Corporate Diversification and the Cost of Capital

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First Draft: November 2008
This Draft: January 2009

Acknowledgments: We thank Paul Fischer, Jerry Hoberg, Chris Jones, John Matsusaka, and workshop participants at Penn State and USC for their helpful comments. We thank Yifeng Zhou for his excellent research assistance.

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ABSTRACT

This paper examines whether coinsurance arising from corporate diversification affects a firm's cost of capital. We develop a model in which diversification reduces systematic risk and leads to a lower cost of capital. This coinsurance benefit of diversification is decreasing in the cross-segment correlation of cash flows. Using measures of implied cost of capital constructed from analyst forecasts, we test and find support for the model's empirical predictions. We find that on average diversified firms have lower implied cost of capital compared to portfolios of stand-alone firms. This difference decreases as cross-segment cash flow and investment correlations increase. Overall, our results are consistent with the coinsurance effect of diversification lowering the cost of capital.

1. Introduction

Whether and how diversification affects cost of capital has important implications for valuation and capital budgeting. Yet it seems we know little to answer these fundamental questions despite a large literature in corporate finance that examines diversification and firm value. The state of our knowledge may in part reflect the conventional view that diversification offers no cost of capital benefit over what investors can achieve through portfolio diversification. The purpose of this study is to reach beyond this conventional view by examining, theoretically and empirically, the coinsurance effect of diversification on cost of capital.

The notion of coinsurance, first introduced by Lewellen (1971), posits that the imperfect correlations between business units' cash flows reduces conglomerate firms' default risk and hence increases their debt capacity. In this paper, we argue that coinsurance can also affect a firm's systematic risk. In particular, we develop a parsimonious model to demonstrate that combining stand-alone firms with imperfectly correlated cash flow can lead to a reduction in asset beta and hence the cost of capital. This coinsurance benefit arises from the merged firm's ability to transfer resources across business segments (i.e., from the segment with high cash flow to the segment with low cash flow) and thereby avoid deadweight costs.[1] Because the scope for coinsurance is greater when overall economic conditions are worse, we show that diversification tends to reduce systematic risk.

Specifically, our model yields two empirical predictions: 1) diversified firms have lower cost of capital than a comparable portfolio of stand-alone firms; and 2) this coinsurance benefit

[1] Our model relies on the standard assumption in the boundaries of the firm literature that it is difficult or costly to write complete state-contingent contracts to transfer resources between cash-rich and cash-poor firms (Gertner, Scharfstein and Stein (1994), Stein (1997), and Rajan, Servaes, and Zingales (2000)). Without this assumption, corporate diversification would offer no benefit over what investors can achieve through portfolio diversification. See Section 2 for a more detailed discussion of the model.

(i.e., the reduction in cost of capital) decreases in the cross-segment correlation of cash flows. We test our model's predictions with a large sample of single- and multi-segment firms spanning 1988 to 2006. Our cost of capital proxy measures the weighted average cost of equity and debt. In particular, we use ex ante measures of the implied cost of equity constructed from analyst forecasts and yields from the Dow Jones Corporate Bond Index to proxy for expected equity and debt returns, respectively.[2] 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 (e.g., Lee, Ng, and Swaminathan (2007), Pastor, Sinha, and Swaminathan (2008)).[3] Our empirical analyses are based on an "excess cost of capital" measure that benchmarks the cost of capital of a conglomerate firm against that of a comparable portfolio of stand-alone firms.[4]

Our results are consistent with the model's predictions. Specifically, we find that excess cost of capital is significantly negative for diversified firms, i.e., 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 in both univariate portfolio and multivariate regression analyses. Put differently, greater coinsurance (i.e., lower cross-segment cash flow correlations) is associated

[2] Our empirical proxy for expected debt returns is admittedly a relatively crude proxy – it measures only the aggregate bond yield and hence it does not capture any cross-sectional 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. We therefore follow an approach that is similar to that in Lamont and Polk, who define total cost of capital as the weighted average of a firm's realized equity returns and the returns on an aggregate bond index. This measure of expected total returns is conceptually superior to a purely expected equity return measure and yet allows us to retain a relatively large sample. To the extent that coinsurance from diversification lowers both cost of equity and cost of debt, our empirical proxy would understate the coinsurance effect on total cost of capital.

[3] We also perform sensitivity tests using alternative implied cost of equity measures based on Claus and Thomas (2001) and Easton (2004) and find that our results are robust to using these alternative measures.

[4] This measure is similar in spirit to the excess value measure (as defined in Berger and Ofek (1995)) used in the diversification literature, which captures the extent to which a diversified firm's total value is different from the sum of imputed values from its segments as stand-alone entities.

with lower cost of capital. These findings are robust to using alternative measures of implied cost of equity capital and to controlling for analyst forecast errors. Taken together, our results are consistent with the coinsurance effect of diversification reducing the cost of capital.

This paper makes several contributions to the literature. First, our study is the first to establish a link between coinsurance and firm systematic risk, and hence between coinsurance and cost of capital. Lewellen's (1971) seminal work has stimulated a stream of research that explores the coinsurance effect of diversification for corporate debt (e.g., Kim and McConnell (1977), Mansi and Reeb (2002)). The debt coinsurance argument has also been extended to incorporate the potential wealth transfer from bondholders to shareholders in conglomerate mergers (e.g., Higgins and Schall (1975), Scott (1977)). However, whether coinsurance from diversification affects cost of capital 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 emerged to study various explanations of its existence (or absence). Most theories relate to cash flow differences between conglomerates and stand-alone firms.[5] Lamont and Polk (2001) are the first to raise the possibility that the discount (or premium) may arise due to differences in expected returns. They address this question by examining the cross-sectional variation in excess values and future returns and find a significant and negative association, suggesting that the diversification discount is explained in

[5] For instance, the diversification discount can be explained by arguments related to empire building (Jensen (1986)), entrenched managers (Shleifer and Vishny (1989)), or inefficient internal capital markets (Rajan, Servaes, and Zingales (2000), Scharfstein and Stein (2000)). Recent research argues that the diversification discount is not caused by diversification per se, but rather arises from self-selection and the endogeneity of the decision to diversify (e.g., Hyland (1997), Campa and Kedia (2002), Graham, Lemmon, and Wolf (2002), Villalonga (2004)). Another stream of recent research demonstrates that the diversification discount can arise as a value-maximizing solution; see, e.g., Matsusaka (2001). Recent research also questions the view that diversified firms are less productive or allocate resources less efficiently than standalone firms, e.g. Schoar (2002), Maksimovic and Phillips (2002).

part by differences in expected returns. While their study introduces the important role of expected returns in understanding the valuation of conglomerates, their main focus is to explain the cross-sectional variation in excess value, and thus the question of *how* diversification affects a firm's cost of capital remains unanswered.[6] Our study fills this gap by providing a theoretical argument and empirical evidence on the coinsurance effect of diversification on the cost of capital.

