likely bias the estimated coefficient on the interest rate towards zero, because long-term interest rates typically fall during economic downturns and when the investment fundamentals are weak. In the event that fundamentals are mismeasured, such endogeneity may lead to a downward bias in the estimate of the interest elasticity of investment demand.

At the firm level, variation in interest rates is due in part to the endogenous decisions made by firms when jointly determining their investment and financial policies. Such endogeneity leads to two types of biases. First, there is the potential for omitted variable bias owing to mismeasured fundamentals. Improvements in fundamentals are likely to increase investment, reduce the probability of default, and raise the price of the existing bonds. Thus, with mismeasured fundamentals, corporate yields may fall and investment spending will rise, a reflection of the endogenous response of interest rates and investment to unobservables. Such bias would lead to an overestimate of the sensitivity of investment demand to the user cost of capital. Second, everything else equal, an increase in investment may lead to higher leverage as firms finance investment using external funds. The increased leverage would raise the likelihood of default, causing an increase in the yield on outstanding corporate debt. This bias, in contrast, would lead to an underestimate of the sensitivity of investment demand to the user cost of capital. In both cases, however, the endogenous response of interest rates is fully reflected in default risk. To the extent that we can measure such risk, we can control for it in our empirical analysis. In particular, we can compute the variation in interest rates that is orthogonal to such default risk and therefore exogenous to the firm's financial policy.

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 three factors are combined into a single measure of default risk called *distance to default*.

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.[7] 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.[8] Using this insight, 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

[7]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 + $0.5 \times$ Long-Term Liabilities), and the default point is calibrated accordingly.

[8]The assumption that all of the firm's assets are traded is clearly inappropriate in most cases. Nevertheless, 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.

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.[9] 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. 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 2 depicts the evolution of the cross-sectional distribution of the expected year-ahead default frequencies for the firms in our sample. Two features are worth noting. First, at any point in time, there is considerable cross-sectional heterogeneity regarding the likelihood of default. And second, there is considerable cyclical variation in the entire cross-sectional distribution of EDFs over the 1991–2005 period.

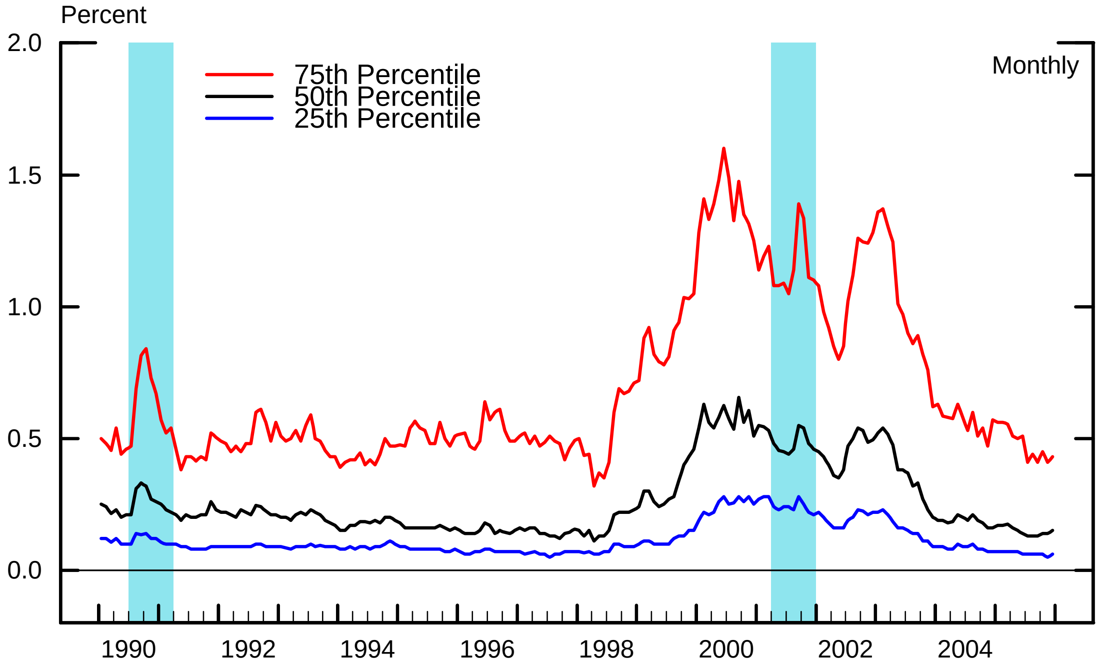
Using the information of firm-specific probabilities of default, we propose a methodology that identifies variation in firm-specific interest rates, and hence the user cost of capital, that is orthogonal to default risk as well as macroeconomic risk factors that influence the price of default risk. Specifically, let $\mathrm{SPR}_{jt_m}$ denote the difference between the firm-specific bond yield and a relevant risk-free interest rate $r_{t_m}$ for firm $j$ at the end of month $m = 1, 2, \ldots, 12$ of year $t$. According to standard asset pricing theory, the credit spread compensates bond holders for the expected cost of default, which equals the expected default probability times the recovery rate (i.e., EDF $\times$ $R$). Movements in credit spreads are also influenced significantly by macro risk factors $f_{t_m}$ and a time-varying liquidity premium $u_{jt_m}$ (see, for example, Elton, Gruber, Agrawal, and Mann [2001]). Following Berndt, Douglas, Duffie, Ferguson, and Schranz [2005], we consider a log specification of the form:

