Corporate Diversification and the Cost of Capital

by

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Abstract

This paper examines whether organizational form matters for a firm’s cost of capital. We develop a model to show that corporate diversification may reduce not only idiosyncratic risk but also systematic risk. Using measures of implied cost of capital constructed from analyst forecasts, we find that diversified firms have on average a lower cost of capital than stand-alone firms. In addition, the cost of capital is lower when the correlation of cash flows among the diversified firm’s segments is lower. The observed cost of capital reduction yields an average value gain of approximately 6% when moving from the highest to the lowest cash flow correlation quintile. Overall, our results are consistent with the coinsurance effect of diversification lowering a firm’s cost of capital.

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1. Introduction

Conglomeration is ubiquitous both in the U.S. and around the world. With over half of total equity market capitalization represented by firms reporting two or more business segments, it is not surprising that a large literature in corporate finance examines whether and how corporate diversification affects firm value. Much of this literature implicitly attributes observed valuation differences between diversified and stand-alone firms to differences in cash flows, perhaps reflecting the conventional view among practitioners and researchers that organizational form does not matter for a firm’s cost of capital because corporate diversification reduces only idiosyncratic risk and not systematic risk.\(^1\) In this paper we present theoretical and empirical evidence that is contrary to the conventional view. In particular, we show that organizational form matters for a firm’s cost of capital – corporate diversification tends to reduce systematic risk and diversified firms have a lower cost of capital than stand-alone firms, consistent with a coinsurance effect.

Prior research inspired by Coase’s (1937) fundamental question about the boundaries of the firm points to various integration costs (Rajan, Servaes, and Zingales (2000), Scharfstein and Stein (2000)) and benefits (Matsusaka and Nanda (2002), Stein (1997)) that come with corporate diversification. In theory, if the cash flows associated with these costs and benefits carry precisely the same systematic risk as the cash flows of the underlying portfolio of businesses do as stand-alone firms, then the conventional view holds – organizational form will not affect cost of capital. Short of this restrictive condition, however, a diversified firm’s cost of capital will

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\(^1\) The conventional view has long been a part of standard finance textbooks such as Brealey, Myers, and Allen or Ross, Westerfield, and Jaffe, and thus may alternatively be referred to as the textbook view. The notion that corporate diversification cannot affect systematic risk is usually covered explicitly in the mergers and acquisitions chapter (e.g., “Systematic variability cannot be eliminated by diversification, so mergers will not eliminate this risk at all,” RWJ, p. 823) or implicitly in the capital budgeting chapter.
differ from that of a portfolio of comparable stand-alone firms. For instance, if integration benefits (costs) carry less (more) systematic risk, then a diversified firm should have a lower cost of capital.

Building on this insight about integration, we argue that organizational form matters for cost of capital by connecting two well-known but separate observations: (i) coinsurance – the imperfect correlation among the cash flows of a diversified firm’s business units – reduces default risk (Lewellen (1971)); and (ii) default risk has a systematic component (Elton, Gruber, Agrawal, and Mann (2001), Almeida and Philippon (2007)). Intuitively, if the reduction in default risk due to coinsurance has a systematic component (i.e., financial distress costs are partly systematic), then corporate diversification should reduce systematic risk. In this paper, we show in a parsimonious model that the coinsurance idea extends to an all-equity firm if one replaces costs of financial distress with other kinds of deadweight costs that even an all-equity firm might face. Our main result is that coinsurance and the ability of a diversified firm to avoid deadweight costs by transferring financial resources from cash-rich units to cash-poor units tend to reduce systematic risk. In addition, we show that the coinsurance effect is more pronounced when the firm’s units have less correlated cash flows.

We empirically examine whether and how organizational form affects a firm’s cost of capital using a large sample of single- and multi-segment firms spanning the period 1988 to 2006. Our cost of capital proxy is the weighted average of cost of equity and cost of debt. We use ex ante measures of the implied cost of equity constructed from analyst forecasts to proxy for

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2 We assume that it is difficult or costly to write complete state-contingent contracts to transfer resources between cash-rich and cash-poor firms. Without this assumption, corporate diversification would offer no benefit over what investors can achieve through portfolio diversification. We also assume that holding financial slack is costly, as otherwise firms would hold the first-best amount of financial slack to avoid any future deadweight cost.

3 We consider several extensions, including the possibility of integration costs arising from inefficient transfers and agency problems within diversified firms, and show that under certain conditions the coinsurance effect can be consistent with a diversification discount.
expected equity returns and yields from the Barclays Capital Aggregate Bond Index to proxy for expected debt returns. We estimate the implied cost of equity based on the approach of Gebhardt, Lee, and Swaminathan (2001), which has been successfully employed in several asset-pricing contexts (Lee, Ng, and Swaminathan (2007), Pastor, Sinha, and Swaminathan (2008)). Our empirical analyses are based on an “excess cost of capital” measure that benchmarks the cost of capital of a diversified firm against that of a comparable portfolio of stand-alone firms.

We find that diversified firms on average have significantly lower cost of capital compared to portfolios of stand-alone firms. Further, we find a significant and positive association between excess cost of capital and cross-segment cash flow correlations, consistent with a coinsurance effect. These findings are robust to using alternative measures of implied cost of equity capital (Claus and Thomas (2001), Easton (2004)) and coinsurance (Duchin (2008)) and to controlling for analyst forecast errors. Finally, our estimates imply an average cost of capital reduction of approximately 3% and an average value gain of approximately 6% when moving from the highest to the lowest cash flow correlation quintile. These magnitudes potentially represent a lower bound on the importance of coinsurance because our proxies do not capture coinsurance that can also arise from alternative dimensions of corporate diversification such as diversity across product lines or geographic areas.

This paper makes several contributions to the literature. First, our study is the first to establish a link between the level of coinsurance among a firm’s business units and the systematic risk of its cash flows, and hence between coinsurance and cost of capital. Following Lewellen’s (1971) seminal work, a stream of research studies coinsurance in the context of

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4 Our empirical proxy for expected debt returns is admittedly a relatively crude proxy, as it is an aggregate measure and hence does not capture any firm-level variation in expected debt returns. However, as Lamont and Polk (2001) point out, debt returns are not readily available for most firms and using a proxy that measures only expected equity returns ignores the importance of debt in a firm’s capital structure. To the extent that coinsurance lowers both cost of equity and cost of debt, our empirical proxy would underststate the effect of coinsurance on total cost of capital.
conglomerate mergers (Higgins and Schall (1975), Scott (1977)) and examines whether such mergers lead to wealth transfers from shareholders to bondholders (Kim and McConnell (1977)). However, whether coinsurance arising from corporate diversification affects cost of capital beyond a tax-based leverage effect has not been explored in prior studies.

Second, our study complements the extant literature on corporate diversification and firm value by exploring an important dimension that thus far has received little attention, namely, cost of capital. Since Lang and Stulz (1994) and Berger and Ofek (1995), the diversification discount has been heavily debated and a wealth of papers have explored various explanations of its existence or absence (Campa and Kedia (2002), Graham, Lemmon, and Wolf (2002), Villalonga (2004)). The discussion in this literature mostly revolves around cash flow differences between conglomerates and stand-alone firms. An exception is Lamont and Polk (2001), who raise the possibility that the discount (or premium) may arise due to differences in expected returns. They find a significant and negative association between excess values and future returns for diversified firms, suggesting that the diversification discount is explained in part by differences in expected returns. However, while their study introduces the important role of expected returns in understanding the valuation of diversified firms, their main focus is to explain the cross-sectional variation in excess value, and not how diversification affects a firm’s cost of capital, which is a focus of the present paper.

Lastly, our evidence has implications for capital budgeting. In practice, managers tend to ignore the coinsurance benefit of enhanced debt capacity and the resulting tax-related reduction in weighted average cost of capital in their capital budgeting decisions, perhaps because they perceive the tax effect to be small. Our results provide two interesting insights on this issue. First, our model shows that there is a coinsurance effect even in the absence of taxes or debt
financing. Second, investors appear to understand the effect of diversification on systematic risk and adjust the discount rate they use in valuing expected future cash flows accordingly. Taken together, our findings suggest that ignoring coinsurance effects and using project-specific discount rates as commonly taught and practiced may yield incorrect (i.e., understated) NPV estimates. In our model, the covariance between a proposed project’s cash flows and those of existing projects determines both the expected level and the systematic risk of synergistic coinsurance cash flows. As a result, covariances matter for capital budgeting (Lintner (1965)).

The remainder of the paper is organized as follows. Section 2 develops our model, which shows that corporate diversification can reduce not only idiosyncratic but also systematic risk. Section 3 discusses the valuation approach we use in estimating the implied cost of equity and its empirical implementation, along with the construction of the excess cost of capital and coinsurance measures. Section 4 describes our sample and data. Section 5 presents our empirical results. Section 6 concludes.

2. A Model of Corporate Diversification and the Cost of Capital

As discussed earlier, the conventional view on corporate diversification is that it reduces only idiosyncratic risk. In this section, we outline a parsimonious model of corporate diversification to demonstrate how integrating business units with imperfectly correlated cash flows under one roof can also lead to a reduction in systematic risk and hence the cost of capital.

We begin with a basic model that assumes all-equity financing and an efficient internal capital market to illustrate the coinsurance effect. We then relax these assumptions and extend

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5 Standard finance textbooks either explicitly cite or implicitly follow Schall (1972) in emphasizing the irrelevancy of covariance and corporate diversification when explaining the stand-alone principle and potential synergy adjustments in capital budgeting. Schall’s analysis rules out by assumption the possibility that synergistic cash flows may be a function of covariance in footnote 13.
the basic model to incorporate debt financing\textsuperscript{6} and the possibility of rent-seeking activities and inefficient transfers in internal capital markets. Finally, we present our model’s testable predictions to conclude the section.

2.1. The Two-state Economy and the Relation between Asset Betas and Expected Returns

Before we introduce firms, we first describe the two-state economy in which we study corporate diversification and the effect of coinsurance on cost of capital.

Suppose that the economy has two dates, \( t \in \{0, 1\} \), and is populated with risk-averse investors. At \( t = 1 \), the economy can be either good (\( g \)) or bad (\( b \)) with probability \( p_g \) and \( (1 - p_g) \), respectively.

In the absence of arbitrage, there exists a strictly positive stochastic discount factor \( m \) \((m_g < m_b)\) that prices all assets with cash flow \( C \) at \( t = 1 \) according to the relation

\[ E[C \cdot m] = V, \]

where \( E \) is the expectation operator and \( V \) is the value of the asset at \( t = 0 \). We are interested in the pricing of traded assets with positive cash flow \( C \in R^+ \).