Third, while it is not the purpose of this study to address the diversification discount debate, our results provide new insights on the debate. In particular, if there is indeed a diversification discount, then the finding of a lower cost of capital, on average, for conglomerates when compared with portfolios of stand-alone firms suggests that the cash flow differences between conglomerates and stand-alone firms are even more pronounced (since prior literature has generally assumed that there is no cost of capital difference between diversified and stand-alone firms).

Lastly, our evidence has implications for capital budgeting. In practice, managers tend to ignore the coinsurance benefit of increased 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 benefit even in the absence of taxes. 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

[6] In particular, Lamont and Polk (2001) do not attempt to explain the effect of diversification on expected returns. They conclude that the negative correlation between excess value and future returns is merely a manifestation of the well-known value effect. The objective of our study is rather different – we provide an explanation for why diversified firms have different expected returns by exploring the coinsurance effect.

findings suggest that ignoring coinsurance effects and using project-specific discount rates as commonly taught and practiced may not yield correct NPV estimates (i.e., understate NPVs).

The remainder of the paper is organized as follows. Section 2 develops the model that demonstrates the coinsurance effect on cost of capital and presents the model's empirical predictions. 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

In this section, we outline a simple model of corporate diversification to demonstrate how coinsurance benefits that arise from integration can lead to a reduction in asset beta and the total cost of capital. To illustrate the main idea, we begin with a model that assumes all-equity financing. We then incorporate debt financing into the model to show that coinsurance benefits also reduce the cost of debt.[7]

In the following subsections, we first describe the two-state economy in which we study the coinsurance effect of corporate diversification on the cost of capital. Because our main point is that corporate diversification reduces asset betas, we begin by establishing the relation between asset betas and equilibrium expected returns in our model. We then describe firm cash flows and present our main analysis regarding the coinsurance effect of corporate diversification on the cost of capital. In the third subsection, we discuss two extensions related to agency costs

[7] 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 (Lewellen (1971)).

and the modeling of coinsurance benefits. Finally, in Section 2.4 we present our model's testable predictions.

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

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 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

$$p_g C_g^i m_g + (1 - p_g) C_b^i m_b = V^i . \tag{1}$$

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* $t = 0$.

$$\frac{E[C^i]}{1 + E[r^i]} = V^i \tag{2}$$

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 $\beta^i \equiv (C_g^i / C_b^i - 1)$. As the following proposition shows, we can use β^i as an analytically convenient measure of the systematic risk of asset i's cash flow.

Proposition 1 *In equilibrium,* $E[r^i]$ *depends only on* β^i *and increases in* β^i.

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

$$1+E[r^i]=\frac{p_g C_g^i+(1-p_g)C_b^i}{p_g C_g^i m_g+(1-p_g)C_b^i m_b}.$$

Restating $E[r^i]$ in terms of β^i in an equilibrium summarized by (p_g, m_g, m_b),

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

Simple algebra shows that

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

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, $\partial E[r^i]/\partial\beta^i > 0$. *Q.E.D.*

2.2. 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.

2.2.1. All-equity Financing

We assume that firms are all-equity financed. This assumption is not necessary, but it streamlines the exposition of our main point that coinsurance benefits of integration reduce asset betas. In the next subsection, we extend the model to show that coinsurance benefits also apply to debt financing.

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 θ and $(1-\theta)$, respectively. The parameter θ depends on the state of the economy. Specifically, the probability of a high outcome is $\theta_g (\theta_b)$ 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 supplier and customer defections, in which case the firm loses L and investors receive 0.[8]

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, and the expected rate of return on a stand-alone firm (S) is determined by $\beta^S \equiv (C_g^S / C_b^S - 1)$, where $C_g^S = \theta_g H$ and $C_b^S = \theta_b H$. A stand-alone firm with

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[8] 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 may be about counterparty exposure when entering into long-term contracts. Further, employees may defect if they think waiting to find new employment until everyone is doing the same would be costly. Hence, L represents the present value of both current and future losses.

$\theta_g > \theta_b$ carries positive systematic risk whereas a stand-alone firm $\theta_g < \theta_b$ 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. A benefit of such a corporate structure is that when one of the projects experiences a low outcome, the suppliers and customers of the project do not lose confidence in the firm if the other project has a high outcome. If both projects experience a low outcome, however, then there is nothing that can be done to avoid the costly supplier and customer defections.[9]

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

$$C_g^D = \theta_g^2(2H) + 2\theta_g(1-\theta_g)(H+L)$$
$$C_b^D = \theta_b^2(2H) + 2\theta_b(1-\theta_b)(H+L).$$

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

[9] To obtain our results, the critical assumption is that a stand-alone firm incurs some loss L when the project's outcome is *l*, but a diversified firm with two projects may avoid this loss or incur a lower loss than L if the outcome of at least one of the two projects is *h*.