$$\log \mathrm{SPR}_{jt_m} = c_{j,t_m} + \alpha \log \mathrm{EDF}_{jt_m}, + \beta R_{jt_m} + f_{t_m} + u_{jt_m} \tag{5}$$

where $\mathrm{EDF}_{jt_m}$ denotes the year-ahead expected default frequency measured at the beginning of the month $t_m$. We include a full set of industry (3/4-digit NAICS) dummies and time

[9]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. We omitted observations with EDFs at the boundary from our analysis.

Figure 2: 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.

dummies to control for industry-specific recovery rates ($R_{jt_m}$) and macro risk factors ($f_{t_m}$). Because the relationship between interest rate spreads and default probabilities may be nonlinear, we allow for this possibility by including higher-order terms of the EDFs in the regression equation 5.

In addition to our option-theoretic measures of expected default risk, we also include external debt-rating information in the regression. Whereas the EDFs incorporate high-frequency news regarding default risk, firm-specific ratings of their senior unsecured debt contain low-frequency information that reflects a variety of firm and industry-specific information not captured by the distance to default. Specifically, letting $l = 1, 2, \ldots, L$ index the $L$ distinct credit ratings, we incorporate this information by including a full set of dummy variables that measure the firm's rating of its senior unsecured debt as additional regressors in the spread equation 5. We do so by specifying that the firm-specific term $c_{jt_m}$ satisfies:

$$c_{jt_m} = \sum_{l=1}^{L} \gamma_l 1(B_{jt_m} = l),$$

where $1(B_{jt_m} = l)$ is the indicator function for firm $j$'s credit rating $B_{jt_m}$ and the beginning of month $t_m$. The parameter $\gamma_l$, therefore, provides an estimate of the average spread for each rating category conditional on expected default.

The residual from this regression, $u_{jt_m}$ is a measure of a firm-specific liquidity premium, which, by construction is orthogonal to default risk as measured by MKMV's EDFs and the firm's external credit rating.[10] Because liquidity shocks are not determined by investment opportunities, they provide a valid instrument for the firm-specific user cost in the investment equation.

Because our investment data are on an annual basis, we construct our instrument—that is, the average annual liquidity shock—as $\hat{u}_{jt} = \sum_{m=1}^{12} \hat{u}_{jt_m}$. These instruments are constructed to match each firm's fiscal year and thus reflect information within the fiscal year $t$ about fluctuations in firm-specific interest rates. These instruments are correlated with the user-cost $C_{jt}^K$ and, according to our argument, uncorrelated with $\epsilon_{jt}$, the unobservable component of the investment demand equation. They may, therefore, be used as valid instruments to estimate the forward-mean differenced regression specified in equation 3. Formally, we are replacing the orthogonality condition $E[\epsilon_{jt} C_{j,t-s}^K \mid \eta_j, \lambda_t] = 0$ with the orthogonality condition $E[\epsilon_{jt} \hat{u}_{j,t-s} \mid \eta_j, \lambda_t] = 0$, which implies that current and lagged values of our liquidity shock $\hat{u}_{jt}$ are valid instrument for the forward-mean differenced error term $\Delta_t^{T_j}(\epsilon_{jt})$.

# 4 Results

In this section, we present our empirical results. We begin with the baseline specification described in equation 2 and estimated over the full sample period (1973–2005). Table 3 reports our baseline regression results. Columns 1 and 2 contain results in levels, whereas columns 3 and 4 contain results using the log-log specification.