In the two-state economy described above, the value of asset \( i \) at \( t = 0 \) with cash flow \( C^i \) at \( t = 1 \) is given by

\[ p_g C^i_g m_g + (1 - p_g) C^i_b m_b = V^i. \]  

\textsuperscript{6} We note that allowing for different tax treatment of debt and equity, and in particular a tax advantage for debt, further reduces total cost of capital because another benefit of corporate diversification may be to increase debt capacity as in Lewellen (1971).
**Definition 1** The expected rate of return on asset $i$, $E[r^i]$, is the discount rate that equates the discounted value of asset $i$’s expected cash flow at $t=1$ to asset $i$’s value at $t=0$.

$$\frac{E[C^i]}{1+E[r^i]} = V^i$$

(2)

In a risk-averse economy, equilibrium expected returns compensate investors for holding assets that offer systematically more cash flow in the good state than in the bad state.

Let $\beta^i \equiv (C^i_s/C^i_b - 1)$. Note that $\beta^i$ is monotone in the conventional measure of systematic risk $-\text{cov}(C^i,m)$ because $m_s < m_b$. This means that we can use $\beta^i$ as an analytically convenient measure of the systematic risk of asset $i$’s cash flow in deriving comparative statics. The following lemma formalizes this.

**Lemma 1** Given an equilibrium summarized by $(p_s,m_s,m_b)$, $E[r^i]$ depends only on $\beta^i$ and increases in $\beta^i$.

**Proof.** Substituting equation (1) into (2),

$$1 + E[r^i] = \frac{p_s C^i_s + (1-p_s)C^i_b}{p_s C^i_s m_s + (1-p_s)C^i_b m_b}.$$  

Restating $E[r^i]$ in terms of $\beta^i$ in an equilibrium summarized by $(p_s,m_s,m_b)$,

$$E[r^i] = \frac{p_s \beta^i + 1}{p_s \beta^i m_s + (1-p_s)m_b + p_s m_s} - 1.$$  

Simple algebra shows that

$$\frac{\partial E[r^i]}{\partial \beta^i} = \frac{p_s(1-p_s)(m_b - m_s)}{[p_s \beta^i m_s + (1-p_s)m_b + p_s m_s]^2}.$$  


Since the probability-adjusted value of cash flow in the bad state $m_b$ is greater than the probability-adjusted value of cash flow in the good state $m_g$, $\partial E[r]/\partial \beta > 0$. \textit{Q.E.D.}

2.2. \textbf{Firm Cash Flows and the Cost of Capital}

Having established the relation between betas and equilibrium expected returns, we now turn to firm cash flows and the cost of capital in our model. A maintained assumption in the model is that it is difficult or costly to write complete state-contingent contracts to transfer resources between cash-rich and cash-poor firms. Without this assumption, corporate diversification would offer no cost of capital benefit over what investors can achieve through portfolio diversification. Another maintained assumption in the model is that holding financial slack is costly. Otherwise, firms would hold first-best amount of financial slack to avoid any future deadweight cost.

2.2.1. \textit{All-equity Financing}

In the basic model, we analyze all-equity firms to show that coinsurance can reduce systematic risk even in the absence of debt financing and taxes. In Section 2.2.3, we extend the model to show that the coinsurance effect can also apply to debt financing.

\textit{One Stand-alone Firm}

Suppose that a stand-alone firm is a project that experiences either a high ($h$) or a low ($l$) outcome with probability $\theta$ and $(1-\theta)$, respectively. The parameter $\theta$ depends on the state of the economy. Specifically, the probability of a high outcome is $\theta_b$ ($\theta_g$) when the economy is good (bad).
Investors receive $H$ when the project’s outcome is $h$. When the project’s outcome is $l$, lack of confidence in the firm leads to costly defections by important stakeholders such as suppliers and customers, in which case the firm incurs a deadweight loss $L$ and investors receive 0.\(^7\)

Further suppose that there are sufficiently many firms in the economy that investors can diversify away firm-specific idiosyncratic risk. Thus, investors only care about the expected cash flow in each state of the economy. The expected rate of return on a stand-alone firm ($S$) is determined by $\beta^S \equiv \left( \frac{C^S_g}{C^S_b} - 1 \right)$, where $C^S_g = \theta_g H$ and $C^S_b = \theta_b H$. A stand-alone firm whose $\theta_g > \theta_b$ carries positive systematic risk whereas a stand-alone firm whose $\theta_g < \theta_b$ carries negative systematic risk. Correspondingly, the former has a higher cost of capital than the latter. Risk-averse investors demand a risk premium for investing in assets that offer more expected cash flow when the economy is good than when the economy is bad.

**Combining Two Stand-alone Firms into One Diversified Firm**

Suppose that two identical stand-alone firms can be combined under one roof.\(^8\) A benefit of such a corporate structure is that when one of the projects experiences a low outcome, important stakeholders of the project do not defect if the other project has a high outcome because the firm has the ability to transfer financial resources from the high-outcome project to

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\(^7\) The assumption about investors receiving nothing is without loss of generality. The loss $L$ and the decision of important stakeholders to defect from an *all-equity* firm after observing a low outcome can be given microfoundation with costly external finance. In a multi-period model, the defection decision of suppliers and customers can be driven by concerns about the willingness of the firm to maintain relationship-specific investments (exceeding the firm’s riskless debt capacity) if the returns on such investments are greater than the cost of internal finance (in insufficient supply following a low outcome) but lower than the cost of external finance. Another concern of outside parties may be counterparty exposure when entering into long-term contracts. Further, employees may defect if they think waiting to find new employment until other employees are doing the same would be costly. Hence, $L$ represents the present value of both current and future losses.

\(^8\) For completeness, we note that the integration possibility we consider is small relative to the size of the economy. Hence, we can take the stochastic discount factor $m$ as exogenous and study the effect of corporate diversification on cash flows and systematic risk without having to consider the general equilibrium effect on $m$. 
the low-outcome project, or alternatively, use the high-outcome project as collateral to obtain external financing for the low-outcome project. Hence, while a stand-alone firm incurs some deadweight loss $L$ when the project’s outcome is $l$, a diversified firm with two projects may avoid this loss if the outcome of at least one of the two projects is $h$.\footnote{Our results hold as long as at least some of the deadweight loss can be avoided. Also, our setup allows for the possibility of contagion ($L < 0$), the opposite of coinsurance ($L > 0$). Indeed, all of our testable implications can be stated in terms of contagion, which our empirical tests reject in favor of coinsurance.} However, if both projects experience a low outcome, then even a diversified firm cannot avoid the deadweight loss.

Enumerating the possible project outcomes $(hh, lh, hl, ll)$ for a diversified firm $(D)$ comprising two stand-alone firms with independent idiosyncratic risks, the cash flows in the good and bad states of the economy are given by

$$C^D_g = \theta^2_s (2H) + 2\theta_s (1 - \theta_s) (H + L) \left( = 2\theta_s H + 2\theta_s (1 - \theta_s) L \right)$$

$$C^D_b = \theta^2_b (2H) + 2\theta_b (1 - \theta_b) (H + L) \left( = 2\theta_b H + 2\theta_b (1 - \theta_b) L \right).$$

Without the terms involving $L$, the expected cash flow of a diversified firm $C^D_e$ equals twice the expected cash flow of a stand-alone firm, $2C^s_e$, for $e \in \{g,b\}$. That is, without real coinsurance, a diversified firm offers nothing that investors cannot achieve on their own by investing in two stand-alone firms.

As the next proposition shows, one implication of coinsurance may be to reduce systematic risk in addition to increasing cash flows $(C^D_g > 2C^s_g, C^D_b > 2C^s_b)$.\footnote{Coinsurance effects arise even if we introduce integration costs to make the analysis cash-neutral. See Section 2.2.2 where we introduce integration costs.}

**Proposition 1** Combining two stand-alone firms with positive systematic and independent idiosyncratic risks reduces systematic risk and cost of capital.

**Proof.** Given Lemma 1, it suffices to show that $\beta^s > \beta^D$. Substituting the cash flows above,
\[ \beta^S = \frac{\theta_s H}{\theta_b H} - 1 \]

\[ \beta^D = \frac{2\theta_s H + 2\theta_s (1 - \theta_s) L}{2\theta_b H + 2\theta_b (1 - \theta_b) L} - 1. \]

Finally, since \( \theta_s > \theta_b \), \( \beta^S > \beta^D \). Q.E.D.

An intuitive way to think about Proposition 1 is that a diversified firm offers two sets of cash flows: (i) the cash flow of two stand-alone firms, and (ii) an additional coinsurance cash flow whose beta,

\[ \beta^{CI} = C^{CI}_g \mid C^{CI}_b - 1, \]

is lower than that of stand-alone firms. This is because the relative probability of avoiding deadweight costs is inversely related to the state of the economy, namely, \( (1 - \theta_s) \) in the good state and \( (1 - \theta_b) \) in the bad state – that is, deadweight costs are partly systematic. Since \( \beta^D \) is a weighted average of \( \beta^S \) and \( \beta^{CI} \), it follows that \( \beta^D \) must be lower than \( \beta^S \) as long as the probability of coinsurance is not zero.

The intuition above also indicates that the way in which coinsurance reduces systematic risk is not specific to our model. A sufficient (though not necessary) condition for our results to hold in a general \( N \)-state economy with states indexed by \( w \in \{1, \ldots, N\} \) is that for any two states \( w' \) and \( w'' \) with stochastic discount factor values \( m(w') \leq m(w'') \), \( \theta(w') \) is greater than or equal to \( \theta(w'') \), and for at least one pair \( m(w') < m(w'') \), \( \theta(w') \) is greater than \( \theta(w'') \). In other words,
our results hold as long as the probability of a high outcome (deadweight loss) increases (decreases) in the state of the economy represented by the value of the stochastic discount factor.