As the next proposition shows, one implication of coinsurance may be to reduce systematic risk in addition to increasing cash flows $(C_g^D > 2C_g^S, C_b^D > 2C_b^S)$.[10]

Proposition 2 *Combining two stand-alone firms with positive systematic and independent idiosyncratic risks reduces systematic risk and leads to a lower cost of capital.*

Proof. Given Proposition 1, it suffices to show that $\beta^S > \beta^D$. Substituting the cash flows above,

$$\beta^S = \frac{\theta_g}{\theta_b} - 1$$

$$\begin{aligned}\beta^D &= \frac{\theta_g^2(2H) + 2\theta_g(1-\theta_g)(H+L)}{\theta_b^2(2H) + 2\theta_b(1-\theta_b)(H+L)} - 1 \\ &= \frac{\underbrace{2\theta_g H}_{2C_e^S} + \underbrace{2\theta_g(1-\theta_g)L}_{\text{Coinsurance}}}{2\theta_b H + 2\theta_b(1-\theta_b)L} - 1.\end{aligned}$$

Finally, since $\theta_g > \theta_b$, $\beta^S > \beta^D$. *Q.E.D.*

An intuitive way to think about Proposition 2 is that a diversified firm offers two sets of cash flow: (i) the cash flow of two stand-alone firms, and (ii) an additional coinsurance cash flow whose beta,

$$\beta^{CI} = C_g^{CI} \big/ C_b^{CI} - 1,$$

is lower than that of stand-alone firms – β^{CI} is less than β^S – because the relative probability of avoiding costly supplier and customer defections is inversely related to the state of the economy, namely, $(1-\theta_g)$ in the good state and $(1-\theta_b)$ in the bad state. Since β^D is a weighted average

[10] Coinsurance effects arise even if we introduce integration costs to make the analysis cash-neutral. See Section 2.3.1 where we introduce agency costs of integration.

of β^S and β^{CI}, it follows that β^D must be lower than β^S as long as the probability of coinsurance is not 0.

The intuition above also indicates that the way coinsurance reduces asset beta and hence cost of capital 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 $w \in \{1,..,N\}$ is that for any two states w' and w'' with stochastic discount factor values $m(w') \le m(w'')$, $\theta(w')$ 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 increases in the state of the economy as indicated 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 $\rho \in [\underline{\rho}, 1]$ represent the correlation of idiosyncratic risks in both states of the economy $e \in \{g, b\}$. Then we have:

$$
\begin{aligned}
p_{hh,e} &= \theta_e(\theta_e + \rho(1-\theta_e)) \\
p_{lh,e} &= (1-\theta_e)(\theta_e - \rho\theta_e) \qquad (= p_{hl,e}) \\
p_{hl,e} &= \theta_e(1-\theta_e - \rho(1-\theta_e)) \qquad (= p_{lh,e}) \\
p_{ll,e} &= (1-\theta_e)(1-\theta_e + \rho\theta_e).
\end{aligned}
$$

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

$$\underline{\rho} = \max\left\{-\frac{\theta_g}{1-\theta_g}, -\frac{1-\theta_g}{\theta_g}, -\frac{\theta_b}{1-\theta_b}, -\frac{1-\theta_b}{\theta_b}\right\}.$$

In addition, joint probabilities are consistent with marginal probabilities.

$$\theta_e = p_{hh,e} + p_{hl,e} = p_{hh,e} + p_{lh,e}$$
$$1-\theta_e = p_{lh,e} + p_{ll,e} = p_{hl,e} + p_{ll,e}$$

The case in which ρ equals 0 corresponds to the case of independence in Proposition 2. When ρ equals 1 (perfect correlation),

$$p_{hh,e} = \theta_e,\ p_{lh,e} = p_{hl,e} = 0,\ p_{ll,e} = (1-\theta_e).$$

The case of perfect correlation for a diversified firm represents a doubling of scale without any coinsurance benefit.

Proposition 3 *The systematic risk and cost of capital of a diversified firm (combining two stand-alone firms with positive systematic risk) increase in* ρ*, and reach those of a stand-alone firm in the limit when* ρ *equals* 1.

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

Using the new probability structure,

$$\beta^D = \frac{\theta_g(\theta_g+\rho(1-\theta_g))(2H)+2\theta_g(1-\theta_g)(1-\rho)(H+L)}{\theta_b(\theta_b+\rho(1-\theta_b))(2H)+2\theta_b(1-\theta_b)(1-\rho)(H+L)}-1$$
$$= \frac{\underbrace{2\theta_g H}_{2C_e^S} + \underbrace{2\theta_g(1-\theta_g)(1-\rho)L}_{\text{Coinsurance}}}{2\theta_b H + 2\theta_b(1-\theta_b)(1-\rho)L}-1.$$

Simple algebra shows that

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

Also, when ρ equals 1, coinsurance cash flows drop out of β^D, and β^D equals β^S. *Q.E.D.*

Proposition 3 demonstrates how a diversified firm benefiting from a higher level of coinsurance should have a lower cost of capital compared to a portfolio of comparable stand-alone firms. Since β^D is a weighted average of β^S and β^{CI}, and the weight on $\beta^{CI}\left(<\beta^S\right)$ is directly proportional to $(1-\rho)$, β^D is always less than or equal to β^S, increases in ρ, and eventually reaches β^S when ρ equals 1. While Propositions 2 and 3 consider the case of identical stand-alone firms, the results generalize to the case in which stand-alone firms have different positive betas. Given that firms almost always have positive betas, the main message of our model covers a wide range of situations.

2.2.2. *Debt Financing*

Thus far we have excluded debt financing to streamline the exposition of our main point that coinsurance benefits reduce asset betas and hence the total cost of capital. We now incorporate debt financing into the main model to show that our results extend naturally to the cost of debt.

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, $0 < \Delta < \left(H - L\right)/2$ so that $\left(H + L\right)/2 < K < H$. Then, depending on the state of the economy $e \in \{g, b\}$, stand-alone bondholders with face value K receive

$$\begin{aligned} B_g^S &= \theta_g \left(H - \Delta\right) \\ B_b^S &= \theta_b \left(H - \Delta\right) \end{aligned},$$

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

$$B_g^D = \theta_g^2(2(H-\Delta)) + 2\theta_g(1-\theta_g)(H+L)$$
$$B_b^D = \theta_b^2(2(H-\Delta)) + 2\theta_b(1-\theta_b)(H+L).$$

Using the expected cash flows above to compute bond betas,

$$\beta_B^S = \frac{\theta_g\left(H-\Delta\right)}{\theta_b\left(H-\Delta\right)}$$
$$\beta_B^D = \frac{2\theta_g(H-\Delta) + 2\theta_g(L+\Delta-\theta_g(L+\Delta))}{\underbrace{2\theta_b(H-\Delta)}_{2B_e^S} + \underbrace{2\theta_b(L+\Delta-\theta_b(L+\Delta))}_{\text{Coinsurance}}}.$$

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, $\beta_B^D < \beta_B^S$, 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 the cost of debt as well.