The firm-specific user cost of capital is an economically important and statistically significant explanatory variable for investment in all specifications reported in Table 3. Depending on the specification, the estimate of the user-cost coefficient varies between -0.591

[10]Berndt, Douglas, Duffie, Ferguson, and Schranz [2005] use credit-default swap (CDS) data to determine the price of default risk by regressing the CDS spread on MKMV's expected default frequency. Blanco, Brennan, and Marsh [2005] argue that the bond prices suffer from liquidity premia owing to the limited supply of each firms bonds at different maturities. Because credit-default swaps are derivatives, they are less likely to suffer from liquidity premia when traded in secondary markets. Indeed, Blanco, Brennan, and Marsh [2005] use the distance between the bond price and the credit default swap as their measure of a liquidity premium, an approach consistent with our proposed methodology. More generally, we simply require that $u_{jt_m}$ reflects movements in firm-specific bond prices that are not endogenously driven by managerial financing decisions when determining investment policy. By constructing a residual, which is orthogonal to the main source of potential endogeneity considered in the asset pricing literature—namely default risk—we believe that we satisfy this criterion.

Table 3: Investment and the Cost of Capital
(Static Specification, 1973–2005)

| | Levels Specification | | Log-Levels Specification | |
|---|---|---|---|---|
| Variable | (1) | (2) | (3) | (4) |
| $\mathrm{MPK}^S_{it}$ | 0.577 | - | 0.629 | - |
| | (0.113) | | (0.034) | |
| $\mathrm{MPK}^\Pi_{it}$ | - | 0.456 | - | 1.758 |
| | | (0.033) | | (0.103) |
| $C^K_{it}$ | -0.563 | -0.591 | -0.574 | -0.495 |
| | (0.077) | (0.069) | (0.063) | (0.059) |
| Time Effects[a] | yes | yes | yes | yes |
| | (0.000) | (0.000) | (0.000) | (0.000) |
| $R^2$ (within) | 0.187 | 0.212 | 0.280 | 0.240 |
| BIC[b] | -17.18 | -17.43 | 8.302 | 8.774 |
| *Panel Dimensions* | Obs $= 8,076$ | $N = 926$ | $\bar{T} = 8.7$ | |

NOTES: Dependent variable is the investment rate in period $t$ ($I_{it}/K_{i,t-1}$). All specifications include fixed firm effects ($\eta_i$) and are estimated by OLS. Heteroscedasticity- and autocorrelation-consistent asymptotic standard errors are computed according to Arellano 1987 and are reported in parentheses.

[a] $p$-values for the test of the null hypothesis of the absence of fixed time effects are reported in parentheses.

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

and -0.495, while the standard errors are on the order of 0.07. Thus, a one-percentage-point rise in the user cost implies a one-half-percentage-point reduction in the rate of investment. The fundamentals are also important determinants of the investment rate, with coefficients that are estimated with reasonable precision.

Let $\eta_k$ and $\eta_c$ denote the elasticity of investment demand with respect to the marginal product of capital and the user cost, respectively. The elasticity of capital with respect to the user cost is then given by the ratio $-\frac{\eta_c}{\eta_k}$. In the log-log specification, using $\mathrm{MPK}^S$ as the fundamental, the long-run elasticity of capital with respect to the user cost may be determined directly from the ratio of the two coefficients. This elasticity is $\frac{0.574}{0.629} = 0.91$, which is not far from the unit elasticity implied by a Cobb-Douglas production function in capital and other inputs. When using $\mathrm{MPK}^\Pi$ to measure the marginal product of capital in the log-log specification, we must normalize the ratio of coefficients by the factor $\mathrm{MPK}^\Pi/(0.5 + \mathrm{MPK}^\Pi)$ to account for the fact that we used $Z_{jt} = \log(0.5 + \mathrm{MPK}^\Pi_{jt})$ as our measure of investment fundamentals. This adjustment factor is 0.4, hence the long-run elasticity is computed as $\frac{0.4\times 1.758}{0.629} = 1.18$, which is again close to unity. Given the mean values