**Combining Two Stand-alone Firms with Correlated Idiosyncratic Risks**

We now turn to the possibility that idiosyncratic risks may be correlated by modeling the structure of the correlation. Let $\rho \in [\rho, 1]$ represent the correlation of idiosyncratic risks in both states of the economy $e \in \{g, b\}$. Then we have:

\[
\begin{align*}
    p_{hh,e} &= \theta_e (\theta_e + \rho(1-\theta_e)) \\
    p_{lh,e} &= (1-\theta_e)(\theta_e - \rho\theta_e) \quad (= p_{hl,e}) \\
    p_{hl,e} &= \theta_e (1-\theta_e - \rho(1-\theta_e)) \quad (= p_{lh,e}) \\
    p_{ll,e} &= (1-\theta_e)(1-\theta_e + \rho\theta_e).
\end{align*}
\]

These probabilities always add up to 1, and individually always fall between 0 and 1 in the specified region of $\rho$ where

\[
\rho = \max \left\{ -\frac{\theta_g}{1-\theta_g}, -\frac{1-\theta_g}{\theta_g}, -\frac{\theta_b}{1-\theta_b}, -\frac{1-\theta_b}{\theta_b} \right\}.
\]

In addition, joint probabilities are consistent with marginal probabilities.

\[
\begin{align*}
    \theta_e &= p_{hh,e} + p_{hl,e} = p_{hh,e} + p_{lh,e} \\
    1-\theta_e &= p_{lh,e} + p_{ll,e} = p_{hl,e} + p_{ll,e}
\end{align*}
\]

The case in which $\rho$ equals 0 corresponds to the case of independence in Proposition 1. When $\rho$ equals 1 (perfect correlation),

\[
p_{hh,e} = \theta_e, \quad p_{lh,e} = p_{hl,e} = 0, \quad p_{ll,e} = (1-\theta_e).
\]
The case of perfect correlation for a diversified firm represents a doubling of scale without any coinsurance effect.

**Proposition 2** The systematic risk and cost of capital of a diversified firm (combining two stand-alone firms with positive systematic risk) increase in $\rho$, and reach those of a stand-alone firm in the limit when $\rho$ equals 1.

**Proof.** Given Lemma 1, it suffices to show that $\partial \beta^D / \partial \rho > 0$ and $\beta^D = \beta^S$ when $\rho$ equals 1.

Using the new probability structure,

$$
\beta^D = \frac{\theta_s (\theta_s + \rho(1-\theta_s))(2H) + 2\theta_b (1-\theta_b)(1-\rho)(H+L)}{\theta_b (\theta_b + \rho(1-\theta_b))(2H) + 2\theta_b (1-\theta_b)(1-\rho)(H+L)} - 1
$$

$$
= \frac{2\theta_s H + 2\theta_s (1-\theta_s)(1-\rho)L}{2\theta_b H + 2\theta_b (1-\theta_b)(1-\rho)L} - 1.
$$

Simple algebra shows that

$$
\frac{\partial \beta^D}{\partial \rho} = \frac{4HL\theta_s \theta_b (\theta_s - \theta_b)}{(2\theta_b H + 2\theta_b (1-\theta_b)(1-\rho)L)^2} > 0.
$$

Also, when $\rho$ equals 1, coinsurance cash flows drop out of $\beta^D$, and $\beta^D$ equals $\beta^S$. Q.E.D.

Proposition 2 demonstrates how a diversified firm with a higher level of coinsurance should have a lower cost of capital compared to a portfolio of stand-alone firms. Since $\beta^D$ is a weighted average of $\beta^S$ and $\beta^{CI}$, and the weight on $\beta^{CI} (< \beta^S)$ is directly proportional to $(1-\rho)$, $\beta^D$ is always less than or equal to $\beta^S$, increases in $\rho$, and eventually reaches $\beta^S$ when $\rho$ equals 1. Propositions 1 and 2 consider the case of identical stand-alone firms. These results
generalize to the case in which stand-alone firms have different positive betas. Given that most businesses have positive betas, the main message of our model covers a wide range of situations.

2.2.2. Agency Problems and Inefficient Transfers as Costs of Integration

Thus far we have assumed that the merged firm operates with an efficient internal capital market that is free of agency problems. In this subsection we relax these assumptions.

In particular, suppose that diversification brings not only coinsurance benefits, but also integration costs in the form of agency problems and inefficient transfers. Indeed, such costs underlie the main conjecture of previous work showing that diversified firms have lower valuations relative to stand-alone firms (Lang and Stulz (1994), Berger and Ofek (1995)). For instance, the agency costs arising from empire building (Jensen (1986)), entrenched managers (Shleifer and Vishny (1989)), inefficient allocation of resources (Shin and Stulz (1998) and Rajan, Servaes, and Zingales (2000)), and cross-subsidization (Scharfstein and Stein (2000)) can lead to lower cash flows.\footnote{Subsequent research questions the view that diversified firms are less productive (Schoar (2002)) or that they allocate resources less efficiently than stand-alone firms (Maksimovic and Phillips (2002)).} These costs can be seen as closing our model to prevent the counterfactual prediction that the entire economy should be owned by one big firm to maximize coinsurance benefits.

It is also worth noting that with integration costs, our predictions can be consistent with either a diversification discount or premium. Specifically, recall that in the basic model diversified firms have not only lower cost of capital, but also higher cash flows compared to portfolios of stand-alone firms. Therefore, our model implies that diversified firms have higher valuations, a prediction that is inconsistent with a large body of empirical work showing that
diversified firms have lower valuations on average. While recent work has challenged the interpretation that diversification leads to lower valuation, the debate is far from settled, and importantly, is not the focus of our paper. Incorporating integration costs into the model allows us to accommodate both valuation possibilities. The extension of the basic model with integration costs is presented below.

Let $A_e^D$ denote the fraction of firm cash flow that is wasted due to rent-seeking activities and inefficient transfers depending on the state of the economy $e \in \{g, b\}$. Then, a diversified firm’s cash flow net of integration costs is given by

$$C_e^{D/A} = C_e^D (1 - A_e^D) \quad \text{for} \quad e \in \{g, b\}.$$  

Whether integration costs increase or decrease systematic risk depends on the relative magnitudes of $A_g^D$ and $A_b^D$. If $A_g^D = A_b^D$, then integration costs do not affect systematic risk beyond reducing firm value. If $A_g^D > A_b^D$, say because bad times discipline managers and survival concerns necessitate efficiency, then integration costs reduce firm beta,

$$\frac{C_g^D (1 - A_g^D)}{C_b^D (1 - A_b^D)} - 1 < \frac{C_g^D}{C_b^D} - 1 \Rightarrow \beta^{D/A} < \beta^D < \beta^S,$$

and add to the coinsurance effect.

Alternatively, if $A_g^D < A_b^D$, then integration costs increase firm beta. However, $A_b^D$ would have to be sufficiently higher than $A_g^D$ for integration costs to overturn the effect of coinsurance $\beta^{CI} (< \beta^S)$ in this case.
2.2.3. Debt Financing

In this subsection, we show that our results extend to debt financing. To see this, suppose that a diversified firm comprises two stand-alone firms, each with a face value of debt $K = H - \Delta$. Further suppose that $K$ is high enough. Specifically, $0 < \Delta < (H - L)/2$ so that $(H + L)/2 < K < H$. Then, depending on the state of the economy $e \in \{g, b\}$, stand-alone bondholders with face value $K$ receive

$$B^S_g = \theta_g (H - \Delta)$$
$$B^S_b = \theta_b (H - \Delta)$$

whereas diversified firm bondholders with face value $2K$ receive

$$B^D_g = \theta^2_g (2(H - \Delta)) + 2\theta_g (1 - \theta_g)(H + L)$$
$$B^D_b = \theta^2_b (2(H - \Delta)) + 2\theta_b (1 - \theta_b)(H + L).$$

Using the expected cash flows above to compute bond betas,

$$\beta^S_B = \frac{\theta_g (H - \Delta)}{\theta_b (H - \Delta)}$$
$$\beta^D_B = \frac{2\theta_g (H - \Delta) + 2\theta_b (L + \Delta - \theta_g (L + \Delta))}{2\theta_b (H - \Delta) + 2\theta_b (L + \Delta - \theta_b (L + \Delta))}.$$ 

Similar to the main model, diversified firm bondholders receive two sets of cash flows whose overall beta is less than the beta of cash flows to stand-alone bondholders. As a result, $\beta^D_B < \beta^S_B$, and the cost of debt for a diversified firm comprising two stand-alone firms is lower than the cost of debt for the two stand-alone firms. In our model, diversified firms enjoy coinsurance benefits that reduce their systematic risk, and as this extension shows, these benefits reduce their cost of debt as well.
2.2.4. **State-contingent Deadweight Loss**

We have assumed so far that $L$, the deadweight loss suffered by stand-alone firms, does not depend on the state of the economy. If, in contrast, such costs were to depend on the state of the economy $e \in \{g, b\}$, our results would continue to hold as long as the beta of coinsurance cash flows,

$$\beta^{CI} = \frac{\theta_g (1-\theta_g) L_g}{\theta_b (1-\theta_b) L_b} - 1,$$

remains less than $\beta^S$.

For instance, if supplier and customer defections are probabilistic and these probabilities are higher during bad times than during good times, then $L_g < L_b$ and we may have understated the coinsurance effect by assuming a state-independent $L$. Defection probabilities may indeed be higher during bad times than during good times if suppliers and customers think that the firm is more likely to forgo important investments due to the greater wedge between internal and external finance during bad times.

2.3. **Testable Predictions**

Our model lends itself to two novel testable predictions about the coinsurance effect of corporate diversification on the total cost of capital. We now state these predictions.

**Prediction 1** A diversified firm, *on average*, has a lower total cost of capital than a portfolio of comparable stand-alone firms.

Prediction 1 follows from Propositions 1 and 2. In our model, a diversified firm is able to avoid deadweight costs that stand-alone firms cannot avoid on their own. The resulting
coinsurance cash flows tend to have lower systematic risk than the underlying stand-alone assets, and this in turn reduces the total cost of capital that investors provide to diversified firms.

**Prediction 2** A diversified firm comprised of businesses with less correlated cash flows has a lower total cost of capital.

Prediction 2 follows from Proposition 2, and provides a cross-sectional test. Because the extent of coinsurance is greater for diversified firms comprised of businesses with less correlated cash flows, investors demand less compensation for providing capital to such firms. In the limit where a firm’s different businesses have perfectly correlated cash flows, there are no coinsurance cash flows and therefore no effect on the total cost of capital.

In the empirical work that follows, we test our model’s predictions using not only the correlation of cash flows, but also the correlation of investment opportunities of the segments comprising a diversified firm. The idea is that coinsurance may lower systematic risk through internal capital markets that help firms avoid deadweight costs of external financing by channeling resources to business units with superior investment opportunities (Matsusaka and Nanda (2002)).