2.3. Extensions

In this subsection, we discuss two extensions of the model. First, we allow for agency costs of integration (Rajan, Servaes, and Zingales (2000), Scharfstein and Stein (2000)), which are thought to offset benefits of integration such as the relaxation of financing constraints when credit is rationed (Stein (1997)), and show that the predictions of our model continue to hold under plausible conditions. Second, we allow coinsurance benefits to depend on the state of the economy.

2.3.1. Agency Costs

In our model, diversified firms have not only lower costs of capital, but also higher cash flows than portfolios of comparable 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, not the point of our paper. It therefore seems important that our model accommodate both valuation possibilities.

Suppose that diversification brings not only coinsurance benefits, but also agency costs. Indeed, agency costs underlie the main conjecture of previous work showing that diversified firms have lower valuations (Lang and Stulz (1994), Berger and Ofek (1995)) because they have lower cash flows due to agency costs and inefficient allocation of resources (Shin and Stulz (1998), and Rajan, Servaes, and Zingales (2000)). Diversified firms may face greater information and incentive problems than stand-alone firms, which in turn may lead to greater influence activities, greater spending on pet projects, and so on (Scharfstein and Stein (2000)). These costs can be seen as closing our model to prevent the counterfactual prediction that the entire economy would be owned by one big firm to maximize coinsurance benefits (Stein (1997)).

Let A_e^D denote the fraction of firm cash flow that is wasted due to agency costs of diversification depending on the state of the economy $e \in \{g, b\}$. Then, a diversified firm's cash flow net of agency costs is given by

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

Whether agency 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 agency costs do not affect systematic risk beyond reducing firm value. If $A_g^D > A_b^D$, say because bad times discipline managers, then agency costs reduce firm beta,

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

and add to the coinsurance effect.

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

2.3.2. Cost of Defections

We have assumed so far that L, the cost of supplier and customer defections, 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,

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

remains less than β^S.

For instance, if supplier and customer defections are probabilistic and these probabilities are higher during bad times than during good times, then β^{CI} is even lower than before, and we may have been conservative in modeling 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.

2.4. 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 2 and 3. In our model, a diversified firm is able to avoid costly supplier and customer defections 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 3, and is a straightforward cross-sectional extension of Prediction 1. Because the probability of coinsurance determines the expected magnitude of risk-reducing coinsurance cash flows, and because the probability 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 consider not only the correlation of cash flows generated by the different businesses comprising a diversified firm, but also the correlation of their investment opportunities to test our model's predictions. The idea is that a significant way in which coinsurance benefits lower systematic risk may be through internal capital markets that help firms avoid costly external finance and channel resources to business units with imperfectly correlated 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.[11] 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

[11] Elton (1999) provides a detailed discussion of the shortcoming of using realized returns as a proxy for expected returns.

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.[12]

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.

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.[13] In the model, the value of the firm at time t is equal to

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

[12] 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)).

[13] 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.

where P_t is the market value of equity at time t, B_t is the book value of equity at time t, NI_{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,

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

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 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

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.[14] 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 returns and the returns on an aggregate bond index. While we also utilize an 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:

$$COC_{i,t} = D_{i,t-1} Y_t^{DJ} + (1 - D_{i,t-1})\, COEC_{i,t}, \qquad \text{(EQN)}$$

[14] 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).

where Y_t^{DJ} is the aggregate bond yield from the Dow Jones Corporate Bond Index, $COEC_{i,t}$ is the implied cost of equity (GLS), and $D_{i,t-1}$ is a firm's book value of debt divided its total value (book value of debt plus market value of common equity plus book value of debt).[15]

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.[16] To the extent that coinsurance reduces the cost of debt (as our model predicts), our results understates 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 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

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[15] 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).

[16] 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.

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 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):

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

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 on 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 each year, we compute average cash flow (investment) using all single-segment firms within the industry, and then regress average industry cash flow (investment) on average market-wide cash flow (investment).[17] Residuals from these regressions represent industry-level idiosyncratic cash flows (investments).

Next, for each year, we estimate the correlation between every possible pair of industries using the time series of industry-level idiosyncratic cash flows (investments) over the preceding ten-year period on a rolling basis. We then use these correlation estimates in constructing our cash flow (investment) coinsurance measures.

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

$$\sum_{s=1}^{n}\sum_{t=1}^{n}\frac{Sales_{is(j)}}{\sum_{u=1}^{n} Sales_{iu}}\frac{Sales_{it(k)}}{\sum_{u=1}^{n} Sales_{iu}}Corr_{[y-10,y-1]}(j,k),$$

where $Sales_{is(j)}$ 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 flow or investment between industries j and k over the ten-year period before year y, estimated using single-segment firms.

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.

[17] Cash flow is defined as operating income before depreciation and investment is measured by capital expenditures.

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.[18] 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 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 39,483 firm-year observations, of which 27,368 (12,115) observations pertain to single-segment (multi-segment) firms.[19] The sample used in the cross-sectional analyses is

[18] 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.

[19] For the sensitivity analyses that are based on the alternative COC measures, the requirement of GLS COC availability is substituted with either a CT or PEG COC availability requirement.

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 revision reflected in positive stock returns, earnings forecasts will be biased downwards, leading to downwardly biased implied cost of equity estimates.[20]

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

[20] 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.

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);[21]
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.

The timeline of the variable measurement is depicted in Figure 1. Note that these additional data requirements constrain our sample to 29,276 observations, of which 20,109 (9,167) 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.

[21] All numbered items refer to the Compustat annual database.

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 is negative and significant (-0.026 and -0.011). For the single-segment subsample, median excess COC is by construction equal to zero; mean excess COC is negative and significant, suggesting that the distribution of excess COC is negatively skewed. Further, we examine the difference in means between the single- and multi-segment subsamples, and find that the difference is negative and significant (0.005 at $p<0.10$), suggesting that diversified firms on average have a lower cost of capital than stand-alone firms.