Table 4: Investment and the Cost of Capital
(Dynamic Specification, 1974–2005)

| | Levels Specification | | Log-Levels Specification | |
|---|---|---|---|---|
| Variable | (1) | (2) | (3) | (4) |
| $\text{MPK}^S_{it}$ | 0.210 | - | 0.427 | - |
| | (0.054) | | (0.063) | |
| $\text{MPK}^\Pi_{it}$ | - | 0.360 | - | 1.579 |
| | | (0.055) | | (0.194) |
| $C^K_{it}$ | -0.535 | -0.623 | -0.394 | -0.443 |
| | (0.086) | (0.087) | (0.081) | (0.079) |
| $I_{i,t-1}/K_{i,t-2}$ | 0.335 | 0.317 | 0.477 | 0.472 |
| | (0.032) | (0.031) | (0.028) | (0.029) |
| Time Effects[a] | yes | yes | yes | yes |
| | (0.000) | (0.000) | (0.000) | (0.000) |
| $\Pr > \lvert m_1 \rvert$ [b] | 0.000 | 0.000 | 0.000 | 0.000 |
| $\Pr > \lvert m_2 \rvert$ [c] | 0.501 | 0.479 | 0.750 | 0.620 |
| $\Pr > J_N$ [d] | 1.000 | 1.000 | 1.000 | 1.000 |
| *Panel Dimensions* | Obs $= 4,911$ $N = 677$ $\bar{T} = 7.3$ | | | |

Notes: Dependent variable is the investment rate in period $t$ ($I_{it}/K_{i,t-1}$). All specifications include fixed firm effects ($\eta_i$), which are eliminated using forward orthogonal deviation transformation. The resulting specification is estimated by GMM using a one-step weighting matrix; see Arellano 2003. The instrument set includes lags 2 to 5 of ($I_t/K_{t-1}$), $\text{MPK}^S_{it}$, $\text{MPK}^\Pi_{it}$, and $C^K_{it}$. Heteroscedasticity- and autocorrelation-consistent asymptotic standard errors are reported in parentheses.

[a] $p$-values for the test of the null hypothesis of the absence of fixed time effects are reported in parentheses.

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

[c] $p$-value for the test of the second-order serial correlation of the transformed residuals.

[d] $p$-value for the Hansen 1982 test of the overidentifying restrictions. This test uses the minimized objective of the corresponding two-step GMM estimator.

of $\text{MPK}^S$, $\text{MPK}^\Pi$, and the user cost $C^K$, the specifications in Table 3 imply long-run elasticities of capital with respect to the user cost between 0.66 and 1.21 for the specifications using the fundamentals constructed from sales and operating income, respectively.

Table 4 reports estimates of the forward mean-differenced specification of the model that allows for a lagged dependent variable as an additional explanatory variable. The user cost of capital is again an economically and statistically significant explanatory variable for investment with coefficients that are similar to those reported in Table 3. The log-log specifications (columns 3 and 4) imply a long-run elasticity of capital with respect to the user cost of 0.92 when using sales to construct the marginal product of capital, and 0.70

when operating income is used as a fundamental. The levels specifications (columns 1 and 2), however, imply substantially higher elasticities—in both cases they are greater that 1.5.

In Table 5, we consider a robustness exercise that allows the coefficient on the investment fundamentals and the user cost to vary across sectors. We report results for the log-log specification, using $\text{MPK}^{\Pi}$ as our measure of investment fundamentals. The top panel of the table reports sector-specific elasticities of investment demand with respect to the marginal product of capital $\text{MPK}^{\Pi}$ and the bottom half contains elasticities with respect to the user cost of capital $C^K$. We consider the full sample period, 1973–2005, as well as the two sub-sample periods, namely 1973–1990 and 1991–2005.

According to entries in Table 5, the elasticity of investment to the user cost of capital is negative and statistically significant across nearly all sectors and time periods. Thus, our finding that an increase in the user cost has a strong negative impact on investment spending is broad-based and is not driven by a small number of observations or data from a single sector. The three largest sectors in terms of number of firms, and percentage of economic activity are the nondurable manufacturing, durable manufacturing and information services. For these sectors, our results for the full sample period imply long-run elasticities that are again close to unity.

## 4.1 Liquidity Shocks and the User Cost of Capital

We now consider a robustness exercise that uses our measure of liquidity shocks as instruments for the user cost of capital. As discussed above, our approach is to regress the spread on credit ratings and expected default frequency, controlling for fixed time and industry effects that capture time variation in the price of risk and industry variation in recovery rates. The residual from this regression is the liquidity premium, which we use as the instrument for the user cost of capital. Because the MKMV data are only available from 1991 onwards, our results are based on the 1991–2005 sub-sample. We first report the interest rate spread regressions. Then we recompute the results reported in Tables 3 and 4 for the 1991-2005 period. This allows us to make clear comparisons with the results that use liquidity shocks to instrument for the user cost of capital.