### 3. Research Design

In Section 3.1, we first discuss our ex ante measures of the implied cost of equity capital. We then present the underlying valuation model and the empirical implementation of the model. Next we describe the construction of our excess cost of capital measure. In Section 3.2, we discuss our measures of coinsurance.
3.1. Implied Cost of Capital

Prior research in finance has generally used ex post realized returns to proxy for expected returns (e.g., Fama and French (1997), Lamont and Polk (2001)). One shortcoming of this approach is that realized returns are noisy proxies for expected returns due to contamination by information shocks.\(^\text{12}\) To address this concern, recent literature in accounting and finance has developed an alternative approach to measuring ex ante returns by estimating the implied cost of capital (e.g., Claus and Thomas (2001), Gebhardt, Lee, and Swaminathan (2001), Easton (2004), Ohlson and Juettner-Nauroth (2005)). The implied cost of capital is the internal rate of return that equates the current stock price to the present value of all expected future cash flows. The expected future cash flows are usually estimated using analysts’ earnings forecasts. In general, these implied cost of capital measures differ in terms of the form of the valuation model and the assumptions regarding terminal value computation.\(^\text{13}\)

In our main analysis, we follow the approach of Gebhardt, Lee, and Swaminathan (2001) (hereafter, GLS) in estimating the implied cost of equity. The GLS measure has been successfully employed in several asset-pricing contexts (e.g., Lee, Ng, and Swaminathan (2007), Pastor, Sinha, and Swaminathan (2008)). We also perform sensitivity tests using two alternative implied cost of equity measures based on Claus and Thomas (2001) and Easton (2004). See Section 5.2.3 for a more detailed discussion.

\(^{12}\) Elton (1999) provides a detailed discussion of the shortcoming of using realized returns as a proxy for expected returns.

\(^{13}\) A discussion of the relative advantages of each method is outside the scope of this paper. Prior research evaluates alternative empirical measures of implied cost of equity and reaches different conclusions on their relative merits and demerits (e.g., Guay et al. (2005), Easton and Monahan (2005), Botosan and Plumlee (2005)).
3.1.1. Valuation Model for Cost of Equity (GLS)

The GLS measure is based on the residual income valuation model, which is derived from the discounted dividend model with an additional assumption of clean-surplus accounting.\(^\text{14}\)

In the model, the value of the firm at time \(t\) is equal to

\[
P_t = B_t + \sum_{i=1}^{\infty} \frac{E_i[N_{t+i} - r_e B_{t+i-1}]}{(1 + r_e)^i},
\]

where \(P_t\) is the market value of equity at time \(t\), \(B_t\) is the book value of equity at time \(t\), \(N_{t+i}\) is net income at time \(t+i\), and \(r_e\) is the implied cost of equity. We assume a flat term structure of interest rates.

GLS further restate the model in terms of ROE, and assume that ROE for each firm reverts to its industry median over a specified horizon. Beyond that horizon, the terminal value is calculated as an infinite annuity of residual ROE,

\[
P_t = B_t + \sum_{i=1}^{T} \frac{FROE_{t+i} - r_e B_{t+i-1}}{(1 + r_e)^i} + \frac{FROE_{t+T} - r_e}{r_e(1 + r_e)^T} B_{t+T-1},
\]

where \(B_{t+i}\) is book value per share estimated using a clean-surplus assumption \((B_{t+i} = B_{t+i-1} - k*FEPS_{t+i} + FEPS_{t+i})\), where \(k\) is the dividend payout ratio and \(FEPS_{t+i}\) is the analyst earnings per share forecast for year \(t+i\), \(FROE_{t+i}\) is future expected return on equity, which is assumed to fade linearly to industry median from year 3 until year \(T\), and all other variables are as defined previously.

3.1.2. Empirical Estimation

Implied Cost of Equity

\(^{14}\) Under the clean-surplus assumption, book value of equity at \(t+1\) is equal to book value of equity at \(t\) plus net income earned during \(t+1\) minus net dividends paid during \(t+1\).
As in GLS, we assume that the forecast horizon, T, is equal to 12 years. We use median consensus forecasts to proxy for the market’s future earnings expectations and require that each observation have non-missing one- and two-year-ahead consensus earnings forecasts ($FEPS_{t+1}$ and $FEPS_{t+2}$) and positive book value of equity. We use three-year-ahead forecasts for future earnings per share, if they are available in I/B/E/S. Otherwise, we estimate $FEPS_{t+3}$ by applying the long-term growth rate to $FEPS_{t+2}$. We use stock price per share and forecasts of both EPS and long-term earnings growth from the I/B/E/S summary tape as of the third Thursday in June of each year. Book value of equity and the dividend payout ratio for the latest fiscal year-end prior to each June are obtained from the Compustat annual database. We assume a constant dividend payout ratio throughout the forecast period. For the first three years, expected ROE is estimated as $FROE_{t+i} = FEPS_{t+i} / B_{t+i-1}$. Thereafter, $FROE$ is computed by linear interpolation to the industry median $ROE$ (we use Fama and French (1997) industry definitions). The cost of equity is calculated numerically employing the Newton-Raphson method. We set the initial value of the cost of equity to 9% in the first iteration; the algorithm is considered to converge if the stock price obtained from the implied cost of equity deviates from the actual stock price by no more than $0.005.

Cost of Capital

Our model predicts that coinsurance reduces systematic risk and hence the total cost of capital. Accordingly, our empirical analyses are based on a weighted-average cost of capital (COC) estimate. To compute this estimate, we follow an approach similar to Lamont and Polk (2001). In particular, Lamont and Polk define total cost of capital as the weighted average of a firm’s realized equity return and the return on an aggregate bond index. While we also utilize an

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15 Book value of equity is Compustat Item #60; the dividend payout ratio is computed as dividends (Compustat Item #21) divided by earnings (Compustat Item #237). If earnings is negative, then the dividend payout ratio is computed as dividends over 6% of total assets (Compustat Item #6).
aggregate bond index to estimate the cost of debt, we use ex ante measures of the implied cost of
equity (i.e., the GLS measure) instead of realized equity returns to proxy for expected equity
returns, and we use bond yields instead of realized bond returns to proxy for expected debt
returns. More specifically, the COC for each firm \( i \) and year \( t \) is computed as follows:

\[
COC_{i,t} = D_{i,t-1} Y_t^{Dj} + (1 - D_{i,t-1}) COEC_{i,t},
\]

where \( Y_t^{Dj} \) is the aggregate bond yield from the Barclays Capital Aggregate Bond Index,\(^{16}\)
\( COEC_{i,t} \) is the implied cost of equity (GLS), and \( D_{i,t-1} \) is a firm’s book value of debt divided by
the total value (book value of debt plus market value of common equity).\(^{17}\)

This cost of capital measure has two limitations. First, our proxy for the cost of debt only
measures the aggregate bond yield and hence it does not capture any firm-specific variation in
expected debt returns.\(^{18}\) To the extent that coinsurance reduces the cost of debt (as our model
predicts), our results understate the coinsurance effect on cost of capital. Second, we use the
book value of debt to calculate leverage. If a firm’s value of debt has increased (decreased) over
time, its book value of debt would understate (overstate) leverage and overstate (understate) cost
of capital (since cost of debt is on average lower than cost of equity). Ex ante we do not suspect
any systematic relation between prior debt value changes and the degree of coinsurance.
Therefore, we do not expect that using book value of debt would systematically bias our results.
Despite these limitations, our measure of total cost of capital is conceptually superior to one that

\(^{16}\) The Barclays Capital Aggregate Bond Index was formerly known as the Lehman Brothers Aggregate Bond Index.
\(^{17}\) Book value of debt is Compustat Item #9; market value of equity is estimated as fiscal year-end stock price
(Compustat #199) times shares outstanding (Compustat Item #25).
\(^{18}\) Using firm-specific bond yields to proxy for the cost of debt is not without limitation. In particular, the yield of a
bond is defined as the discount rate that equates the bond’s promised repayment (as opposed to its expected future
cash flow) to its market price. Hence, it captures both the firm’s exposure to systematic risk and its exposure to
idiosyncratic risk (non-systematic risk of default). Given that coinsurance may reduce not only the systematic risk
component (as our model predicts), but also the non-systematic risk of default (Lewellen (1971)), it would be
difficult to disentangle the two coinsurance effects and hence attribute our results (i.e., the positive association
between cost of capital and cross-segment correlations) solely to the former if we construct our cost of capital
measure using bond yields as a proxy for expected cost of debt.
measures only the cost of equity capital, because it takes into consideration the importance of debt in a firm’s capital structure.

3.1.3. Excess Cost of Capital

Our tests compare the COC of a diversified firm to the as-if (imputed) COC of a portfolio of its segments as stand-alone businesses. For that purpose, we compute the excess COC as a natural logarithm of the ratio of a firm’s COC to its imputed COC. The excess COC that is lower (greater) than zero is consistent with diversification reducing (increasing) the firm’s cost of capital. The magnitude of excess COC can be interpreted as a logarithmic percentage difference between the firm’s actual COC and its imputed counterpart.

The firm’s imputed COC is calculated as a weighted average of the imputed COC of its segments (the calculation is equivalent to computing the value-weighted return on a portfolio of stocks):

$$iCOC_i = \sum_{k=1}^{n} \frac{iMV_{ik} \cdot iCOC_{ik}}{\sum_{k=1}^{n} iMV_{ik}},$$

where $n$ is the number of the firm’s segments, $iCOC_{ik}$ is the imputed COC of segment $k$, equal to the segment’s industry median COC, and $iMV_{ik}$ is the imputed market value of segment $k$, calculated as in Berger and Ofek (1995).

The procedure of estimating segments’ imputed market values is described in detail in Berger and Ofek (1995). In short, the estimation consists of two steps: (1) estimating the median ratio of total capital to sales for all single-segment firms in the industry to which the segment belongs, and (2) multiplying the segment’s sales with the median industry ratio. The definition of
industry is based on the narrowest SIC grouping that includes at least five single-segment firms with at least $20 million in sales and has a non-missing COC estimate.

3.2. Coinsurance Measures: Cross-segment Correlations

Our model calls for a measure of coinsurance between the firm’s segments, which in our model is the idiosyncratic correlation among the segments’ future free cash flows. A precise measurement of coinsurance, however, is difficult to obtain because the distribution of segments’ future free cash flows is not observable. Moreover, using historical data at the segment level to estimate coinsurance is also problematic because firm composition is usually not constant over time. Accordingly, we construct two empirical proxies of coinsurance based on historical industry-level data. To ensure that the correlations do not contain systematic risk, the computation is performed in two stages.

First, for each 2-digit SIC code industry and year, we compute the industry-level idiosyncratic cash flows (investments) as residuals from regressing average industry cash flows (investments) on average market-wide cash flows (investment).\textsuperscript{19} The averages are computed using all single-segment firms within each industry.

Next, for each year, we estimate correlations between every possible pair of industry-level idiosyncratic cash flows (investments) using the preceding ten-year period. We then use these correlation estimates to construct our cash flow (investment) coinsurance measures.