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

The summary statistics discussed above provide preliminary evidence on the coinsurance effect of diversification on the cost of capital. In this section, we directly test the coinsurance hypothesis by examining the relation between excess COC and our measures of coinsurance, cash flow correlations and investment correlations, which measure the degree of correlation between business units' cash flows and investments, respectively. Hence, a higher (lower) correlation means lower (greater) coinsurance. In the following subsections, we present our results from the univariate portfolio tests and multivariate regression tests of the coinsurance hypothesis.

5.2.1. Univariate Analysis

Table 2 reports results from our univariate analysis. In particular, 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 correlation (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).[22] Therefore, they serve as an interesting benchmark group for the coinsurance analysis.

Because the two sets of portfolio results are very similar, we focus our discussion on those from the cash flow correlation portfolios. Consistent with our coinsurance hypothesis, we observe a monotonic positive relation between cash flow correlations and excess COC across the five quintiles, with the mean difference between Q5 and Q1 positive and significant. The mean difference between the single-segment sample and Q1 is also positive and significant. Overall, these results provide preliminary support for Prediction 2 – the coinsurance benefit (i.e., a lower cost of capital) is decreasing in cross-segment cash flow and investment correlations.

[22] 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.

5.2.2. Multivariate Analysis

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

The first set of specifications, Models 1 and 2, regress excess COC on our main variables of interest, cash flow and investment correlations, respectively, along with all control variables except for the number of segments (NUMSEG) and the log of market capitalization (SIZE). The reason we exclude NUMSEG and SIZE in these specifications is as follows. Larger firms or firms with more segments are more likely to have business units with imperfect cash flow correlations. Therefore, NUMSEG and SIZE are likely to capture some degree of coinsurance, and including these two measures in the regressions would over control for the coinsurance effect from the correlation measures.

In the second set of specifications, Models 3 and 4, we use NUMSEG and SIZE respectively, as an alternative measure 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 imperfect 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 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 coefficient 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).[23]

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 benefit from diversification increases as cross-segment cash flow correlation decreases. In the following subsections, we provide several sensitivity tests to ensure that the results from our 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 enjoy any coinsurance benefit. 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

[23] 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.

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). Their model includes the following predictors that proxy for systematic biases in analyst forecasts: (1) past earnings surprises, (2) past stock returns, (3) recent analyst earnings forecast revisions, and variables related to overreaction to past information, namely, (4) book-to-market ratio, (5) forward earnings-to-price ratio, (6) long-term growth forecast, (7) past sales growth, (8) investments in property, plant, and equipment, (9) investments in other long-lived assets, and (10) the accrual component of earnings. In addition, we include a variable measuring the forecast horizon. 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 Swaminanthan (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,[24] g is the terminal growth rate of residual earnings equal to the expected inflation rate (risk-free rate minus 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

[24] 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.

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

$$r_e = \sqrt{(FEPS_{t+2} - FEPS_{t+1}) / P_t}\ .$$

The PEG COE can be estimated only for firms where two-year-ahead EPS forecasts exceed one-year-ahead EPS forecasts.

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 COE 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 on 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.[25]

6. Conclusion

This paper studies the coinsurance effect of diversification on the cost of capital. Our model shows that combining stand-alone firms with imperfectly correlated cash flow can lead to a reduction in systematic risk and hence the cost of capital. This coinsurance benefit of diversification is decreasing in the cross-segment correlation of cash flows. Our empirical analysis provides 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 document a significant and positive association between excess cost of capital and cash flow/investment correlations (our coinsurance measures), suggesting that a greater

[25] 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 lower cross-segment correlations (i.e., greater coinsurance) are associated with higher debt ratings (i.e., lower default risk). We acknowledge that debt ratings merely proxy for a firm's default risk and we therefore do not draw inferences on coinsurance and the cost of debt from this exercise.

degree of coinsurance yields a lower cost of capital. Overall, our results are consistent with coinsurance affecting a firm's systematic risk and hence reducing 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 model relies on the assumption that it is difficult or costly to write complete state-contingent contracts to transfer resources between cash-rich and cash-poor firms (Gertner, Scharfstein, and Stein (1994), Stein (1997), and Rajan, Servaes, and Zingales (2000)). Without this assumption in the model, corporate diversification would offer no benefit over what investors can achieve through portfolio diversification. Our empirical results support the theory, and expand the conventional view by drawing attention to the potential coinsurance benefits of diversification.

REFERENCES

Berger, Philip G., and Eli Ofek, 1995, Diversification's effect on firm value, *Journal of Financial Economics* 37, 39–65.

Botosan, Christine A., and Marlene A. Plumlee, 2005, Assessing alternative proxies for the expected risk premium, *The Accounting Review* 80, 21–53.

Campa, José Manuel, and Simi Kedia, 2002, Explaining the diversification discount, *Journal of Finance* 57, 1731–1762.

Claus, James J. and Jacob K. Thomas, 2001, Equity risk premium as low as three percent? Evidence from analysts' earnings forecasts for domestic and international stocks, *Journal of Finance* 56, 1629–1666.

Duchin, Ran, 2008, Cash holdings and corporate diversification, Working paper, University of Michigan.

Easton, Peter D., 2004, PE ratios, PEG ratios, and estimating the implied expected rate of return on equity capital, *The Accounting Review* 79, 73–96.

Easton, Peter D., and Steven J. Monahan, 2005, An evaluation of accounting-based measures of expected returns, *The Accounting Review* 80, 501–538.

Elton, Edwin J., 1999, Expected return, realized return, and asset pricing tests, *Journal of Finance* 54, 1199–1220.

Fama, Eugene F., and Kenneth R. French, 1992, The cross-section of expected stock returns, *Journal of Finance* 47, 427–465.

Fama, Eugene F., and Kenneth R. French, 1997, Industry costs of equity, *Journal of Financial Economics* 43, 153–193.