In addition to adopting an instrumental variables estimation approach, we also modify our construction of the user cost by using industry-level (3/4-digit NAICS) data, as opposed to sectoral data, for the key components of the user cost in equation 1. That is, the relative price of capital $P^K_{st}/P^Q_{st}$, the depreciation rate $\delta_{st}$, and the expected capital gains term $E_t[\Delta P^K_{s,t+1}/P^K_{st}]$ are allowed to vary across 51 industries. Our motivation here is that such data are preferable to use when constructing the user cost but, unfortunately, are only available from 1985 onward. Thus all regressions in this section incorporate industry-level data when constructing the user cost of capital.

Table 5: Investment and the Cost of Capital
(Static Sectoral Specification)

| | | Elasticity w.r.t. $MPK^{\Pi}$ by Sample Period[a] | | | | | |
|---|---|---|---|---|---|---|---|
| | | 1973–2005 | | 1973–1990 | | 1991–2005 | |
| Sector | $N$ | *Est.* | *S.E.* | *Est.* | *S.E.* | *Est.* | *S.E.* |
| Mining | 44 | 0.471 | 0.089 | 0.413 | 0.103 | 0.493 | 0.131 |
| Utilities | 74 | 0.303 | 0.111 | -0.028 | 0.113 | 0.613 | 0.151 |
| Construction | 8 | 0.448 | 0.177 | 0.901 | 0.189 | 0.234 | 0.144 |
| Mfg. (nondurable) | 220 | 0.608 | 0.058 | 0.607 | 0.082 | 0.674 | 0.074 |
| Mfg. (durable) | 230 | 0.463 | 0.039 | 0.484 | 0.057 | 0.491 | 0.051 |
| Wholesale Trade | 25 | 0.397 | 0.127 | 0.325 | 0.366 | 0.401 | 0.094 |
| Retail Trade | 73 | 0.558 | 0.137 | 0.725 | 0.205 | 0.783 | 0.136 |
| Transportation | 51 | 0.560 | 0.155 | 0.963 | 0.199 | 0.252 | 0.231 |
| Information | 118 | 0.516 | 0.103 | 0.537 | 0.144 | 0.495 | 0.127 |
| Services | 83 | 0.373 | 0.099 | 0.229 | 0.091 | 0.457 | 0.123 |
| | | Elasticity w.r.t. $C^K$ by Sample Period | | | | | |
| | | 1973–2005 | | 1973–1990 | | 1991–2005 | |
| Sector | $N$ | *Est.* | *S.E.* | *Est.* | *S.E.* | *Est.* | *S.E.* |
| Mining | 44 | -0.279 | 0.060 | -0.144 | 0.055 | -0.564 | 0.377 |
| Utilities | 74 | -0.290 | 0.060 | -0.219 | 0.076 | -0.542 | 0.257 |
| Construction | 8 | -0.079 | 0.681 | -0.613 | 0.158 | 2.142 | 0.853 |
| Mfg. (nondurable) | 220 | -0.694 | 0.069 | -0.438 | 0.093 | -1.078 | 0.229 |
| Mfg. (durable) | 230 | -0.666 | 0.083 | -0.491 | 0.102 | -0.817 | 0.141 |
| Wholesale Trade | 25 | -1.041 | 0.145 | -1.202 | 0.296 | -1.086 | 0.178 |
| Retail Trade | 73 | -1.260 | 0.233 | -0.796 | 0.314 | -1.617 | 0.303 |
| Transportation | 51 | -0.595 | 0.131 | -0.468 | 0.145 | -0.862 | 0.332 |
| Information | 118 | -0.638 | 0.137 | -0.551 | 0.154 | -0.716 | 0.179 |
| Services | 83 | -1.023 | 0.174 | -0.541 | 0.129 | -1.915 | 0.414 |
| $R^2$ (within) | | 0.253 | | 0.287 | | 0.276 | |
| Obs./$N$ | | 8,076/926 | | 3,554/703 | | 4,522/787 | |

NOTES: Dependent variable is $\log(I_{it}/K_{i,t-1})$. The explanatory variables are $\log(0.5 + \text{MPK}^{\Pi}_{it})$ and $\log C^K_{it}$. All specifications include fixed firm effects ($\eta_i$) and fixed time effects ($\lambda_t$) and are estimated by OLS. Heteroscedasticity- and autocorrelation-consistent asymptotic standard errors are computed according to Arellano 1987.

[a]The reported elasticities and the associated standard errors are adjusted for the fact that the dependent variable is transformed as $\log(0.5 + \text{MPK}^{\Pi}_{it})$.