Specifically, we compute a sales-weighted correlation measure $\rho^i_{y(n)}$ for firm $i$ in year $y$ with $n$ business segments as

\textsuperscript{19} Cash flow is defined as operating income before depreciation and investment is measured by capital expenditures.
\[
\sum_{s=1}^{n} \frac{\sum_{i=1}^{n} Sales_{i(s(j)}}}{\sum_{i=1}^{n} Sales_{i}} \frac{\sum_{i=1}^{n} Sales_{i(k)}}{\sum_{i=1}^{n} Sales_{i}} Corr_{[y-10,y-1]}(j,k),
\]

where \( Sales_{i(s(j)}} \) is the sales of firm \( i \)'s business segment \( s \) operating in industry \( j \) (similarly for business segment \( t \) operating in industry \( k \)), and \( Corr_{[y-10,y-1]}(j,k) \) is the idiosyncratic correlation of average industry cash flows or investments between industries \( j \) and \( k \) over the ten-year period before year \( y \), estimated using single-segment firms. We obtain quantitatively similar results using an alternative coinsurance measure, which further includes the standard deviation of average industry cash flows and investments (Duchin (2008)).

Note that a single-segment firm’s sales-weighted cash flow or investment correlation measure equals one by definition. This is also true for a multi-segment firm whose segments operate in the same industry.

4. Sample and Data

4.1. Sample Selection

We obtain our sample from the intersection of the Compustat and I/B/E/S databases for the period 1988 to 2006.\(^\text{20}\) We construct cost of capital measures by combining firm-level accounting information from the Compustat annual files with analyst forecasts from I/B/E/S. The excess cost of capital measures and the coinsurance measures require availability of segment disclosures from the Compustat segment-level files.

Additionally, we impose the following restrictions on our dataset used in the main analyses. First, we follow Berger and Ofek (1995) and require that (1) all firm-years have at least

\(^{20}\) The start of our sample period in 1988 is determined by our use of cross-segment correlation estimates based on prior ten-year single-segment data, which start in 1978.
$20 million in sales to avoid distorted valuation multiples; (2) the sum of segment sales must be within 1% of total sales of the firm to ensure the integrity of segment data; (3) all of the firm’s segments for a given year must have at least five firms in the same 2-digit SIC code industry with non-missing capital-sales ratios and GLS COC estimates; and (4) all firms with at least one segment in the financial industry (SIC codes between 6000 and 6999) are excluded from the sample. Second, we require the following data to estimate the GLS COC measure: (1) one- and two-year-ahead earnings forecasts; (2) either a three-year-ahead earnings forecast or the long-term growth earnings forecast and a positive two-year-ahead earnings forecast; and (3) positive book value of equity. The full sample with available GLS excess cost of capital estimates consists of 38,369 firm-year observations, of which 26,454 (11,915) observations pertain to single-segment (multi-segment) firms. The sample used in the cross-sectional analyses is further constrained by the availability of control variables. We discuss our control variables in the next subsection.

4.2. Control Variables for Cross-sectional Analysis

We include the following sets of control variables in our cross-sectional regression analysis.

Market Anomalies

To ensure that our results are distinct from the well-documented asset-pricing anomalies (Fama and French (1992) and Jegadeesh and Titman (1993)), we control for size, book-to-market, and momentum as proxied by the log of market capitalization, the book-to-market ratio, and lagged buy-and-hold returns over the past 12 months, respectively. Including a measure of momentum also controls for sluggishness in analyst forecasts. In particular, forecast sluggishness
can induce a negative relation between recent returns and the cost of capital measures. Recent revisions in the stock market’s earnings expectations, although immediately reflected in stock prices, may not be incorporated in analyst forecasts on a timely basis. In the case of an upward (downward) revision reflected in positive (negative) stock returns, earnings forecasts will be biased downwards (upwards), leading to downwardly (upwardly) biased implied cost of equity estimates.\footnote{It is possible that we are “over-controlling” for other factors by including size and the book-to-market ratio in our regressions. First, book-to-market may be associated with forward-looking betas from the conditional asset-pricing model (e.g. Petkova and Zhang (2005)). Since we hypothesize that coinsurance affects cost of capital by affecting forward-looking betas, we may be “throwing the baby out with the bath water” by including book-to-market as a control variable. Second, size may serve as an alternative proxy for the extent of coinsurance. Larger firms are likely to have a larger number of unrelated projects, which can lead to a higher degree of coinsurance.}

In addition, we include I/B/E/S’s long-term growth forecast to control for the anomaly documented by LaPorta (1996). In particular, LaPorta finds that forecasted long-term growth in earnings is negatively associated with average realized returns. This could lead to a negative relation between the long-term growth forecast and the implied cost of equity measure. On the other hand, the implied COC measures may be mechanically positively related to the long-term growth forecasts. Understating (overstating) expected long-term growth would lead to understating (overstating) the implied COC, and in turn to a positive relation between the two measures.

**Analyst Forecast Dispersion**

We control for dispersion in analysts’ forecasts, as measured by the log of the standard deviation in analyst forecasts. Gebhardt et al. (2001) show that the GLS COC measure is positively correlated with dispersion in analysts’ forecasts. If forecast dispersion is a measure of analyst disagreement, we expect it to increase in the degree of diversification or the degree of unrelatedness of business segments due to a higher complexity of the forecasting task. A higher
degree of coinsurance will therefore be associated with higher forecast dispersion and higher cost of capital. Failure to control for forecast dispersion could bias our tests against finding the cost of capital effect of coinsurance.

**Leverage**

Finally, we control for leverage. Theory predicts that the weighted average cost of capital declines with leverage due to tax-shield benefits of debt, and that the cost of equity increases with leverage due to increased financial risk. To the extent that firms with greater cross-segment coinsurance take on more debt (Lewellen (1971)), our results are affected. We therefore control for these leverage effects.

**Variable Definitions**

We summarize the definitions of the control variables below:

- **Log(market capitalization)** = Natural logarithm of fiscal year-end stock price times shares outstanding from Compustat (#199 * #25);\(^{22}\)

- **Leverage** = Book value of long-term debt divided by the sum of the book value of long-term debt and the market value of equity from Compustat (#9 / (#9 + #199 * #25));

- **Book-to-market** = Ratio of book value of equity to market value of equity from Compustat (#60 / (#199 * #25));

- **Log(forecast dispersion)** = Natural logarithm of the standard deviation in analysts’ one-year-ahead earnings forecasts from I/B/E/S;

- **Long-term growth forecast** = Consensus (median) long-term growth forecast from I/B/E/S;

- **Lagged 12-month return** = Buy-and-hold return on the firm’s stock from the beginning of June \((t-1)\) until the end of May \((t)\), estimated using monthly returns from CRSP.

\(^{22}\) All numbered items refer to the Compustat annual database.
The timeline of the variable measurement is depicted in Figure 1. Note that these additional data requirements constrain our sample to 29,153 observations, of which 20,046 (9,107) observations pertain to single-segment (multi-segment) firms. Some of the sensitivity analyses impose further data restrictions on the sample, as discussed in the corresponding sections of the paper.

5. Empirical Results

In this section, we first present summary statistics for our main variable of interest: excess cost of capital (excess COC). We then test our main hypothesis – the coinsurance effect of diversification on cost of capital – by examining the relation between excess COC and our measures of cross-segment cash flow/investment correlations and we present results from both univariate and multivariate tests. Following our main empirical analysis, we present results from various robustness tests.

5.1. Summary Statistics: Excess Cost of Capital

Recall that a diversified firm’s excess COC measures the extent to which the firm’s cost of capital is higher or lower than the sum of the imputed cost of capital from its segments as stand-alone firms. On average, we expect diversified firms to have a lower cost of capital relative to portfolios of comparable stand-alone firms (Prediction 1).

In Table 1, we present summary statistics for multi- and single-segment firms separately. For the multi-segment subsample, both mean and median excess COC are negative and significant (-0.040 and -0.025). For the single-segment subsample, the median excess COC is zero by construction because the imputed COC values are calculated based on industry medians, but the reported figure is different from zero due to additional sample restrictions. The mean
excess COC is negative and significant, suggesting that the distribution of excess COC is negatively skewed. Accordingly, we use the mean excess COC for single-segment firms as a benchmark for evaluating the excess COC of multi-segment firms. The difference in means between the single- and multi-segment subsamples is negative and significant (0.01 at $p<0.01$), suggesting that the cost of capital of diversified firms is on average 1% lower than that of comparable portfolios of stand-alone firms.\footnote{Throughout the paper, we imply logarithmic percentages whenever we discuss percentage differences. For small percentage values, logarithmic percentages and absolute percentages are approximately the same.} The modest result is likely due to the pooling of all multi-segment firms, of which many operate within a single industry and thus enjoy little cross-segment coinsurance. In the next section, we investigate whether the cost of capital effect from diversification varies with the extent of cross-segment coinsurance.

5.2. Cross-sectional Analysis of Cost of Capital and Coinsurance

In this section, we explore the cross-sectional relation between excess COC and our measures of coinsurance as proxied by cross-segment industry-level cash flow and investment correlations. A higher (lower) correlation means lower (greater) coinsurance.

5.2.1. Univariate Analysis

Table 2 reports results from the univariate analysis. We sort our sample of multi-segment firms into quintiles based on the two coinsurance measures and compute the average excess COC for each portfolio. The results for the cash flow (investment) correlation portfolios are reported in the left (right) panel. We also present results for the single-segment firms. Note that the single-segment firms can be viewed as the extreme observations with respect to the degree of
coinsurance – stand-alone firms by definition have zero coinsurance (i.e., cash flow and investment correlations are equal to one by definition).  

Because the two sets of portfolio results are very similar, we focus our discussion on those from the cash flow correlation portfolios. Consistent with the coinsurance hypothesis, we observe a monotonic increase in excess COC from the highest coinsurance firms (Q1) to the lowest coinsurance firms (Q5). The mean difference between Q5 and Q1 is a statistically significant 3.2%, i.e. diversification leads to a cost of capital reduction that is 3.2% higher in magnitude for firms in the highest coinsurance quintile (Q1) compared to firms in the lowest coinsurance quintile (Q5). Similarly, the mean difference between the cost of capital of single-segment firms and firms in the highest coinsurance quintile (Q1) is 2.9%, suggesting a significant coinsurance effect from diversification. Overall, these results are consistent with Prediction 2 – diversified firms comprised of businesses with less correlated cash flows have lower total cost of capital.

5.2.2. Multivariate Analysis

In this subsection, we estimate the cross-sectional relation between excess COC and our measures of coinsurance after controlling for the set of firm characteristics discussed in Section 4.2. Before we turn to the results reported in Table 3, we first discuss our regression specifications.