Gebhardt, William R., Charles M. C. Lee, and Bhaskaran Swaminathan, 2001, Towards an ex–ante cost of capital, *Journal of Accounting Research* 39, 135–176.

Gertner, Robert H., David S. Scharfstein, and Jeremy C. Stein, 1994, Internal versus external capital markets, *Quarterly Journal of Economics* 109, 1211–1230.

Graham, John R., Michael Lemmon, and Jack Wolf, 2002, Does corporate diversification destroy value? *Journal of Finance* 57, 695–720.

Guay, Wayne R., S. P. Kothari, and Susan Shu, 2005, Properties of implied cost of capital using analysts' forecasts, Working paper, MIT.

Higgins, Robert C., and Lawrence D. Schall, 1975, Corporate bankruptcy and conglomerate merger, *Journal of Finance* 30, 93–113.

Hyland, David C., 1997, Why firms diversify: An empirical examination, Unpublished doctoral dissertation, Ohio State University, Columbus, Ohio.

Jegadeesh, Narasimhan, and Sheridan Titman, 1993, Returns to buying winners and selling losers: Implications for stock market efficiency, *Journal of Finance* 48, 65–91.

Jensen, Michael C., 1986, Agency costs of free cash flow, corporate finance, and takeovers, *American Economic Review* 76, 323–329.

Kaplan, Robert S., and Gabriel Urwitz, 1979, Statistical models of bond ratings: A methodological inquiry. *The Journal of Business* 52, 231–261.

Kim, E. Han, and John J. McConnell, 1977, Corporate mergers and the co–insurance of corporate debt, *Journal of Finance* 32, 349–365.

Lamont, Owen A., and Christopher Polk, 2001, Cash flows vs. returns, *Journal of Finance* 56, 1693–1721.

Lang, Larry H.P., and René M. Stulz, 1994, Tobin's Q, corporate diversification, and firm performance, *Journal of Political Economy* 102, 1248–1280.

La Porta, Rafael, 1996, Expectations and the cross-section of stock returns, *Journal of Finance* 51, 1715–1742.

Lee, Charles M. C., David Ng, and Bhaskaran Swaminathan, 2007, Testing international asset pricing models using implied costs of capital, *Journal of Financial and Quantitative Analysis*, forthcoming.

Lewellen, Wilbur G., 1971, A pure financial rationale for the conglomerate merger, *Journal of Finance* 26, 527–537.

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

Maksimovic, Vojislav, and Gordon Phillips, 2002, Do conglomerate firms allocate resources inefficiently? *Journal of Finance* 57, 721–767.

Mansi, Sattar A., and David M. Reeb, 2002, Corporate diversification: What gets discounted? *Journal of Finance* 57, 2167–2183.

Matsusaka, John, 2001, Corporate diversification, value maximization, and organizational capabilities, *Journal of Business* 74, 409–431.

Matsusaka John G., and Vikram Nanda, 2002, Internal capital markets and corporate refocusing, *Journal of Financial Intermediation* 1, 176–211.

Ogneva, Maria, K. R. Subramanyam, and K. Raghunandan, 2007, Internal control weakness and cost of equity: Evidence from SOX Section 404 disclosures, *The Accounting Review* 82, 1255–1297.

Ohlson, James A., and Beate E. Juettner-Nauroth, 2005, Expected EPS and EPS growth as determinants of value, *Review of Accounting Studies* 10, 349–365.

Pastor, Lubos, Meenakshi Sinha, and Bhaskaran Swaminathan, 2008, Estimating the intertemporal risk–return tradeoff using the implied cost of capital, *Journal of Finance* 63, 2859–2897.

Petersen, Mitchell A., 2008, Estimating standard errors in finance panel data sets: Comparing approaches, *The Review of Financial Studies*, forthcoming.

Petkova, Ralitsa, Lu Zhang, 2004, Is value riskier than growth?, *Journal of Financial Economics* 78, 187–202.

Rajan, Raghuram, Henri Servaes, and Luigi Zingales, 2000, The cost of diversity: The diversification discount and inefficient investment, *Journal of Finance* 60, 35–80.

Scharfstein, David S., and Jeremy C. Stein, 2000, The dark side of internal capital markets: Divisional rent–seeking and inefficient investment, *Journal of Finance* 55, 2537–2564.

Schoar, Antoinette S., 2002, Effects of corporate diversification on productivity, *Journal of Finance* 57, 2379–2403.

Scott Jr., James H., 1977, On the theory of conglomerate mergers, *Journal of Finance* 32, 1235–1250.

Shin, Hyun–Han, and René M. Stulz, 1998, Are internal capital markets efficient? *Quarterly Journal of Economics* 113, 531–552.

Shleifer, Andrei, and Robert W. Vishny, 1989, Managerial entrenchment: The case of manager–specific investments, *Journal of Financial Economics* 25, 123–139.

Stein, Jeremy C., 1997, Internal capital markets and the competition for corporate resources, *Journal of Finance* 52, 111–133.

Villalonga, Belén, 2004, Does diversification cause the "diversification discount"? *Financial Management* 33, 5–23.

FIGURE 1
Timeline of Variable Measurement for a Year *t* Observation (Assuming December Fiscal Year-End)