Table 6: Credit Spreads and Expected Default Risk

| | Specification | | | |
|---|---|---|---|---|
| Variable | (1) | (2) | (3) | (4) |
| $\text{EDF}_{i,t-1}$ | 1.771 | 1.018 | 1.018 | 0.654 |
| | (0.049) | (0.041) | (0.042) | (0.040) |
| $\text{EDF}^2_{i,t-1}$ | -0.204 | -0.114 | -0.113 | -0.061 |
| | (0.009) | (0.007) | (0.007) | (0.007) |
| $\text{EDF}^3_{i,t-1}$ | 0.007 | 0.004 | 0.004 | 0.002 |
| | (0.000) | (0.000) | (0.000) | (0.000) |
| $R^2$ | 0.529 | 0.664 | 0.674 | 0.746 |
| BIC[a] | 175.02 | 155.68 | 154.46 | 142.14 |
| Ratings Effects[b] | no | yes | yes | yes |
| | | (0.000) | (0.000) | (0.000) |
| Industry Effects[c] | no | no | yes | yes |
| | | | (0.000) | (0.000) |
| Time Effects[d] | no | no | no | yes |
| | | | | (0.000) |
| *Panel Dimensions* | Obs $= 58,037$ | $N = 872$ | $\bar{T} = 66.6$ | |

NOTES: Estimation period: Monthly data from February 1991 to December 2005. Dependent variable is the credit spread in month $t$ measured relative to the comparable maturity Treasury yield ($\text{SPR}_{it}$). All specifications include a constant (not reported) and are estimated by OLS. Heteroscedasticity- and autocorrelation-consistent asymptotic standard errors are reported in parentheses.

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

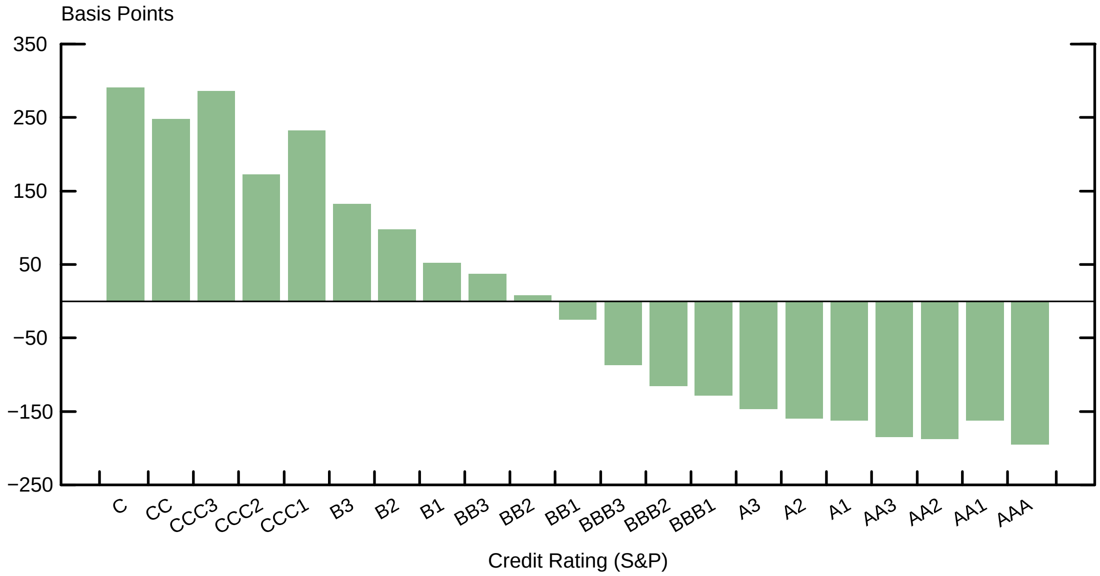
[b]$p$-values for the test of the null hypothesis of the absence of fixed credit ratings effects are reported in parentheses.

[c]$p$-values for the test of the null hypothesis of the absence of fixed industry effects are reported in parentheses.

[d]$p$-values for the test of the null hypothesis of the absence of fixed time effects are reported in parentheses.