In the first set of regressions, Models 1 and 2, we regress excess COC on our coinsurance measures, cash flow and investment correlations respectively, and control for all variables except

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Note that conceptually coinsurance can also be present in single-segment firms. Our measures of coinsurance are constrained by the use of industry segment definition as a measure of diversity. There are other dimensions of diversity (such as product line diversity) that can create coinsurance even in single-segment firms. See Section 5.2.2 for a more detailed discussion.
for the number of segments (NUMSEG) and the log of market capitalization (SIZE). We exclude NUMSEG and SIZE because they are likely to capture some degree of coinsurance. Larger firms or firms with more segments are more likely to have business units with imperfect cash flow correlations. Therefore, including these two measures in the regressions would over control for the coinsurance effect from the correlation measures.

In the second set of regressions, Models 3 and 4, we use NUMSEG and SIZE respectively, as alternative measures of coinsurance. As discussed earlier, NUMSEG and SIZE are both measures of firm size, and hence they are likely to capture not only the extent of coinsurance that is reflected in cash flow/investment correlations, but also some degree of coinsurance that is not captured by the correlation measures. In particular, cash flow and investment correlations measure the coinsurance effect arising from diversity at the segment level (imperfect cash flow correlations across business segments). Conceptually, however, coinsurance can be present even in single-segment firms if these firms have diverse product lines generating imperfectly correlated cash flows. Since the number of product lines is likely to be positively related to firm size and the number of segments, we use NUMSEG and SIZE as an alternative proxy for coinsurance in Models 3 and 4. To the extent that these measures capture some degree of coinsurance, our hypothesis predicts a negative association.

In the last set of specifications, Models 5 and 6, we include all control variables, including NUMSEG and SIZE. While NUMSEG and SIZE might capture some degree of coinsurance, it is difficult to disentangle other possible “size effects” from the coinsurance effect. Also, the cash flow and investment correlation measures are direct proxies for the coinsurance effect as they directly take into consideration the correlation of cash flows/investments across
segments. We therefore view this last specification as the most demanding test of our correlation measures and also of our coinsurance hypothesis.

The results from the three sets of regression specifications are presented in Table 3. The robust standard error for each variable (heteroskedasticity consistent and double clustered by firm and year (Petersen (2008)) is reported in brackets below its corresponding coefficient. Because the results across the two correlation measures are qualitatively and statistically similar, we focus our discussion on the cash flow correlation regressions.

Consistent with the univariate test results, the coefficient on cash flow correlations is positive and significant in both Models 1 and 5 (with p<0.01). Even after controlling for NUMSEG and SIZE, the coefficients on cash flow correlations remain positive and significant, suggesting that greater coinsurance (lower cash flow correlations) is associated with lower cost of capital. In Models 3 and 4, we find a negative and significant coefficient on NUMSEG and SIZE, respectively, with p<0.10 and p<0.01. This result suggests that larger firms and firms with more segments, which tend to have more product lines, tend to have a lower cost of capital. As noted earlier, while this result is consistent with the coinsurance hypothesis, it is difficult to attribute the finding solely to the coinsurance effect as size may also proxy for other factors (e.g., information environment) that can affect the cost of capital. We therefore draw inferences primarily from our main regression specifications (i.e., Models 5 and 6).

Overall, our univariate and multivariate test results are consistent with our second prediction: firms with lower cross-segment cash flow correlations have lower cost of capital, i.e., the coinsurance effect increases as cross-segment cash flow correlation decreases. In the following subsections, we provide several sensitivity tests to ensure that the results from our

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25 Note that to the extent that coinsurance is present across different product lines (or any other dimension of diversity that is finer than industry segments), our correlation measures understate the effect of coinsurance.
main cross-sectional analysis are robust to excluding single-segment firms, controlling for analyst forecast errors, using alternative measures of implied cost of equity, and ignoring the firm’s capital structure.

5.2.3. Robustness Tests

Excluding Single-segment Firms

Our main regression analysis in the previous subsection includes both single- and multi-segment firms. Prior research that examines the diversification discount generally also includes stand-alone firms as a benchmark (e.g., Berger and Ofek (1995)). In the context of our study, including single-segment firms yields a more powerful test as single-segment firms by definition do not have any cross-segment coinsurance. Nevertheless, to ensure that our results are not spuriously driven by other differences between stand-alone and diversified firms, we perform our main analysis on a subsample of multi-segment firms. The results, reported in Table 4, are qualitatively and statistically similar to those reported in Models 5 and 6 of Table 3. In particular, the coefficients on cash flow and investment correlations are both positive and significant (at p<0.01). This finding suggests that our results on coinsurance and cost of capital are not driven by differences across single- and multi-segment firms.

Analyst Forecast Errors

A potential limitation of the implied cost of equity measures is the measurement error arising from the bias in analyst forecasts. To address this concern, we control for one-year and two-year-ahead unexpected and expected forecast errors in our main regression models. In particular, we follow Ogneva, Subramanyam, and Raghunandan (2007) and estimate expected forecast errors using the prediction model in Liu and Su (2005). Our parsimonious version of the
model includes the following predictors that proxy for systematic biases in analyst forecasts: (1) past stock returns, (2) recent analyst earnings forecast revisions, and variables related to overreaction to past information, namely, (3) forward earnings-to-price ratio, (4) long-term growth forecast, (5) investments in property, plant, and equipment. Estimation of the predicted forecast error is performed in two steps. First, we regress realized forecast errors from the previous period on the set of predictors. Second, we combine the coefficients derived from the first step with current values of predictors to arrive at the predicted forecast error. Estimation is performed separately for one- and two-year-ahead forecast errors. Unexpected forecast errors are computed as the difference between realized errors and their predicted component. Because one-year and two-year-ahead expected errors are highly collinear, we use the average expected errors over the two years as the control measure. The results, reported in Table 5, continue to show a positive and significant coefficient on cash flow and investment correlations, suggesting that the negative relation between coinsurance and cost of capital is unlikely to be driven by systematic differences in analyst forecast biases between single- and multi-segment firms.

*Alternative Measures of Implied Cost of Equity Capital*

In our main analysis, we estimate implied cost of equity (COE) using the approach of Gebhardt, Lee, and Swaminathan (2001) and Lee, Ng, and Swaminathan (2007) – see Section 3.1. In this subsection, we perform our main cross-sectional analysis using two alternative measures of implied COE to construct excess COC. The first implied COE measure, CT COE, is estimated following the approach of Claus and Thomas (2001) (hereafter, CT). Similar to the GLS COE measure, the CT COE measure is an internal rate of return from the residual income valuation model. The CT model uses five years of earnings forecasts (compared to twelve years in the GLS model) and assumes that the terminal growth in residual income is equal to the
expected inflation rate (compared to zero in the GLS model). The CT expression for price per share at time \( t \) is:

\[
P_t = B_t + \sum_{i=1}^{5} \frac{FEPS_{t+i} - r_e B_{t+i-1}}{(1+r_e)^i} + \frac{FEPS_{t+5} - r_e B_{t+4}}{(r_e - g)(1+r_e)^5},
\]

where \( B_{t+i} \) is the book value per share computed using the clean-surplus assumption, \( FEPS_{t+i} \) is the \( i \)-period-ahead earnings per share forecast,\(^{26}\) \( g \) is the terminal growth rate of residual earnings equal to the expected inflation rate (nominal risk-free rate minus a real risk-free rate of 3%), and \( r_e \) is the cost of equity capital. The implied cost of equity is estimated through an iterative procedure described in detail in Section 3.1.2.

The second COE measure, PEG COE, is based on Easton’s (2004) specification of the Ohlson and Juettner-Nauroth (2005) abnormal earnings growth model. The model equates the price of one share to the sum of capitalized one-year-ahead EPS and the capitalized abnormal growth in EPS. The Ohlson and Juettner-Nauroth model does not rely on clean-surplus accounting, and therefore is simply a reformulation of the discounted dividend model. Easton makes two simplifying assumptions, zero future dividends and zero growth in abnormal earnings changes beyond two years, to arrive at the PEG model:

\[
P_t = \frac{FEPS_{t+2} - FEPS_{t+1}}{(r_e)^2},
\]

where all variables are as previously defined. From the above model, PEG COE is calculated as a function of the forward earnings-to-price ratio and the expected earnings growth rate:

\[
 r_e = \sqrt{g \frac{FEPS_{t+2}}{P_t}}, \quad \text{where} \quad g = \frac{(FEPS_{t+2} - FEPS_{t+1})}{FEPS_{t+1}}
\]

\(^{26}\) We use three-, four-, and five-year-ahead forecasts for future earnings per share when available in I/B/E/S. If any of these forecasts is unavailable, we estimate the corresponding values by applying the long-term growth rate to the two-year-ahead forecast.
The PEG COE can be estimated only for firms where two-year-ahead EPS forecasts exceed one-year-ahead EPS forecasts. In addition, we restrict the estimation to firms with forward earnings-to-price ratios greater than 0.5%. We incorporate the predicted earnings long-term growth rate ($ltg$) in the estimation by setting $g$ equal to an average of one-year ahead earnings growth rate and $ltg$. The additional winsorization procedures include restricting $ltg$ to be less than 50%, restricting the one-year ahead growth rate to fall between $ltg$ and 1, and restricting PEG COE to be less than 1.

The results of our main analysis using these two alternative measures of implied cost of equity are reported in Table 6. On the left (right) panel, excess COC is computed using CT (PEG) as a proxy for implied cost of equity. Consistent with our earlier findings, the coefficients on cash flow and investment correlations are positive and significant. Our cross-sectional results are robust to using CT or PEG as a proxy for cost of equity capital.

**Capital Structure and Cost of Capital**

As discussed earlier, because the essence of our model is the reduction in asset beta (systematic risk) that arises from coinsurance, the model’s predictions pertain to total cost of capital (i.e., both cost of equity and cost of debt). As such, we employ an empirical proxy that measures the weighted average of the cost of equity and debt capital. Because debt returns are not readily available for most firms, we follow an approach similar to Lamont and Polk (2001) and use aggregate bond yields to proxy for the cost of debt. While this approach ignores any firm-specific variation in expected debt returns, it is conceptually superior to using a pure cost of equity measure because it takes into consideration the importance of debt in a firm’s capital structure. In this subsection, we examine whether our main results are sensitive to the inclusion/exclusion of the variation in capital structure in the cost of capital measure. In
particular, we perform the main cross-sectional analysis on an excess cost of equity measure that is constructed similar to excess cost of capital. The results, reported in Table 7, show a positive and significant coefficient on both cash flow and investment correlations, suggesting that our main findings are at least partially driven by the cost of equity component. An interesting extension is to examine whether our results hold also for the cost of debt.