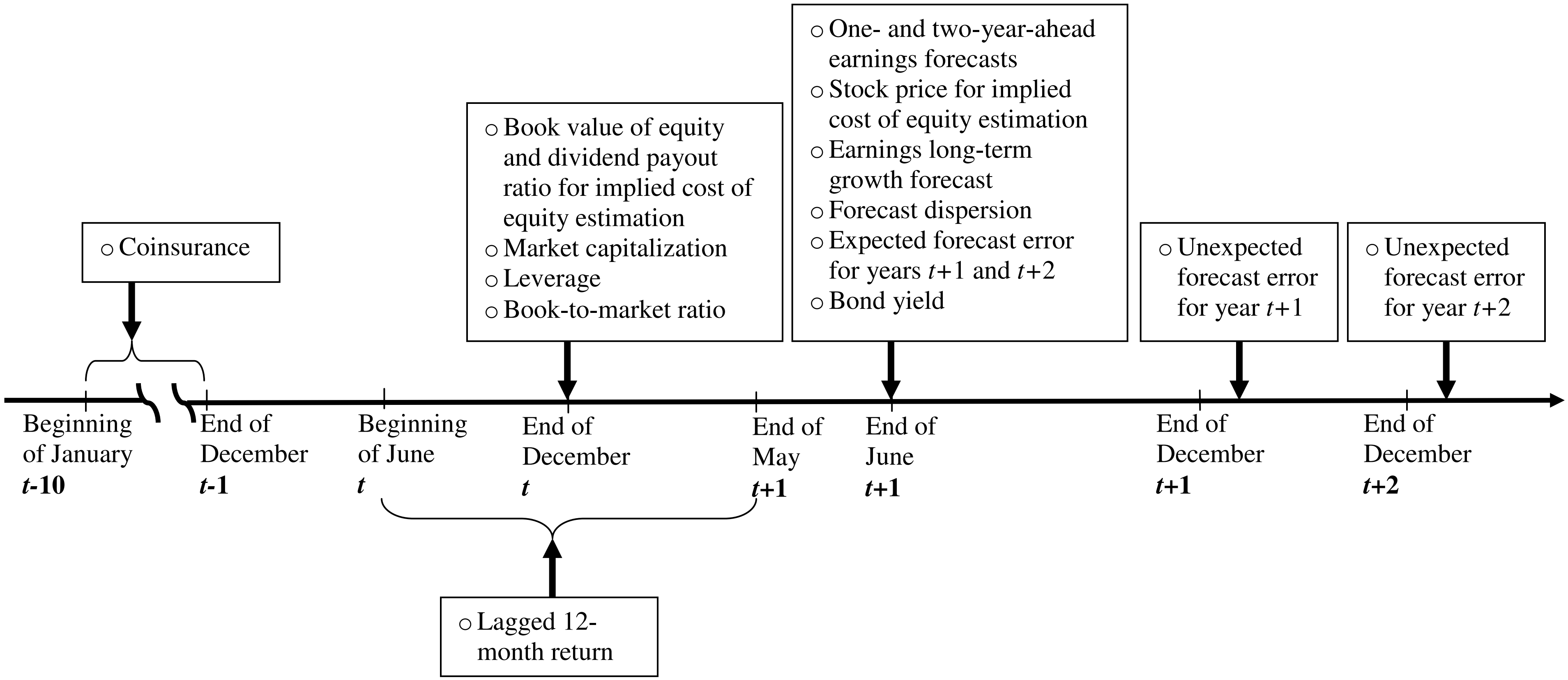

TABLE 1
Summary Statistics: Excess Cost of Capital

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 and the yields from the Dow Jones Corporate Bond Index.

	Obs.	Mean	Std. Dev.	Lower Quartile	Median	Upper Quartile
Single-Segment (SS)	27,368	-0.021***	0.266	-0.118	0.000	0.110
Multi-Segment (MS)	12,115	-0.026***	0.266	-0.139	-0.011***	0.115
MS-SS		-0.005*			-0.011***	

TABLE 2

Univariate Analysis on Excess Cost of Capital and Cross-segment Correlations

This table presents univariate portfolio test results on excess cost of capital. The sample period spans from 1988 to 2006. Multi-segment firms are sorted into quintiles based on their cash flow and investment correlations. 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 and the yields from the Dow Jones Corporate Bond Index. Cash flow and investment correlations (across segments) measure the idiosyncratic correlation of average industry cash flow and investment, respectively, between the industries that a firm's business segments operate in over the last ten years, estimated using single-segment firms.

	Firms Sorted by Cash Flow Correlations				Investment Correlations			
	Obs.	Sort Variable	Excess COC		Obs.	Sort Variable	Excess COC	
Multi-Segment Firms								
Q1 (Lowest Correlation)	1,834	0.396	-0.056		1,834	0.371	-0.052	
Q2	1,833	0.709	-0.040		1,833	0.698	-0.042	
Q3	1,834	0.928	-0.038		1,834	0.925	-0.040	
Q4	1,833	1.000	-0.026		1,833	1.000	-0.025	
Q5 (Highest Correlation)	1,833	1.000	-0.025		1,833	1.000	-0.026	
Single-Segment Firms	20,109	1.000	-0.030		20,109	1.000	-0.030	
Q5 - Q1			0.031	***			0.026	***
Single-Segment - Q1			0.026	***			0.022	***

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 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 and the yields from the Dow Jones Corporate Bond Index. Cash flow and investment correlations (across segments) measure the idiosyncratic correlation of average industry cash flow and investment, respectively, between the industries that a firm's business segments operate in over the last ten years, estimated using single-segment firms. 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).

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
Cash flow correlations	0.056***				0.056***	
	[0.014]				[0.014]	
Investment correlations		0.054***				0.053***
		[0.015]				[0.014]
Number of segments			-0.005*		0.008**	0.008**
			[0.003]		[0.003]	[0.003]
Logarithm of market capitalization				-0.026***	-0.027***	-0.027***
				[0.005]	[0.005]	[0.005]
Leverage	-0.134***	-0.134***	-0.135***	-0.136***	-0.137***	-0.137***
	[0.025]	[0.025]	[0.026]	[0.025]	[0.025]	[0.025]
Book-to-market	0.190***	0.190***	0.191***	0.139***	0.136***	0.137***
	[0.018]	[0.018]	[0.018]	[0.019]	[0.019]	[0.019]
Logarithm of forecast dispersion	0.004	0.004	0.004	0.009***	0.009***	0.009***
	[0.003]	[0.003]	[0.003]	[0.003]	[0.003]	[0.003]
Long-term growth forecast	-0.158	-0.159	-0.156	-0.258**	-0.258**	-0.258***
	[0.103]	[0.102]	[0.100]	[0.103]	[0.100]	[0.100]
Lagged 12-month return	-0.091***	-0.091***	-0.091***	-0.090***	-0.090***	-0.090***
	[0.009]	[0.009]	[0.009]	[0.007]	[0.007]	[0.007]
Constant	-0.099***	-0.097***	-0.040	0.178***	0.122**	0.125**
	[0.034]	[0.036]	[0.025]	[0.060]	[0.058]	[0.062]
Observations	29,276	29,276	29,276	29,276	29,276	29,276
R-squared	0.118	0.118	0.117	0.138	0.139	0.139

TABLE 4
Multivariate Regressions of Excess Cost of Capital on Cross-Segment Correlations: Multi-Segment Sample

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 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 and the yields from the Dow Jones Corporate Bond Index. Cash flow and investment correlations (across segments) measure the idiosyncratic correlation of average industry cash flow and investment, respectively, between the industries that a firm's business segments operate in over the last ten years, estimated using single-segment firms. 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).