Table 6 reports the regression results from estimating equation 5. We consider four separate specifications. The first specification includes a linear, squared, and cubed term of the expected default frequency. We then augment this baseline specification sequentially with credit ratings effects, industry effects, and time effects. Although the higher-order EDF terms are statistically significant, the linear term dominates in all four specifications. The EDF terms alone yield an $R^2$ equal to 0.529, which is similar to that reported by Berndt, Douglas, Duffie, Ferguson, and Schranz [2005] who regressed credit default swap spreads on the EDF. As expected, the inclusion of credit rating dummies reduces the effect of the

Figure 3: The Effect of Ratings on Credit Spreads




NOTES: This figure depicts the estimated effect of external credit ratings on the interest rate spreads.

expected default frequency and significantly raises the explanatory power of the regression. Although industry effects do not appear to have a large impact, the addition of the time dummies—which is meant to capture variation in macro risk factors—also significantly improves the fit of the regression and reduces the coefficient on the EDF. With all four factors included, the regression explains 74% of the monthly variation in the corporate credit spreads. As discussed above, the residual from this regression is our measure of the liquidity shock.

Figure 3 displays the estimated coefficients on the ratings indicators that are included in the regression. As expected, these coefficients imply a strong negative relationship between ratings and credit spreads. Our estimates imply a 400 basis point differential in yield spreads between C-rated firms and AAA-rated firms. Most of this variation occurs between the B3 and A3 ratings categories.

Tables 7 and 8 replicate the results reported in Tables 3 and 4 for the 1991–2005 sub-sample. The coefficients reported for this subperiod have somewhat higher standard errors but otherwise are of similar magnitude as those obtained for the full sample. An exception here is the coefficient on the sales-based measure of MPK, which appears to be biased downward relative to the full sample results. In all cases, the user cost coefficient is estimated to be higher for the sub-sample than the full sample. It is possible that the effect is stronger because the user cost constructed for the 1991–2005 uses better price information and thus

Table 7: Investment and the Cost of Capital
(Static Specification, 1991–2005)

| | Levels Specification | | Log-Levels Specification | |
|---|---|---|---|---|
| Variable | (1) | (2) | (3) | (4) |
| $\text{MPK}^S_{it}$ | 0.333 | - | 0.737 | - |
| | (0.156) | | (0.043) | |
| $\text{MPK}^\Pi_{it}$ | - | 0.459 | - | 1.857 |
| | | (0.046) | | (0.160) |
| $C^K_{it}$ | -0.678 | -0.637 | -0.769 | -0.527 |
| | (0.123) | (0.137) | (0.095) | (0.090) |
| Time Effects[a] | yes | yes | yes | yes |
| | (0.000) | (0.000) | (0.000) | (0.000) |
| $R^2$ (within) | 0.180 | 0.231 | 0.330 | 0.253 |
| BIC[b] | -10.21 | -10.50 | 3.687 | 4.183 |
| *Panel Dimensions* | Obs $= 4,532$ | $N = 789$ | $\bar{T} = 5.7$ | |

NOTES: Dependent variable is the investment rate in period $t$ ($I_{it}/K_{i,t-1}$). All specifications include fixed firm effects ($\eta_i$) and are estimated by OLS. Heteroscedasticity- and autocorrelation-consistent asymptotic standard errors are computed according to Arellano 1987 and are reported in parentheses.

[a]$p$-values for the test of the null hypothesis of the absence of fixed time effects are reported in parentheses.

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

has less measurement error.

Table 9 reports the dynamic specification using the liquidity shock as an instrument for the user cost. It is thus directly comparable to Table 8, which uses lagged values of the user cost as an instrument. With the exception of the first column, which uses the sales-based measure of fundamentals in a levels specification, the user cost coefficient is highly significant and estimated to be very similar across the specification that instruments with the user cost versus the specification that instruments with the liquidity shock. These results strongly suggest that the effect of interest rates on investment embedded in our user cost framework are not driven by endogeneity between interest rates and investment policy at the firm-level.

# 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

Table 8: Investment and the Cost of Capital
(Dynamic Specification, 1991–2005)

| | Levels Specification | | Log-Levels Specification | |
|---|---|---|---|---|
| Variable | (1) | (2) | (3) | (4) |
| $\text{MPK}_{it}^{S}$ | 0.138 | - | 0.533 | - |
| | (0.045) | | (0.078) | |
| $\text{MPK}_{it}^{\Pi}$ | - | 0.355 | - | 1.617 |
| | | (0.079) | | (0.308) |
| $C_{it}^{K}$ | -0.555 | -0.679 | -0.617 | -0.478 |
| | (0.134) | (0.141) | (0.137) | (0.132) |
| $I_{i,t-1}/K_{i,t-2}$ | 0.388 | 0.367 | 0.449 | 0.486 |
| | (0.047) | (0.045) | (0.037) | (0.039) |
| Time Effects$^a$ | yes | yes | yes | yes |
| | (0.000) | (0.000) | (0.000) | (0.000) |
| $\Pr > \lvert m_1 \rvert^b$ | 0.000 | 0.000 | 0.000 | 0.000 |
| $\Pr > \lvert m_2 \rvert^c$ | 0.495 | 0.393 | 0.782 | 0.514 |
| $\Pr > J_N{}^d$ | 0.997 | 0.988 | 0.989 | 0.979 |
| *Panel Dimensions* | Obs $= 3,033$ $\quad N = 566$ $\quad \bar{T} = 5.4$ | | | |