5.2.4. Economic Significance

To evaluate the economic significance of the coinsurance-related reduction in cost of capital we estimate the magnitude of the corresponding increase in firm value. In a simple Gordon growth model, under a zero dividend growth assumption, a 1% decrease in cost of capital approximately translates into a 1% increase in firm value. However, the relation between cost of capital and firm value is in general non-linear and it depends on other inputs in the valuation formula—expected earnings and earnings growth. To estimate the cost of capital effect on firm value, we compare the actual firm values to the as-if firm values calculated using imputed cost of capital (i.e. the cost of capital on a comparable portfolio of single-segment firms).

Specifically, we estimate the as-if market value of the firm based on the GLS valuation model (see Section 3.1.1):

$$MV_{iCOE} = D_{t-1} + \left[ B_t + \sum_{i=1}^{T} \frac{FROE_{t+i} - iCOE}{(1 + iCOE)^{t+i-1}} B_{t+i-1} + \frac{FROE_{t+i} - iCOE}{iCOE(1 + iCOE)^{t+i-1}} B_{t+i-1} \right]$$

In a supplemental analysis, we examine the relation between coinsurance and default risk, as proxied by credit ratings. In particular, we regress excess debt ratings on cash flow and investment correlations (controlling for the variables used in Kaplan and Urwitz (1979)). The results (untabulated) show a negative and significant association between excess debt ratings and the correlation measures, suggesting that higher cross-segment correlations (i.e., lower coinsurance) are associated with lower debt ratings (i.e., higher default risk). We acknowledge that debt ratings merely proxy for a firm’s total default risk (idiosyncratic plus systematic) and we therefore do not draw inferences on coinsurance and the cost of debt from this exercise.
where $D_{t-1}$ is the book value of debt for the latest fiscal year, $iCOE$ is the imputed cost of equity, and all other variables are as defined in Section 3.1.1. The “excess value” attributable to differences in cost of capital is calculated as the natural logarithm of the ratio of actual firm value ($MV$) to as-if firm value ($MV^{COC}$), where actual value is the sum of the book value of debt for the latest fiscal year and the market value of equity at the time of the cost of capital estimation. This measure of excess value captures the percentage gain or loss in market value resulting from the coinsurance effect on cost of capital.

The untabulated univariate results suggest a 5.5% (6%) average gain in total value when moving from the lowest to the highest coinsurance quintile and a 2% (1.8%) average gain in total value when moving from single-segment firms to the highest coinsurance quintile, where the degree of coinsurance is measured using cash-flow (investment) correlations. The corresponding median gains in total value are 5.6% (6.2%) and 5.2% (5.1%), respectively. Overall, these results are consistent with coinsurance effects of diversification having economically significant effects on firm value.

6. Conclusion

This paper studies the effect of coinsurance on a firm’s cost of capital. Our model shows that combining stand-alone firms with imperfectly correlated cash flows can lead to a reduction in systematic risk and hence the combined firm’s cost of capital, and that this coinsurance effect is decreasing in the cross-segment correlation of cash flows. Our empirical analysis provides

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28 For comparison, the average gain in total value implied by the differences in CT or PEG cost of capital is 2.3% (2.8%) for CT or 6.1% (7.8%) for PEG when moving from the lowest to the highest coinsurance quintile, and 1.5% (2.1%) for CT or 7.9% (8.9%) for PEG when moving from single-segment firms to the highest coinsurance quintile, where the degree of coinsurance is measured using cash-flow (investment) correlations. For analyses using the PEG cost of capital, we compare the as-if firm values to firm values calculated based on actual implied cost of capital estimates. We do not use actual firm values because they are not always equal to the latter due to the winsorization routines described in Section 5.2.3.
evidence consistent with the model’s predictions. In particular, we find that on average diversified firms have lower cost of capital than portfolios of comparable single-segment firms. We also find a significant and positive association between excess cost of capital and cross-segment cash flow correlations – these findings imply a 6% value gain when moving from the lowest to the highest cash flow correlation quintile. Overall, our results are consistent with coinsurance affecting a firm’s systematic risk and hence its cost of capital.

A novel contribution of this paper is that it establishes a link between coinsurance and systematic risk, and hence between coinsurance and cost of capital. Our empirical results represent a major challenge to the conventional view that corporate diversification offers no cost of capital benefit over what investors can achieve through portfolio diversification. Further, the evidence of a coinsurance effect on firms’ cost of capital has implications for both valuation and capital budgeting, and suggests that ignoring such an effect may yield incorrect firm value and NPV estimates. Equally fundamental are the implications for mergers and acquisitions. We see these questions as exciting avenues for future research.
REFERENCES


Liu, Jing, and Wei Su, 2005, Forecasting analysts’ forecast errors, Working paper, UCLA.


FIGURE 1
Timeline of Variable Measurement for a Year t Observation (Assuming December Fiscal Year-End)
This table reports summary statistics for excess cost of capital. The statistics are computed over the period 1988 to 2006 for a sample of single- and multi-segment firms. Excess cost of capital is defined as the natural logarithm of the ratio of a firm’s cost of capital to its imputed cost of capital calculated using a portfolio of comparable stand-alone firms. A firm’s cost of capital is measured as the weighted average of the implied cost of equity based on the approach of Gebhardt, Lee, and Swaminathan (2001) and the yields from the Barclays Capital Aggregate Bond Index. *** indicates statistical significance at the 1% level.

<table>
<thead>
<tr>
<th></th>
<th>Obs.</th>
<th>Mean</th>
<th>Std. Dev.</th>
<th>Lower Quartile</th>
<th>Median</th>
<th>Upper Quartile</th>
</tr>
</thead>
<tbody>
<tr>
<td>Single-Segment (SS)</td>
<td>20,046</td>
<td>-0.030***</td>
<td>0.219</td>
<td>-0.125</td>
<td>-0.001***</td>
<td>0.093</td>
</tr>
<tr>
<td>Multi-Segment (MS)</td>
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<td>0.225</td>
<td>-0.150</td>
<td>-0.025***</td>
<td>0.093</td>
</tr>
<tr>
<td>MS-SS</td>
<td></td>
<td>-0.010***</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

TABLE 1
Summary Statistics: Excess Cost of Capital
TABLE 2
Univariate Analysis on Excess Cost of Capital and Cross-segment Correlations

This table presents univariate test results on excess cost of capital. The sample period spans from 1988 to 2006. Excess cost of capital is defined in Table 1. Multi-segment firms are sorted into quintiles based on their cash flow and investment correlations. Cash flow and investment correlations measure cross-segment correlation of average industry cash flow and investment based on single-segment firms. *** indicates statistical significance at the 1% level.

| Firms Sorted by | Cash-Flow Correlations | | | Investment Correlations | | |
|-----------------------------------|-------------------------|------------|-------------------------|------------|-------------------------|
|                                   | Obs. | Sort Variable | Excess COC | Obs. | Sort Variable | Excess COC |
| Multi-Segment Firms               |      |               |            |      |               |            |
| Q1 (Lowest Correlation)           | 1,822 | 0.396 | -0.059 | 1,822 | 0.372 | -0.057 |
| Q2                                | 1,821 | 0.710 | -0.044 | 1,821 | 0.699 | -0.044 |
| Q3                                | 1,822 | 0.928 | -0.038 | 1,822 | 0.925 | -0.041 |
| Q4                                | 1,821 | 0.999 | -0.029 | 1,821 | 0.999 | -0.033 |
| Q5 (Highest Correlation)          | 1,821 | 1.000 | -0.028 | 1,821 | 1.000 | -0.023 |
| Single-Segment Firms              | 20,046 | 1.000 | -0.030 | 20,046 | 1.000 | -0.030 |
| "Q5" - "Q1"                       |      |               |            | 0.032 | *** |            |
| "Single-Segment" - "Q1"           |      |               |            | 0.029 | *** |            |


TABLE 3
Multivariate Regressions of Excess Cost of Capital on Measures of Coinsurance

This table presents regressions of excess cost of capital on measures of coinsurance. The regressions are estimated over the period 1988 to 2006 for a sample of single- and multi-segment firms. Excess cost of capital is defined in Table 1. Cash flow and investment correlations are defined in Table 2. The control variables are defined in Section 4. Robust standard errors (heteroskedasticity consistent and double clustered by firm and year) are reported in brackets. ***, **, or * indicates that the coefficient estimate is significant at the 1%, 5%, or 10% level (respectively).

<table>
<thead>
<tr>
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<th>Model 1</th>
<th>Model 2</th>
<th>Model 3</th>
<th>Model 4</th>
<th>Model 5</th>
<th>Model 6</th>
</tr>
</thead>
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<td>Cash flow correlations</td>
<td>0.057***</td>
<td>0.055***</td>
<td></td>
<td></td>
<td></td>
<td>0.052***</td>
</tr>
<tr>
<td></td>
<td>[0.014]</td>
<td>[0.015]</td>
<td></td>
<td></td>
<td></td>
<td>[0.014]</td>
</tr>
<tr>
<td>Investment correlations</td>
<td>0.056***</td>
<td>-0.005*</td>
<td>-0.026***</td>
<td>-0.027***</td>
<td>-0.027***</td>
<td></td>
</tr>
<tr>
<td></td>
<td>[0.015]</td>
<td>[0.003]</td>
<td>[0.005]</td>
<td>[0.005]</td>
<td>[0.005]</td>
<td>[0.005]</td>
</tr>
<tr>
<td>Number of segments</td>
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<td>-0.005*</td>
<td>-0.026***</td>
<td>-0.027***</td>
<td>-0.027***</td>
<td></td>
</tr>
<tr>
<td></td>
<td></td>
<td>[0.003]</td>
<td>[0.005]</td>
<td>[0.005]</td>
<td>[0.005]</td>
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</tr>
<tr>
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<td></td>
<td></td>
<td></td>
<td></td>
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</tr>
<tr>
<td></td>
<td>-0.177***</td>
<td>-0.177***</td>
<td>-0.178***</td>
<td>-0.178***</td>
<td>-0.178***</td>
<td>-0.178***</td>
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<tr>
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<td>[0.027]</td>
<td>[0.027]</td>
<td>[0.027]</td>
</tr>
<tr>
<td>Leverage</td>
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<td>0.192***</td>
<td>0.192***</td>
<td>0.141***</td>
<td>0.139***</td>
<td>0.139***</td>
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<tr>
<td></td>
<td>[0.019]</td>
<td>[0.019]</td>
<td>[0.019]</td>
<td>[0.019]</td>
<td>[0.019]</td>
<td>[0.019]</td>
</tr>
<tr>
<td>Book-to-market</td>
<td>0.004</td>
<td>0.004</td>
<td>0.004</td>
<td>0.009***</td>
<td>0.009***</td>
<td>0.009***</td>
</tr>
<tr>
<td></td>
<td>[0.003]</td>
<td>[0.003]</td>
<td>[0.003]</td>
<td>[0.003]</td>
<td>[0.003]</td>
<td>[0.003]</td>
</tr>
<tr>
<td>Logarithm of forecast dispersion</td>
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<td></td>
<td></td>
<td></td>
<td></td>
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<tr>
<td></td>
<td>-0.175*</td>
<td>-0.176*</td>
<td>-0.174*</td>
<td>-0.272***</td>
<td>-0.273***</td>
<td>-0.273***</td>
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<tr>
<td></td>
<td>[0.105]</td>
<td>[0.104]</td>
<td>[0.102]</td>
<td>[0.103]</td>
<td>[0.100]</td>
<td>[0.099]</td>
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<tr>
<td>Long-term growth forecast</td>
<td>-0.091***</td>
<td>-0.091***</td>
<td>-0.091***</td>
<td>-0.090***</td>
<td>-0.090***</td>
<td>-0.089***</td>
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<tr>
<td></td>
<td>[0.009]</td>
<td>[0.009]</td>
<td>[0.009]</td>
<td>[0.007]</td>
<td>[0.007]</td>
<td>[0.007]</td>
</tr>
<tr>
<td>Lagged 12-month return</td>
<td>-0.092***</td>
<td>-0.091**</td>
<td>-0.031</td>
<td>0.185***</td>
<td>0.130***</td>
<td>0.132**</td>
</tr>
<tr>
<td></td>
<td>[0.035]</td>
<td>[0.036]</td>
<td>[0.026]</td>
<td>[0.058]</td>
<td>[0.057]</td>
<td>[0.060]</td>
</tr>
<tr>
<td>Constant</td>
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<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>-0.092***</td>
<td>-0.091**</td>
<td>-0.031</td>
<td>0.185***</td>
<td>0.130***</td>
<td>0.132**</td>
</tr>
<tr>
<td></td>
<td>[0.035]</td>
<td>[0.036]</td>
<td>[0.026]</td>
<td>[0.058]</td>
<td>[0.057]</td>
<td>[0.060]</td>
</tr>
</tbody>
</table>