Cash flow correlations	0.044***	
	[0.015]	
Investment correlations		0.041***
		[0.014]
Number of segments	0.013***	0.013***
	[0.003]	[0.003]
Logarithm of market capitalization	-0.029***	-0.029***
	[0.007]	[0.007]
Leverage	-0.179***	-0.179***
	[0.038]	[0.038]
Book-to-market	0.165***	0.165***
	[0.029]	[0.029]
Logarithm of forecast dispersion	0.006	0.005
	[0.004]	[0.004]
Long-term growth forecast	-0.185*	-0.186*
	[0.103]	[0.101]
Lagged 12-month return	-0.082***	-0.082***
	[0.010]	[0.010]
Constant	0.094	0.095
	[0.071]	[0.074]
Observations	9,167	9,167
R-squared	0.124	0.124

TABLE 5

Multivariate Regressions of Excess Cost of Capital on Cross-Segment Correlations: Controlling for Analyst Forecast Errors

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 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 and the yields from the Dow Jones Corporate Bond Index. Cash flow and investment correlations (across segments) measure the idiosyncratic correlation of average industry cash flow and investment, respectively, between the industries that a firm's business segments operate in over the last ten years, estimated using single-segment firms. The construction and definitions of expected and unexpected forecast errors are given in Section 5.2.3. 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).

Cash flow correlations	0.051***	
	[0.015]	
Investment correlations		0.046***
		[0.014]
Number of segments	0.008**	0.008**
	[0.004]	[0.004]
Logarithm of market capitalization	-0.025***	-0.025***
	[0.005]	[0.005]
Leverage	-0.137***	-0.138***
	[0.028]	[0.028]
Book-to-market	0.146***	0.146***
	[0.018]	[0.018]
Logarithm of forecast dispersion	0.003	0.003
	[0.003]	[0.003]
Long-term growth forecast	-0.311***	-0.311***
	[0.098]	[0.098]
Lagged 12-month return	-0.004	-0.004
	[0.010]	[0.010]
Unexpected analyst forecast error in year +1	-0.226***	-0.226***
	[0.067]	[0.067]
Unexpected analyst forecast error in year +2	-0.221***	-0.221***
	[0.036]	[0.036]
Average predicted analyst forecast error in years +1 and +2	-5.026***	-5.030***
	[0.692]	[0.699]
Constant	0.087*	0.091*
	[0.047]	[0.052]
Observations	23,522	23,522
R-squared	0.192	0.191

TABLE 6

Multivariate Regressions of Excess Cost of Capital on Cross-Segment Correlations: Alternative Measures of Cost of Capital

This table presents regressions of excess cost of capital on cross-segment correlations using two alternative approaches 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. The CT and PEG implied cost of equity measures are constructed following the approach of Claus and Thomas (2001) and Easton (2004), respectively. 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 and the yields from the Dow Jones Corporate Bond Index. Cash flow and investment correlations (across segments) measure the idiosyncratic correlation of average industry cash flow and investment, respectively, between the industries that a firm's business segments operate in over the last ten years, estimated using single-segment firms. 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).

	CT		PEG	
Cash flow correlations	0.027***		0.036*	
	[0.009]		[0.019]	
Investment correlations		0.026***		0.056***
		[0.009]		[0.017]
Number of segments	0.011***	0.011***	-0.006*	-0.005
	[0.002]	[0.002]	[0.004]	[0.004]
Logarithm of market capitalization	-0.024***	-0.024***	-0.062***	-0.062***
	[0.004]	[0.004]	[0.005]	[0.005]
Leverage	-0.085***	-0.085***	-0.113***	-0.112***
	[0.017]	[0.017]	[0.028]	[0.028]
Book-to-market	-0.077***	-0.077***	0.036***	0.035***
	[0.008]	[0.008]	[0.010]	[0.010]
Logarithm of forecast dispersion	0.018***	0.017***	0.065***	0.066***
	[0.002]	[0.002]	[0.005]	[0.005]
Long-term growth forecast	0.197***	0.196***	0.425***	0.423***
	[0.042]	[0.042]	[0.078]	[0.078]
Lagged 12-month return	-0.061***	-0.061***	-0.133***	-0.133***
	[0.008]	[0.008]	[0.012]	[0.012]
Constant	0.189***	0.189***	0.488***	0.467***
	[0.033]	[0.036]	[0.070]	[0.069]
Observations	27,640	27,640	28,744	28,744
R-squared	0.083	0.083	0.210	0.211

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 implied cost of equity to its imputed implied cost of equity calculated using a portfolio of comparable stand-alone firms. Cash flow and investment correlations (across segments) measure the idiosyncratic correlation of average industry cash flow and investment, respectively, between the industries that a firm's business segments operate in over the last ten years, estimated using single-segment firms. 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).

Cash flow correlations	0.087***	
	[0.017]	
Investment correlations		0.076***
		[0.016]
Number of segments	0.011***	0.010***
	[0.004]	[0.004]
Logarithm of market capitalization	-0.030***	-0.030***
	[0.006]	[0.006]
Leverage	0.037*	0.037*
	[0.022]	[0.022]
Book-to-market	0.180***	0.180***
	[0.022]	[0.022]
Logarithm of forecast dispersion	0.002	0.002
	[0.004]	[0.004]
Long-term growth forecast	-0.169	-0.168
	[0.117]	[0.116]
Lagged 12-month return	-0.099***	-0.099***
	[0.008]	[0.008]
Constant	0.010	0.021
	[0.067]	[0.072]
Observations	29,265	29,265
R-squared	0.144	0.143