NOTES: Dependent variable is the investment rate in period $t$ ($I_{it}/K_{i,t-1}$). All specifications include fixed firm effects ($\eta_i$), which are eliminated using forward orthogonal deviation transformation. The resulting specification is estimated by GMM using a one-step weighting matrix; see Arellano 2003. The instrument set includes lags 2 to 5 of $(I_t/K_{t-1})$, $\text{MPK}_{it}^{S}$, $\text{MPK}_{it}^{\Pi}$, and $C_{it}^{K}$. Heteroscedasticity- and autocorrelation-consistent asymptotic standard errors are reported in parentheses.

$^a$$p$-values for the test of the null hypothesis of the absence of fixed time effects are reported in parentheses.

$^b$$p$-value for the test of the first-order serial correlation of the transformed residuals.

$^c$$p$-value for the test of the second-order serial correlation of the transformed residuals.

$^d$$p$-value for the Hansen 1982 test of the overidentifying restrictions. This test uses the minimized objective of the corresponding two-step GMM estimator.

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 investment spending, our estimation results imply a robust and quantitatively important effect of interest rates on investment at the firm-level. Roughly speaking, a one percentage point rise in the user cost implies somewhere between a 70 to 130 basis point reduction in investment. Consistent with previous results documented in Cummins, Hassett, and Hubbard [1994, 1996], our results imply a strong link between the user

Table 9: Investment, Cost of Capital, and Liquidity Shocks
(Dynamic Specification, 1991–2005)

| | Levels Specification | | Log-Levels Specification | |
|---|---|---|---|---|
| Variable | (1) | (2) | (3) | (4) |
| $\text{MPK}_{it}^{S}$ | 0.123 | - | 0.673 | - |
| | (0.040) | | (0.120) | |
| $\text{MPK}_{it}^{\Pi}$ | - | 0.349 | - | 1.254 |
| | | (0.087) | | (0.379) |
| $C_{it}^{K}$ | -0.278 | -0.664 | -0.766 | -0.485 |
| | (0.243) | (0.172) | (0.198) | (0.156) |
| $I_{i,t-1}/K_{i,t-2}$ | 0.398 | 0.369 | 0.454 | 0.461 |
| | (0.051) | (0.049) | (0.044) | (0.050) |
| Time Effects$^a$ | yes | yes | yes | yes |
| | (0.000) | (0.000) | (0.000) | (0.000) |
| $\Pr > \lvert m_1 \rvert^b$ | 0.000 | 0.000 | 0.000 | 0.000 |
| $\Pr > \lvert m_2 \rvert^c$ | 0.517 | 0.414 | 0.856 | 0.637 |
| $\Pr > J_N{}^d$ | 0.999 | 1.000 | 1.000 | 0.976 |
| *Panel Dimensions* | Obs $= 3,118$ $N = 582$ $\bar{T} = 5.4$ | | | |

NOTES: Dependent variable is the investment rate in period $t$ ($I_{it}/K_{i,t-1}$). All specifications include fixed firm effects ($\eta_i$), which are eliminated using forward orthogonal deviation transformation. The resulting specification is estimated by GMM using a one-step weighting matrix; see Arellano 2003. The instrument set includes lags 2 to 5 of $(I_t/K_{t-1})$, $\text{MPK}_{it}^{S}$, $\text{MPK}_{it}^{\Pi}$, and the estimate of the average liquidity shock in period $t$ (see text for details). Heteroscedasticity- and autocorrelation-consistent asymptotic standard errors are reported in parentheses.

[a] $p$-values for the test of the null hypothesis of the absence of fixed time effects are reported in parentheses.

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

[c] $p$-value for the test of the second-order serial correlation of the transformed residuals.

[d] $p$-value for the Hansen 1982 test of the overidentifying restrictions. This test uses the minimized objective of the corresponding two-step GMM estimator.

cost of capital and investment spending at the firm level. These findings have important implications for the conduct of both monetary and fiscal policy.

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