Observations | 29,153 | 29,153 | 29,153 | 29,153 | 29,153 | 29,153 |
R-squared      | 0.123  | 0.123  | 0.122  | 0.144  | 0.145  | 0.145  |
This table presents regressions of excess cost of capital on cross-segment correlations for a subsample of multi-segment firms. The regressions are estimated over the period 1988 to 2006. Excess cost of capital is defined in Table 1. Cash flow and investment correlations are defined in Table 2. The control variables are defined in Section 4. Robust standard errors (heteroskedasticity consistent and double clustered by firm and year) are reported in brackets. ***, **, or * indicates that the coefficient estimate is significant at the 1%, 5%, or 10% level (respectively).

<table>
<thead>
<tr>
<th>Regression Variable</th>
<th>Coefficient</th>
<th>Standard Error</th>
</tr>
</thead>
<tbody>
<tr>
<td>Cash flow correlations</td>
<td>0.043***</td>
<td>[0.015]</td>
</tr>
<tr>
<td>Investment correlations</td>
<td>0.041***</td>
<td>[0.014]</td>
</tr>
<tr>
<td>Number of segments</td>
<td>0.012***</td>
<td>[0.003]</td>
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<tr>
<td>Logarithm of market capitalization</td>
<td>-0.028***</td>
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<td>Leverage</td>
<td>-0.234***</td>
<td>[0.042]</td>
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<td>Book-to-market</td>
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<td>Logarithm of forecast dispersion</td>
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<td>Long-term growth forecast</td>
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<td>[0.101]</td>
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<tr>
<td>Lagged 12-month return</td>
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<tr>
<td>Constant</td>
<td>0.097</td>
<td>[0.069]</td>
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</table>

| Observations                        | 9,107       | 9,107          |
| R-squared                           | 0.134       | 0.134          |
This table presents regressions of excess cost of capital on cross-segment correlations, controlling for expected and unexpected analyst forecast errors. The regressions are estimated over the period 1988 to 2006 for a sample of single- and multi-segment firms. Excess cost of capital is defined in Table 1. Cash flow and investment correlations are defined in Table 2. The construction of expected and unexpected analyst forecast errors follows Liu and Su (2005) and Ogneva, Subramanyam, and Raghunandan (2007). The rest of the control variables are defined in Section 4. Robust standard errors (heteroskedasticity consistent and double clustered by firm and year) are reported in brackets. ***, **, or * indicates that the coefficient estimate is significant at the 1%, 5%, or 10% level (respectively).

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<th>Cash flow correlations</th>
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<tr>
<td>Investment correlations</td>
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<td>Logarithm of market capitalization</td>
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<tr>
<td>Leverage</td>
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<tr>
<td>Book-to-market</td>
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<tr>
<td>Logarithm of forecast dispersion</td>
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<td>Long-term growth forecast</td>
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<tr>
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<td>Unexpected analyst forecast error in year +1</td>
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<td>-0.210**</td>
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<table>
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<th>Observations</th>
<th>23,270</th>
<th>23,270</th>
</tr>
</thead>
<tbody>
<tr>
<td>R-squared</td>
<td>0.189</td>
<td>0.189</td>
</tr>
</tbody>
</table>
This table presents regressions of excess cost of capital on cross-segment correlations using two alternative approaches, CT and PEG, instead of GLS to derive the implied cost of equity. The regressions are estimated over the period 1988 to 2006 for a sample of single- and multi-segment firms. Excess cost of capital is defined in Table 1, and CT and PEG implied cost of equity are computed based on the approach of Claus and Thomas (2001) and Easton (2004), respectively. Cash flow and investment correlations are defined in Table 2. The control variables are defined in Section 4. Robust standard errors (heteroskedasticity consistent and double clustered by firm and year) are reported in brackets. ***, **, or * indicates that the coefficient estimate is significant at the 1%, 5%, or 10% level (respectively).

<table>
<thead>
<tr>
<th></th>
<th>CT</th>
<th>PEG</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>Cash flow correlations</strong></td>
<td>0.026***</td>
<td>0.051***</td>
</tr>
<tr>
<td></td>
<td>[0.010]</td>
<td>[0.011]</td>
</tr>
<tr>
<td><strong>Investment correlations</strong></td>
<td>0.029***</td>
<td>0.056***</td>
</tr>
<tr>
<td></td>
<td>[0.009]</td>
<td>[0.012]</td>
</tr>
<tr>
<td><strong>Number of segments</strong></td>
<td>0.010***</td>
<td>0.002</td>
</tr>
<tr>
<td></td>
<td>[0.002]</td>
<td>[0.002]</td>
</tr>
<tr>
<td><strong>Logarithm of market capitalization</strong></td>
<td>-0.025***</td>
<td>-0.029***</td>
</tr>
<tr>
<td></td>
<td>[0.004]</td>
<td>[0.003]</td>
</tr>
<tr>
<td><strong>Leverage</strong></td>
<td>-0.109***</td>
<td>-0.124***</td>
</tr>
<tr>
<td></td>
<td>[0.013]</td>
<td>[0.021]</td>
</tr>
<tr>
<td><strong>Book-to-market</strong></td>
<td>-0.083***</td>
<td>0.010</td>
</tr>
<tr>
<td></td>
<td>[0.010]</td>
<td>[0.006]</td>
</tr>
<tr>
<td><strong>Logarithm of forecast dispersion</strong></td>
<td>0.018***</td>
<td>0.030***</td>
</tr>
<tr>
<td></td>
<td>[0.002]</td>
<td>[0.002]</td>
</tr>
<tr>
<td><strong>Long-term growth forecast</strong></td>
<td>0.187***</td>
<td>0.517***</td>
</tr>
<tr>
<td></td>
<td>[0.041]</td>
<td>[0.046]</td>
</tr>
<tr>
<td><strong>Lagged 12-month return</strong></td>
<td>-0.060***</td>
<td>-0.072***</td>
</tr>
<tr>
<td></td>
<td>[0.008]</td>
<td>[0.007]</td>
</tr>
<tr>
<td><strong>Constant</strong></td>
<td>0.203***</td>
<td>0.153***</td>
</tr>
<tr>
<td></td>
<td>[0.033]</td>
<td>[0.035]</td>
</tr>
</tbody>
</table>

| Observations | 26,280 | 26,280 | 27,302 | 27,302 |
| R-squared    | 0.194  | 0.195  | 0.088  | 0.088  |
TABLE 7
Multivariate Regressions of Excess Cost of Equity Capital on Cross-Segment Correlations

This table presents regressions of excess cost of equity capital on cross-segment correlations. The regressions are estimated over the period 1988 to 2006 for a sample of single- and multi-segment firms. Excess cost of equity is defined as the natural logarithm of the ratio of a firm’s cost of equity based on the approach of Gebhardt, Lee, and Swaminathan (2001) to its imputed cost of equity calculated using a portfolio of comparable stand-alone firms. Cash flow and investment correlations are defined in Table 2. The control variables are defined in Section 4. Robust standard errors (heteroskedasticity consistent and double clustered by firm and year) are reported in brackets. ***, **, or * indicates that the coefficient estimate is significant at the 1%, 5%, or 10% level (respectively).

<table>
<thead>
<tr>
<th>Variable</th>
<th>Coefficient 1</th>
<th>Coefficient 2</th>
</tr>
</thead>
<tbody>
<tr>
<td>Cash flow correlations</td>
<td>0.084***</td>
<td></td>
</tr>
<tr>
<td>Investment correlations</td>
<td>0.074***</td>
<td>0.010***</td>
</tr>
<tr>
<td>Number of segments</td>
<td>0.011***</td>
<td>0.010***</td>
</tr>
<tr>
<td>Logarithm of market capitalization</td>
<td>-0.029***</td>
<td>-0.029***</td>
</tr>
<tr>
<td>Leverage</td>
<td>0.016</td>
<td>0.016</td>
</tr>
<tr>
<td>Book-to-market</td>
<td>0.193***</td>
<td>0.193***</td>
</tr>
<tr>
<td>Logarithm of forecast dispersion</td>
<td>0.003</td>
<td>0.002</td>
</tr>
<tr>
<td>Long-term growth forecast</td>
<td>-0.142</td>
<td>-0.141</td>
</tr>
<tr>
<td>Lagged 12-month return</td>
<td>-0.097***</td>
<td>-0.097***</td>
</tr>
<tr>
<td>Constant</td>
<td>-0.001</td>
<td>0.008</td>
</tr>
</tbody>
</table>

| Observations                            | 29,150        | 29,150        |
| R-squared                               | 0.158         | 0.158         |